# A VAR Analysis for the Uncovered Interest Parity and the Ex-Ante Purchasing Power Parity:

## The Role of Macroeconomic and Financial Information

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**Abstract**

This study revisits the relation between the uncovered interest parity (UIP), the *ex ante* purchasing power parity (EXPPP) and the real interest parity (RIP) using a VAR approach for the US dollar, the British sterling and the Japanese yen interest rates, exchange rates and changes in prices. The original contribution is on developing some joint coefficient-based tests for the three parity conditions at a long horizon. Particularly, test results are derived from the implied slope coefficients obtained by rewriting the UIP, the EXPPP and the RIP as a set of cross-equation restrictions in the VAR (see also Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007). Consistent with the idea of some form of proportionality among the three parity conditions, we find a "forward premium" bias in both the UIP - as it is normally found in empirical analysis (e.g. Diez de los Rios and Sentana, 2007) - and the *expectational* PPP. The latter result is new in the literature and stands on having uncertainty both on the future exchange rate and price dynamics. The overall results confirm the UIP to be currency-based and the EXPPP to be horizon-dependent. Moreover, we find (weak) evidence that conditioning the VAR on variables having a strong forward-looking component (i.e. share prices) help recover a unitary coefficient in the UIP equation.

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# 1 Introduction

The interest for the uncovered interest parity (UIP) and the purchasing power parity (PPP) represents a key element in the analysis of the economic and financial arbitrage conditions on international markets.

According to the definition of PPP, the latter is defined as the exchange rate between two currencies that would equate national and foreign prices when expressed in a common currency. For PPP to hold, no arbitrage opportunities across market locations exist. A general result of the studies on PPP is that this condition does not seem to hold during floating exchange rate periods but it has performed better in other historical periods, as the prefloat international standard phase (Cheung and Lai, 1993). At that time, the faith in PPP essentially derived from the prevailing theory according to which price movements were dominated by monetary factors, and gathered the constancy of the nominal exchange rate. Indeed, under the hypothesis of long-run neutrality of money, the PPP was not susceptible of measurement errors and/or goods markets inefficiencies (see Froot and Rogoff, 1994; Sarno and Taylor, 2001). When the Bretton Woods period came to an end, the exceptional volatility of the floating exchange period could no longer be explained by standard theories, so that the collapse of PPP started soon to be imputed to the low power of testing - with all evidence reporting against the existence of PPP, at least at short horizons[1] - or to the existence of unidirectional goods markets imperfections (i.e. price stickiness, role of tradables vs. non-tradables goods, non linearities).[2]

The empirical support in favour of the UIP is on the contrary very mixed (Bekaert *et al.*, 2007; Meredith and Chinn, 2004; Diez de los Rios and Santena, 2007; Evans, 1998). The UIP predicts high yield currencies to be expected to depreciate in order to offset international capital markets arbitrage opportunities. Tests results have mostly pointed out a rejection of the UIP over the recent floating period at both high and low frequencies, as documented by the "forward premium" puzzle (a negative regression coefficient); with measurement errors (a stationary time-dependent risk premium) or violations of the rational expectations assumption (see Section 2.2) being usually the explanation provided for the finding.[3] If, on the one side, the evidence in favour of the UIP at long horizons is

[1] As soon as the floating period occurred, the collapse of PPP started to be explained by the overshooting exchange rate model proposed by Dornbusch (1976). On the empirical ground, some valid statistical results were achieved when the PPP started to be tested as a long run equilibrium condition. Some contributions such as Edison (1987), Lothian and Taylor (1996) and Taylor (2002) found the PPP to empirically hold in the long run (for one century data or more) with an half-life of about 4 years for the major industrialized countries. Such results were however not exempted from severe critiques, as long samples were found to be very inappropriate because of differences in the RER behavior not only across different historical periods but mostly across different nominal exchange rate regimes (Taylor, Peel and Sarno, 2001). For a survey see Rogoff (1996); MacDonald (1991), (1993), (1998); Taylor (2002).

[2] The relation between exchange rates and national price levels might be affected by non linearities (international transaction costs) in the real exchange rate adjustments (Taylor *et al.*, 2001; Cheung and Lai, 1993). Equivalently sticky prices in local currency may lead to PPP deviations (Engle and Rogers, 1996).

[3] One of the most striking feature of the exchange rate behaviour in UIP testing is the presence of a "forward premium" puzzle, predicting high interest rate currencies to appreciate rather then depreciate as UIP would suggest. The "carry trade" consists indeed in borrowing low-interest rate currencies and investing in high interest rate currencies, by exploiting this anomaly (see Diez de los Rios and Sentana, 2007).

recognized in the attempt of getting rid of short run exchange rate swings, on the other, the presence of speculation would suggest evidence in favour of short-run UIP. In the short run, it is very likely shocks and structural changes to drive exchange rates away from the long run equilibrium (Edison, 1987). Hence, addressing the UIP as a long run relation implies market frictions - preventing a prompt and full response of the exchange markets to interest rate changes - to completely die out. Instead, the presence of speculative activities suggests it is the long-term UIP - rather than its short-term version - to be affected by market frictions, as it is very unlikely trading desks to keep capital binded in long-term contracts (Chaboud and Wright, 2005).

Across the PPP and the UIP puzzles, more recent empirical analysis (Juselius, 1991; 1992; 1995; Johansen and Juselius, 1992; Pesaran *et al.*, 2000; Cheng, 1999; Throop, 1993; Zhou and Mahadavi, 1996; Hunter, 1991) have found evidence in favour of a PPP-UIP joint relation, emphasizing the role of government budget deficits in determining real exchange rate (RER) *disequilibria*. Short-run deviations in the RER are expected to involve real factors acting through the current account - as foreign net asset position or international imbalances - which would require a relative supply of cash flows for the balance of payment to be equilibrated back (e.g Edison, 1987).[4]

In order to test for the PPP and the UIP jointly, we introduce a third parity condition: the real interest parity (RIP). The RIP can be shown to be a combination of the *ex ante* PPP (EXPPP) and the UIP (Cumby and Obstfeld, 1980; Mishkin, 1982; Jore *et al.*, 1993), so that any couple in between these three parities naturally implies the third relation (Marston, 1997; Campbell *et al.*, 2007).

The analysis focuses on the US dollar, the British sterling and the Japanese yen interest rates, exchange rates and changes in prices. Drawing on a VAR approach, we revisit the relation between the UIP, the EXPPP and the RIP by developing some joint coefficient tests obtained from a set of VAR cross-equation restrictions (Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007) and consistent with the idea of *present value* models (Campbell and Shiller, 1987). The focus is on the long horizon in order to help recover the PPP.

In the present setting we confirm the existence of a "forward premium" puzzle for the UIP, as normally found in the literature (e.g. Diez de los Rios and Sentana, 2007). Consistent with the existence of some form of proportionality between the three parities, we find a "forward" bias also for the *expectational* version of the PPP (EXPPP); standing the latter result on having uncertainty both on future exchange rate and inflation dynamics. The whole results confirm the UIP to be currency-based while the EXPPP to be rather horizon-dependent. Finally, augmenting the original VAR framework with macroeconomic and financial variables, we find (weak) evidence that variables having a strong forward looking component (i.e. share prices) help recover a unitary coefficient in the UIP equation.

[4]Hence, the lack of short run adjustments in prices is necessarily compensated by changes in the interest rate spread.

The reminder of the paper is organized as follows. Section 2 presents the theoretical model. Section 3 introduces the econometric methodology. Section 4 presents our main results. Section 4 concludes.

# 2 Uncovered Interest Parity and the Purchasing Power Parity

## 2.1 Uncovered Interest Parity

The uncovered interest parity (UIP) follows from the definition of the (log) covered interest parity (CIP), relying itself on the assumption of arbitrage between spot and forward foreign exchange markets. Drawing on Fama (1984), a *risk-free* arbitrage condition exists if (in logs):

$$f_{t+l} - s_t = i_{t,l} - i^*_{t,l},$$

where $i_{t,l}$ represents the yield of a bond with maturity $l$ at time t in the home country, and $f_{t+l}$ is the forward value of the home vs. foreign spot nominal exchange rate, $s$, expiring $l$-periods ahead. The expression above is regardless of investors preferences (*unbiasedness* hypothesis).[5] Assuming individuals to be risk-adverse makes the forward rate to differ from the expected future spot rate, $E_t s_{t+l}$, by a premium compensating for the risk of holding assets denominated in a foreign currency (see also Fraga, 1985; Mark and Wu, 1998; Hai *et al.*, 1997). Hence,

$$f_{t+l} - E_t s_{t+l} = v_{t+l},$$

where $v_{t,t+l}$ is an *ex ante* risk premium. Substituting $v$ into the CIP gives the standard UIP,

$$E_t \Delta s_{t+l} = i_{t,l} - i^*_{t,l} - v_{t+l},$$

suggesting that the excess of home interest rate over the foreign one ($i^*$), compounded over $l$ periods, is equal to the expected depreciation of the home currency over the same period, and allowing for a risk premium. So defined, the risk premium can be positive or negative depending on whether investors would require an "excess return" to compensate for the risk of holding a particular currency. For the forward premium to be a predictor of $E_t s_{t+l}$, the UIP can be tested at the $l - th$ period horizon with the following regression (obtained by iterative substitutions)

$$\frac{1}{l}\sum_{j=1}^{l} E_t \Delta s_{t+j} = \alpha_l^{uip} + \beta_l^{uip}(i_{t,l} - i^*_{t,l}) + \epsilon_{t+l}^{uip}, \qquad (1)$$

[5] For further details see Green (1992).

under the null that $\beta^{uip} = 1$, and where the time-varying premium enters $\epsilon^{uip}_{t+l}$.

## 2.2 Purchasing Power Parity

The purchasing power parity (PPP) is defined as the exchange rate between two currencies that would equate national and foreign price levels when expressed in a common currency (Sarno and Taylor, 2002). The starting point for considering such a parity is the law of one price (LOP) in logs, asserting that for any good $i$:

$$p_t(i) = p_t^*(i) + s_t,$$

where $p_t(i)$ and $p_t^*(i)$ describe the current price for the good $i$ in the home and in the foreign economy respectively, and $s$ is the home vs. foreign nominal exchange rate. The statement underlying this law is nothing but a standard goods market arbitrage condition; net of tariffs, transportation costs and trade barriers.[6] If the LOP (at least theoretically) holds for every good $i$, the same rule is expected to hold when relying on identical baskets of goods:

$$p_t = p_t^* + s_t,$$

where $p_t$ and $p_t^*$ describe the current price levels in both the foreign and the home country. Many empirical tests do not compare however identical basket of goods, but use different countries CPIs (consumer price indices) or WPIs (wholesale price indices).[7] Constant price differentials are indeed obtained by using the so called relative consumption-based PPP (Froot and Rogoff, 1994):

$$\Delta p_t = \Delta p_t^* + \Delta s_t,$$

where $\Delta$ is the difference operator. This relation predicts the relative inflation rate across countries to be necessarily compensated by changes in the nominal exchange rate.[8] Taking expectations on both sides and reformulating it at the $l-th$ period horizon, the PPP can be expressed in *expectational* terms (EXPPP), as

$$\frac{1}{l}E_t\Delta s_{t+l} = \frac{1}{l}E_t(\Delta p_{t+l} - \Delta p^*_{t+l}).$$

[6]As a matter of fact this relation can in principle hold exclusively for highly traded goods, as gold for instance (e.g. Mussa, 1986; MacDonald and Taylor, 1992; Sarno and Taylor, 2002).

[7]The PPP has indeed no reason to hold unless the two countries share identical consumption bundles. As underlined by Froot and Rogoff (1994), in principle it might be possible to construct international price indices for identical baskets of good, though there have been "very few attempts and the literature has developed in other directions".

[8]The latter specification is more appropriate given the price inflation statistical properties (see Johansen, 1991; Juselius, 1995).

The formulation is normally augmented with a term, $o_{t+l}$, imposing a departure of the real exchange rate (RER) from the PPP equilibrium, as:[9]

$$\frac{1}{l}\sum_{j=1}^{l} E_t\Delta s_{t+j} = \alpha_l^{ppp} + \beta_l^{ppp}\left[\frac{1}{l}\sum_{j=1}^{l}\left(E_t\Delta p_{t+j} - E_t\Delta p_{t+j}^{*}\right)\right] + \epsilon_{t+l}^{ppp}, \tag{2}$$

with the RER deviations ($o_{t+l}$) being captured by $\epsilon_{t+l}^{ppp}$. If markets are *efficients*, equation (2) ensures commodity speculators to keep expected deviations from PPP in line under $\beta^{ppp} = 1$ (Roll, 1979).

### 2.3 Uncovered Interest Parity and the Purchasing Power Parity

The PPP and the UIP can be tested jointly by accounting for a third condition, the real interest parity (RIP). The RIP refers to the equality between the home and the foreign real interest rates, as:

$$r_{t+l} = r_{t+l}^{*},$$

where real rates are defined according to the Fisher's (1097) parity condition, $r_{t+l} = i_{t,l} - E_t\Delta p_{t+l}$. According to Marston (1997), the real interest parity holds as soon as capital and goods markets are in equilibrium.[10] In fact, adding and subtracting the term $E_t\Delta s_{t+l}$ in the expression above shows how the RIP becomes a relation conditional on the joint validity of the UIP and the EXPPP:

$$r_{t+l} - r_{t+l}^{*} =$$

$$= \left(i_{t,l} - i_{t,l}^{*} - E_t\Delta s_{t+l}\right) - \left(E_t\Delta p_{t+l} - E_t\Delta p_{t+l}^{*} - E_t\Delta s_{t+l}\right).$$

As it is constructed, the RIP does not allow for frictions in the behaviour of both markets and investors. Clearly, if an "excess return" and a RER deviations term exist, the RIP would necessarily allow for an erratic component $\xi_{t+l}$, which - by definition - must equal $(v_{t+l} - o_{t+l})$. In light of the above, the RIP is normally tested with the following regression, where $\epsilon_{t+l}^{rip}$ is a linear function of the UIP and the EXPPP *premia*, i.e. $\epsilon_{t+l}^{rip} = \epsilon(v_{t+l}, o_{t+l})$:

$$\left(i_{t,l} - i_{t,l}^{*}\right) = \alpha_l^{rip} + \beta_l^{rip}\left[\frac{1}{l}\sum_{j=1}^{l}\left(E_t\Delta p_{t+j} - E_t\Delta p_{t+j}^{*}\right)\right] + \epsilon_{t+l}^{rip}. \tag{3}$$

[9] The term measures the real exchange rate (RER) observed deviations. The definition of the (log) real exchange rate is indeed $rer_t = p_t - p_t^{*} - s_t$.

[10] See also MacDonald and Nagayasu (1999).

As before, the efficient markets hypothesis simply imply the joint UIP-EXPPP restriction that $\beta^{rip} = 1$.

# 3 Econometric Methodology

## 3.1 Deriving Restrictions on the VAR

An obvious problem in testing the above parity conditions is the absence of observations on market expectations of future exchange rate and inflation movements. Substituting expected values with the actual ones (see Figure 1) does not seem a convenient solution, as we induce further uncertainty given *ex post* exchange rate and/or inflation forecast errors (see Marston, 1997).[11]

In order to estimate equations (1), (2) and (3), we consider a 3-dimensional VAR of $I(0)$ variables, i.e.:

$$y_t = [\Delta s_t, (i_{t,120} - i^*_{t,120}), (\Delta p_t - \Delta p^*_t)].$$

where $i$ and $i^*_t$ are 10 yrs constant maturity Treasury bonds, $s_t$ is the bilateral nominal exchange rate (monthly average, denominated in US dollars) and $(\Delta p_t - \Delta p^*_t)$ describes the inflation spread. In this paper the US are regarded as foreign economy with all the variables expressed with a star superscript. In the present setting we consider dollar-based bilateral parities for the British sterling and the Japanese yen. Further, in the paper we consider some additional macroeconomic and financial variables as industrial production, monetary aggregates (M3), reserve assets and share prices. All data are seasonally adjusted, when needed, and taken in monthly frequencies from the OECD.stat database. Price indices (*cpi*-based), exchange rates and the macroeconomic/financial variables are transformed in month-on-month changes. The sample covers the period from 1975-1 to 2008-6. The series for the long term interest rate for Japan starts from 1989-1.

Table 1 reports some descriptive statistics for the variables. Over the overall sample, changes in the appreciation rate are found to be highly volatile but very little autocorrelated. Instead, interest rate and inflation spreads display stronger ACF up to the 4-th order. Spreads are very persistent but they do not display a *near-I(1)* problem as they are not as autocorrelated as $i$ and $\Delta p$ themselves (see Bekaert *et al.*, 2007).

We begin by determining the VAR order $K$ by means of the standard *information criteria* and select the number of lags for which at least two criteria are congruous. Namely $K = 3$ for the UK vs. US

[11] Advocating that the three parity conditions need to simultaneously hold (Marston, 1997) implies moreover some form of proportionality among the slope coefficients (discussed further in Section 3.4). Not observing this relationship empirically (whenever we depart from the null $\beta^{uip,ppp,rip} = 1$) is nonetheless not surprising, as the PPP and the UIP are fundamentally different. The PPP is a long run relation whose adjustment is expected to be backward looking, whereas the UIP is forward looking (Mishkin, 1982). The rationale for their combination stands indeed on correctly modeling market expectations.

system and $K = 1$ for the Japan vs. US system (Table 3). For each system we then reformulate $y_t$ in the standard companion form with $z_t^{'} = (y_t, y_{t-1}, ..., y_{t+1-K})$. Disregarding any constant term, the following compact form applies:

$$z_t = Az_{t-1} + e_t,$$

where the parameters matrix $A$ is a $(3K \times 3K)$ dimensional matrix with $k$ (for $k = 1, 2, ..., K$) VAR matrices stacked horizontally in the first 3 rows, a $3(K-1)$ identity matrix underneath these parameters on the left hand corner, and zero elsewhere. The innovation vector $e_t$ is assumed to have variance equal to $\Sigma$.

In this framework, testing for the parities outlined in Section 2 imposes different restrictions on the companion parameters in A. This methodology allows for multi-horizon tests, as expectations are accounted as forecasts formed from a function of past observations, i.e. $E(z_{t+j}|z_t) = A^j z_t$, and consistent with the idea of *present value* models (see Campbell and Shiller, 1987). [12]

By letting $e_n$ to be an indicator column vector that selects the n-*th* variable in the companion VAR, testing for (1),(2) and (3) results into a set of $n = 3K$ non-linear cross equation restrictions on the $3n$ coefficients of A. Using straightforward algebra, the UIP implies (Bekaert and Hodrick, 2001; Bekaert *et al.*, 2007):

$$\frac{1}{l}\sum_{j=1}^{l} e_{\Delta s}^{'} A^j z_t = (e_{i-i^*})^{'} z_t, \tag{4}$$

which, using *geometric series* results, for the 120-months horizon is:[13]

$$\frac{1}{120} e_{\Delta s}^{'} C = (e_{i-i^*})^{'}, \tag{5}$$

with $C = A(I - A^{120})(I - A)^{-1}$. Similarly, the relative EXPPP over the same horizon imposes the restrictions:

$$\frac{1}{120} e_{\Delta s}^{'} C = \frac{1}{120}(e_{\Delta p - \Delta p^*})^{'} C, \tag{6}$$

and so for the real interest parity:

$$(e_{i-i^*})^{'} = \frac{1}{120}(e_{\Delta p - \Delta p^*})^{'} C, \tag{7}$$

[12] The assumption such that $E(z_{t+j}|z_t) = A^j z_t$ exploits the law of iterated expectations. For a proof see King and Kurmann (2002).

[13] In order for the matrix (I-A) to be invertible its corresponding eigenvalues must lay inside the unit circle. This is clearly the case for the VAR being stationary.

## 3.2 Implied VAR Statistics

The set of restrictions (5)-(7) allows the estimation of the implied slope coefficients that are analogous to the one reported in Section 2. In our 3-dimensional VAR, the implied 120-months regression slope for the UIP is

$$\beta_{120}^{uip} = \frac{\frac{1}{120} e_{\Delta s}' C\Psi (e_{i-i^*})}{(e_{i-i^*})'\Psi(e_{i-i^*})}, \tag{8}$$

where $\Psi$ is the unconditional variance of $z_t$, computed from $vec(\Psi) = (I - A \otimes A)^{-1} vec(\Sigma)$. The numerator in equation (8) is the covariance between the expected future rate of appreciation and the interest rate differential, whereas the denominator is the variance of the interest rate spread. Analogously, for the EXPPP the implied slope coefficients for the 120-month horizon is respectively:[14]

$$\beta_{120}^{ppp} = \frac{\frac{1}{120} e_{\Delta s}' C\Psi C' \frac{1}{120}(e_{\Delta p - \Delta p^*})}{\frac{1}{120}(e_{\Delta p - \Delta p^*})' C\Psi C' \frac{1}{120}(e_{\Delta p - \Delta p^*})}, \tag{9}$$

and similarly for the RIP:

$$\beta_{120}^{rip} = \frac{(e_{i-i^*})'\Psi C' \frac{1}{120}(e_{\Delta p - \Delta p^*})}{\frac{1}{120}(e_{\Delta p - \Delta p^*})' C\Psi C' \frac{1}{120}(e_{\Delta p - \Delta p^*})}. \tag{10}$$

On the same root, to characterize UIP-EXPPP-RIP deviations we start computing three distinct statistics for each condition (Bekaert *et al.*, 2007). The tests are performed following the same set of restrictions as in equations (5)-(7), with $C$ and $\Psi$ fully capturing exchange rates changes, interest rates and inflation spread dynamics in the VAR.[15]

Under the UIP, the expected exchange rate change should be perfectly correlated with the interest rate differential, and they are expected to have equal variability. Hence:

$$CORR^{uip} = corr\left(\frac{1}{l} E_t \sum_{j=1}^{l} \Delta s_{t+j}, i_{t,l} - i_{t,l}^*\right),$$

and

$$VR^{uip} = var\left(\frac{1}{l} E_t \sum_{j=1}^{l} \Delta s_{t+j}\right) / var\left(i_{t,l} - i_{t,l}^*\right).$$

where CORR accounts for correlation and VR is the variance ratio statistics.

Analogously, from EXPPP we would expect the expected exchange rate change to be perfectly correlated with the expected inflation rate differential over the same horizon, and that the two variables

[14]These coefficient are comparable to direct OLS coefficients when $l = 1$, as $C = A$ (e.g., Bekaert and Hodrick, 2001).

[15]All the following statistics are expressed as a function of variances and covariances between the variables, by means of the methodology outlined before.

share the same variability, i.e.

$$CORR^{ppp} = corr\left(\frac{1}{l}E_t\sum_{j=1}^{l}\Delta s_{t+j}, \frac{1}{l}E_t\sum_{j=1}^{l}(\Delta p_{t+j} - \Delta p^*_{t+j})\right),$$

and

$$VR^{ppp} = var\left(\frac{1}{l}E_t\sum_{j=1}^{l}\Delta s_{t+j}\right)/var\left(\frac{1}{l}E_t\sum_{j=1}^{l}(\Delta p_{t+j} - \Delta p^*_{t+j})\right).$$

The same CORR and VR statistics are evaluated for the RIP condition.

Finally, to characterize UIP-EXPPP deviations we calculate the standard deviation (SD) of the residual from each equation, e.g.

$$SD^{uip} = \left[var\left(\epsilon^{uip}_{t+l}\right)\right]^{1/2},$$

where residuals are computed from each equation under the null, e.g. $\epsilon^{uip}_{t+l} = \frac{1}{l}\left(E_t\sum_{j=1}^{l}\Delta s_{t+j}\right) - (i_{t,l} - i^*_{t,l})$.

### 3.3 Montecarlo Analysis

It is well known that standard tests based on lagged dependent variables may lead to over-rejections in small samples. Such a poor sample property arise in the context of the estimation of $AR$ processes, particularly as serial correlation induces non-strict exogeneity in the regressors (Mariott and Pope, 1954; Kendall, 1954).[16] This might turn to be a crucial point when discriminating across near-proximate tests results (e.g. Bekaert *et al.* 1997; 2007).

To bias-correct VAR-coefficients we bootstrap the original VAR-residuals in a *i.i.d.* fashion, so to generate 50.000 data sets. In order to diminish the effect of initial conditions, the temporal bootstrap dimension has been augmented by 1.000 observations (yielding therefore a time series dimension which equals the original number of entries shifted up to 1000 data points) which are then discarded when the estimation is performed.[17] For each of the 50.000 samples we recalculate the VAR parameters. The bias is estimated as the difference between the original VAR parameters and the mean of the new estimates, based on the Montecarlo replications. Bias corrected coefficients are hence obtained by adding back the biases to the original VAR estimates. This yields a set of corrected parameters which are used to construct the point estimates for the betas and the statistics described in Section 2, representing the *alternative* of violation of the parity conditions hypothesis (see Section 3.2).

The empirical distribution of both the coefficients and the statistics is analogously derived by simu-

[16]Being the regressors lagged dependent variables, parameter estimates suffer from small sample bias, although they are consistent.

[17]Standard errors are heteroskedasticity consistent and corrected for MA terms up to the $l-1$ order, using a Newey-West window (see Chinn and Meredith, 2004; 2005). We avoid using the Hansen-Hodrick (1980) estimator as this has the tendency to produce non-positive-definite variance-covariance matrices.

lating 50.000 data samples from the original VAR residuals. At each bootstrap draw, bias correction is implemented on the parameters of interest and new implied VAR coefficients and statistics are obtained. Relevant quantiles are then computed from the empirical distribution obtained as described above. Coefficients and statistics point estimates are reported in Table 4, together with their empirical moments.

# 4 Results from the VAR

In Table 4 we focus on the second row results, reporting bias-corrected estimates.[18] In all cases, point estimates for the UIP are broadly consistent with the ones found in Bekaert *et al.* (2007), although our findings report evidence at a longer horizon (120-month). Based on the estimated $\beta$s in both the UIP and the EXPPP case the expected changes in the nominal exchange rate are found to be negatively correlated with the interest rate spread and with the expected inflation differential respectively.[19] If this is not surprising in the context of UIP - given the existence of a "forward premium" puzzle - it is surprising under the EXPPP hypothesis. Nonetheless, it can be argued that the goods market condition is an *expectational* version of the standard PPP. Hence, a negative sign may not be wrong (albeit it is not obvious) and possibly stands on people bearing on the uncertainty of unforeseen exchange rate and price changes. Also in light of the estimated signs, the $\beta$ coefficients are consistent with the existence of proportionality of the type $\beta^{rip} = \frac{\beta^{ppp}}{\beta^{uip}}$ among the three parity conditions.[20]
The dimension of the correlation coefficients gives further insight on the validity of the UIP, EXPPP and RIP. As the statistics inherit the sign from the implied slope coefficient, in both systems correlation among the numerator and the denominator in each equation - (8), (9) and (10) - is broadly the same (0.9 in absolute value). Indeed, for the RIP to hold one would expect the regressors in (1) and (2) to display the same statistical properties (and hence, the UIP and the EXPPP correlation coefficients to be sensitively close). Alternatively, one might think of more substantial deviations to occur in the EXPPP rather than in the UIP case, or *viceversa* (see also Gokey, 1994).
In both systems (UK vs US and Japan vs. US), the VR for the UIP is below unity, pointing to the absence of a constant volatility ratio among the expected exchange rate changes and the interest rate differential, and to the denominator in $VR^{uip}$ being bigger than the numerator. Alternatively, for the EXPPP and the RIP we find the ratio to be higher than one, suggesting a steadier behaviour of the expected inflation differential with respect to exchange rates and interest rates, and consistent with

[18]The coefficients for the UIP and the RIP are found to be downward biased (for further discussion see Bekaert and Hodrick, 1993; Bekaert *et al.*, 2007), whilst the bias on the PPP coefficient depends on the system considered (an upward bias for the UK - US system and a downward bias for the Japan - US system).

[19]Consistent with Bekaert *et al.* (2007), we find that Meredith and Ching's (2004) finding of UIP better holding at longer horizons - with slope coefficients significantly close to unity - is simply a matter of sample selection.

[20]The proportionality should hold given that the parities are in logs and given the definition of the RIP in (3).

the idea of goods prices to be less volatile than interest rates.

The standard deviation of the errors (SD), capturing the variability of the *risk premia* for each equation (see Section 2.1), is in all cases close (or higher) than unity, being consistent with a time-varying *risk premium* explanation. In this respect, deviations are higher in the UIP case whereas smaller for the EXPPP, corroborating the idea that UIP (and not PPP) rejections are generally more likely at a longer horizon (Chaboud and Wright, 2005; Bekaert *et al.*, 2007).

The remainder of Table 4 reports some information based on the empirical distribution of the implied coefficients and statistics. Together with the standard four moments (mean, variance, skewness and kurtosis), Table 4 reports the fractiles at the 2.5% and 97.5%. A normality Jarque-Brera test is also reported at the bottom of the Table.

Based on the empirical distribution, all slope coefficients fall in between the 2.5% and 97.5% fractiles. For the *UK vs. US system*, when considering a two sided test we reject the null of the UIP slope coefficient being equal to one, i.e. $\beta_l^{uip} = 1$ ($p$-value: 0.001). This is in line with the findings in Bekaert *et al.* (2007) against the UIP using USD-GBS data at different thresholds. The probability of having a unitary coefficient in the EXPPP is accepted instead ($p$-value: 0.570). For the RIP, a one-sided test (the distribution is skewed to the right) confirms the rejection of $\beta^{rip} = 1$ ($p$-value: 0.000), being consistent with a UIP rejection (see Marston, 1997).

For the *Japan - US* system the results report a decisive non-rejection of the three parities. Such a finding support the idea of the UIP to be currency-dependent (Bekaert *et al.*, 2007) rather than horizon-based, whereas for the PPP the evidence goes in the opposite direction (Lothian and Taylor, 1996; Taylor, 2002). The failure of the RIP in the UK vs. US system but not in the Japan vs. US system is moreover consistent with the assumption that any couple of parity among UIP, EXPPP and RIP necessarily implies the third relation (Marston, 1997). For the EXPPP and the RIP, the estimated confidence intervals (2.5% and 97.5% fractiles) are nonetheless large, increasing the likelihood of type I error.

## 4.1 Augmenting the VAR

Since the UIP and the EXPPP coefficients estimates are non-positive (Table 4), we have assessed the possible role of expectations mispecifications by augmenting the information set to include macroeconomic and financial variables, some expected to feature forward looking properties. With reference to a *present value* model (Campbell and Shiller, 1987), this allows conditioning the information set agents use in forming expectations as follows

$$E(z_{t+j}|z_t, H_t) = A^j z_t,$$

where $H_t$ includes: (i) industrial production growth differentials, i.e. $\Delta y_t - \Delta y_t^*$; (ii) broad money aggregates growth differentials (M3), i.e. $\Delta m_t - \Delta m_t^*$; (iii) reserve assets growth differentials, i.e. $\Delta ra_t - \Delta ra_t^*$; and (iv) share price return differentials, i.e. $\Delta sp_t - \Delta sp_t^*$. In all cases, differentials are considered as "domestic vs. US" differences. The inclusion of variables besides the one predicted by economic theory is in line with the literature describing the evolution of exchange rates as a function of macroeconomic fundamentals other than prices (PPP) and interest rates (UIP). Exogenous regressors are primarily aimed at capturing cross-country macroeconomic developments and international imbalances. For sake of simplicity, the regressors are let entering the VAR only contemporaneously.

We summarize our main results in Table 5 to Table 8, whereas the histograms of the newly simulated beta coefficients are reported in Figure 3 to 6. Once again, the focus is on bias-corrected results. As rejection of the UIP and RIP hypothesis is found only for the UK - US system, in what follows we focus on the results for the latter pair of countries.

Overall, the results are not sensitively affected by the inclusion of exogenous regressors, as a "forward premium" bias persists in all cases.

Considering productivity growth differentials and foreign reserve assets, the $p$-value for the non-rejection of the null $H_0 : \beta^{uip,rip} = 1$ (in absolute value) does not increase (see Table 5 and 7). The broad money (M3) growth differential helps instead reducing the coefficient bias in the RIP equation for the UK - US system, showing a better fit of the RIP distribution by reducing skewness (right). Nonetheless, also in this case there is yet not clear evidence of a non-rejection at the 5% significance level (see Table 7). Conditioning on the share price return differential analogously helps center the distribution for the UIP in the UK - US system over a mean value of about -0.6, yet not allowing to reject the null $\beta^{uip} = -1$ at the 2.5% level (the test is one-sided). This result deserves further discussion as share prices reflect investors confidence in the stock market evaluation in each period, hence having a strong forward-looking component. Although the "forward premium bias" does not disappear, the above result possibly provides (weak) evidence in favour of theories predicting foreign exchange rate *premia* (e.g. Fama, 1987) in explaining UIP deviations. Indeed, share prices, by proxing the perceptions about future cyclical economic developments, have an impact in shaping the magnitude of the UIP coefficient.

# 5 Conclusions

Drawing on a VAR approach, in this paper we revisited the evidence on the uncovered interest parity (UIP), the *ex ante* purchasing power parity (EXPPP) and the real interest parity (RIP) for the UK vs. US and the Japan vs. US data from 1975-2008. The evidence is based on developing some joint coefficient-based tests obtained by rewriting our theoretical relations as a set of cross-equation restrictions in the VAR (Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007). The results point out the existence of a "forward premium" bias in both the UIP and the EXPPP equations, consistently with the existence of some form of proportionality among the three parities. A "forward premium bias" in the EXPPP case is new in the literature and can be explained by arguing that the tested equation is an *expectational* version of the standard PPP, i.e. implying people bearing on the uncertainty of future exchange rate and price changes.

To better characterize these results, we augment the original VAR framework with exogenous regressors. In this respect, we find the results not to be sensitively affected by the inclusion of exogenous variables, as a "forward premium" bias persists in all cases. Alternatively, conditioning on share prices return differentials yields a better fitting of the UIP relation in the UK - US system. This result provides (weak) support to the role of foreign exchange rate *premia* (e.g. Fama, 1987) in explaining UIP deviations.

The overall results are moreover consistent with the idea of the UIP to be currency-dependent (Bekaert *et al.*, 2007) rather than horizon-based, whilst for the EXPPP the evidence goes in the opposite direction (Lothian and Taylor, 1996; Taylor, 2002). The failure of RIP in the UK vs. US system but not in the Japan vs. US system is consistent with the assumption that any couple of parity among UIP, EXPPP and RIP necessarily implies the third relation (Marston, 1997). In all cases, the statistical explanation of the results must be however taken cautiously because of the large standard errors associated with the EXPPP and RIP coefficient estimates.

Overall the results are in line with the literature describing the evolution of exchange rates as a function of macroeconomic fundamentals other than interest rates (UIP) and prices (EXPPP). Future research could fruitfully be devoted to the assessment of the role of economic fundamentals in shaping international exchange rate and inflation dynamics, together with their expectations.

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Table 1: Summary and Descriptive Statistics

| | $\Delta s_t$ | $i_t - i_t$ | $\Delta p_t - \Delta p_t^*$ |
|---|---|---|---|
| *UK - US System* | | | |
| Mean | 0.561 | 1.396 | 1.060 |
| Variance of sample mean | 1.458 | 0.089 | 0.361 |
| Maximum | 132.185 | 8.300 | 42.386 |
| Minimum | -120.312 | -2.0300 | -17.852 |
| AC(1) | 0.330 | 0.973 | 0.252 |
| AC(2) | -0.018 | 0.930 | 0.096 |
| AC(3) | 0.040 | 0.896 | 0.043 |
| AC(4) | 0.055 | 0.876 | 0.062 |
| *Japan - US System* | | | |
| Mean | -3.086 | -2.959 | -2.374 |
| Variance of sample mean | 1.665 | 0.058 | 0.325 |
| Maximum | 95.588 | -0.858 | 25.024 |
| Minimum | -123.181 | -4.969 | -24.844 |
| AC(1) | 0.297 | 0.965 | 0.081 |
| AC(2) | 0.040 | 0.919 | -0.200 |
| AC(3) | 0.060 | 0.885 | -0.094 |
| AC(4) | 0.033 | 0.857 | 0.016 |

*Notes*: The *cpi*-inflation and the appreciation rate are taken as month-on-month changes. Figures for the nominal exchange rate and inflation have been multiplied by 1200.

Table 2: Lag-length Selection Criteria

| Lags | Akaike Information Criterion | Bayesian Information Criterion | Hannan-Quinn |
|---|---|---|---|
| *UK - US System* | | | |
| 0 | 7964.458 | 7976.39 | 7969.172 |
| 1 | 6738.387 | 6785.929< | 6757.06 |
| 2 | 6703.745 | 6786.617 | 6736.095 |
| 3 | 6681.621< | 6799.536 | 6727.362< |
| 4 | 6686.466 | 6839.129 | 6745.303 |
| *Japan - US System* | | | |
| 0 | 4227.568 | 4237.847 | 4231.694 |
| 1 | 3593.353 | 3634.149< | 3609.534< |
| 2 | 3581.827 | 3652.643 | 3609.567 |
| 3 | 3571.399< | 3671.72 | 3610.183 |
| 4 | 3580.753 | 3710.038 | 3630.041 |

Table 3: VAR Dynamics and Bias-Corrected Coefficients

| | $\Delta s_t$ | $i_t - i_t^*$ | $\Delta p_t - \Delta p_t^*$ |
|---|---|---|---|
| *UK - US System* | | | |
| $\Delta s_{t-1}$ | 0.394 | 0.001 | 0.035 |
| Bias-corrected | 0.386 | 0.001 | 0.035 |
| (s.e.) | (0.051) | (0.001) | (0.012) |
| $i_{t-1} - i_{t-1}^*$ | -5.077 | 1.321 | 1.428 |
| Bias-corrected | -5.143 | 1.312 | 1.443 |
| (s.e.) | (3.677) | (0.049) | (0.871) |
| $\Delta p_{t-1} - \Delta p_{t-1}^*$ | 0.011 | 0.005 | 0.134 |
| Bias-corrected | 0.009 | 0.005 | 0.127 |
| (s.e.) | (0.206) | (0.003) | (0.049) |
| $\Delta s_{t-2}$ | -0.180 | 0.000 | -0.004 |
| Bias-corrected | -0.182 | 0.000 | -0.004 |
| (s.e.) | (0.054) | (0.001) | (0.013) |
| $i_{t-2} - i_{t-2}^*$ | 7.246 | -0.590 | -0.948 |
| Bias-corrected | 7.264 | -0.584 | -0.945 |
| (s.e.) | (5.734) | (0.076) | (1.359) |
| $\Delta p_{t-2} - \Delta p_{t-2}^*$ | 0.382 | 0.003 | -0.044 |
| Bias-corrected | 0.381 | 0.003 | -0.048 |
| (s.e.) | (0.209) | (0.003) | (0.049) |
| $\Delta s_{t-3}$ | 0.095 | 0.001 | 0.014 |
| Bias-corrected | 0.089 | 0.001 | 0.014 |
| (s.e.) | (0.051) | (0.001) | (0.012) |
| $i_{t-3} - i_{t-3}^*$ | -3.497 | 0.240 | 0.752 |
| Bias-corrected | -3.611 | 0.234 | 0.749 |
| (s.e.) | (3.594) | (0.048) | (0.851) |
| $\Delta p_{t-3} - \Delta p_{t-3}^*$ | 0.059 | -0.007 | -0.079 |
| Bias-corrected | 0.055 | -0.007 | -0.082 |
| (s.e.) | (0.205) | (0.003) | (0.049) |
| *Japan - US System* | | | |
| $\Delta s_{t-1}$ | 0.294 | 0.001 | 0.008 |
| Bias-corrected | 0.287 | 0.001 | 0.008 |
| (s.e.) | (0.048) | (0.001) | (0.010) |
| $i_{t-1} - i_{t-1}^*$ | -0.631 | 0.959 | 1.128 |
| Bias-corrected | -0.647 | 0.949 | 1.117 |
| (s.e.) | (0.960) | (0.012) | (0.203) |
| $\Delta p_{t-1} - \Delta p_{t-1}^*$ | -0.162 | 0.005 | 0.162 |
| Bias-corrected | -0.164 | 0.005 | 0.156 |
| (s.e.) | (0.234) | (0.003) | (0.050) |

*Notes*: In the Table we report both actual and bias-corrected coefficients. The coefficients are SUR regression estimates with robust standard errors (Newey-West/Bartlett), where we correct for MA terms up to the $l-1$ order (see Chinn and Meredith, 2004; 2005).

Table 4: Three Parity Conditions at Long Horizon

| | Coefficients | | | Additional Statistics | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | UIP | | | PPP | | | RIP | | |
| | $\beta^{uip}$ | $\beta^{ppp}$ | $\beta^{rip}$ | CORR | VR | SD | CORR | VR | SD | CORR | VR | SD |
| *UK - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.257 | -0.858 | 3.365 | -0.963 | 0.071 | 2.058 | -0.948 | 0.819 | 0.907 | 0.993 | 11.485 | 1.157 |
| Bias corrected | -0.234 | -1.030 | 4.457 | -0.950 | 0.061 | 1.770 | -0.929 | 1.230 | 0.658 | 0.989 | 20.297 | 1.118 |
| Mean | -0.222 | -1.334 | 5.859 | -0.619 | 0.113 | 1.631 | -0.573 | 5.642 | 0.586 | 0.975 | 46.174 | 1.067 |
| Median | -0.220 | -1.134 | 5.285 | -0.924 | 0.065 | 1.590 | -0.880 | 1.954 | 0.536 | 0.983 | 29.065 | 1.063 |
| Max | 0.996 | 16.669 | 28.864 | 0.996 | 1.954 | 4.796 | 0.999 | 981.914 | 3.700 | 0.998 | 6705.4 | 1.661 |
| Min | -1.391 | -19.153 | -65.025 | -0.998 | 0.001 | 0.186 | -0.999 | 0.003 | 0.047 | -0.885 | 2.206 | 0.597 |
| St.dev. | 0.231 | 1.606 | 2.562 | 0.590 | 0.133 | 0.413 | 0.602 | 16.534 | 0.333 | 0.040 | 82.284 | 0.118 |
| Skewness | -0.046 | -1.153 | 1.098 | 1.665 | 2.620 | 0.621 | 1.558 | 19.441 | 1.019 | -20.686 | 26.470 | 0.178 |
| Kurtosis | 3.318 | 9.132 | 24.615 | 4.292 | 14.103 | 3.877 | 3.959 | 672.3 | 4.783 | 700.3 | 1566.4 | 3.020 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -0.684 | -5.073 | 2.810 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.227 | 1.157 | 12.351 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.001 | 0.570 | 0.000 | – | – | – | – | – | – | – | – | – |
| *Japan - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.906 | -15.113 | 16.397 | -0.988 | 0.841 | 1.585 | -0.988 | 233.765 | 0.810 | 0.983 | 278.000 | 0.780 |
| Bias corrected | -0.549 | -14.178 | 24.777 | -0.962 | 0.325 | 1.058 | -0.963 | 216.932 | 0.413 | 0.959 | 667.488 | 0.654 |
| Mean | -0.425 | -3.896 | 4.812 | -0.566 | 0.461 | 0.896 | -0.299 | 883.877 | 0.378 | 0.332 | 2128.1 | 0.592 |
| Median | -0.403 | -3.918 | 10.718 | -0.919 | 0.230 | 0.841 | -0.696 | 102.779 | 0.296 | 0.899 | 430.767 | 0.585 |
| Max | 1.492 | 639.799 | 577.633 | 0.999 | 18.371 | 6.749 | 1.000 | 1028892 | 5.788 | 1.000 | 3000667 | 1.192 |
| Min | -4.284 | -445.854 | -1325 | -1.000 | 0.001 | 0.037 | -1.000 | 0.076 | 0.009 | -1.000 | 7.352 | 0.313 |
| St.dev. | 0.496 | 19.598 | 30.903 | 0.644 | 0.622 | 0.393 | 0.745 | 7569.9 | 0.285 | 0.833 | 16297 | 0.099 |
| Skewness | -0.278 | 0.573 | -2.425 | 1.438 | 3.734 | 1.007 | 0.620 | 75.931 | 1.828 | -0.746 | 130.132 | 0.408 |
| Kurtosis | 3.439 | 57.997 | 87.423 | 3.477 | 39.586 | 5.852 | 1.707 | 8691.2 | 10.452 | 1.706 | 23072.2 | 3.227 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -1.460 | -41.435 | -57.045 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.492 | 33.395 | 55.057 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.121 | 0.930 | 0.998 | – | – | – | – | – | – | – | – | – |

*Notes*: The first two rows of the Table reports the point estimates for the beta coefficients and for the implied VAR statisitcs, both actual and bias-corrected. CORR is the correlation statistics, VR is the variance ratio and SD is the standard error of the residual in each equation under the null of, i.e. UIP. The remainder of the Table reports the empirical distribution obtained by 50.000 simulations. Jarque-Bera refers to the $p$-value of the Jarque-Bera normality test.

Table 5: Three Parity Conditions at Long Horizon. Conditioning on Productivity Growth Differential

| | Coefficients | | | Additional Statistics | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | UIP | | | PPP | | | RIP | | |
| | $\beta^{uip}$ | $\beta^{ppp}$ | $\beta^{rip}$ | CORR | VR | SD | CORR | VR | SD | CORR | VR | SD |
| *UK - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.260 | -0.856 | 3.319 | -0.963 | 0.073 | 2.062 | -0.949 | 0.813 | 0.918 | 0.993 | 11.164 | 1.150 |
| Bias corrected | -0.237 | -0.998 | 4.261 | -0.952 | 0.062 | 1.794 | -0.932 | 1.146 | 0.685 | 0.990 | 18.510 | 1.115 |
| Mean | -0.226 | -1.260 | 5.478 | -0.636 | 0.112 | 1.660 | -0.594 | 4.551 | 0.613 | 0.980 | 38.218 | 1.066 |
| Median | -0.223 | -1.089 | 4.981 | -0.929 | 0.065 | 1.620 | -0.889 | 1.731 | 0.563 | 0.985 | 25.626 | 1.062 |
| Max | 0.820 | 5.647 | 39.571 | 0.997 | 1.694 | 4.074 | 0.999 | 1440.3 | 2.996 | 0.999 | 3316.9 | 1.555 |
| Min | -1.287 | -25.152 | -1.378 | -0.998 | 0.001 | 0.275 | -0.999 | 0.007 | 0.047 | -0.029 | 2.185 | 0.646 |
| St.dev. | 0.226 | 1.440 | 2.241 | 0.579 | 0.130 | 0.414 | 0.591 | 13.015 | 0.339 | 0.023 | 54.671 | 0.117 |
| Skewness | -0.039 | -1.289 | 2.009 | 1.743 | 2.540 | 0.606 | 1.643 | 39.605 | 0.967 | -10.288 | 17.110 | 0.199 |
| Kurtosis | 3.293 | 9.114 | 11.264 | 4.573 | 13.093 | 3.713 | 4.245 | 3424.8 | 4.424 | 245 | 652.9 | 3.003 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -0.678 | -4.597 | 2.700 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.218 | 1.048 | 11.196 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.001 | 0.545 | 0.000 | – | – | – | – | – | – | – | – | – |
| *Japan - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.981 | -16.068 | 16.099 | -0.989 | 0.984 | 1.651 | -0.990 | 263.4 | 0.875 | 0.984 | 267.757 | 0.782 |
| Bias corrected | -0.638 | -15.566 | 23.476 | -0.970 | 0.432 | 1.138 | -0.971 | 257.1 | 0.482 | 0.963 | 594.399 | 0.664 |
| Mean | -0.515 | -4.471 | 5.084 | -0.657 | 0.549 | 0.973 | -0.341 | 1134 | 0.427 | 0.354 | 2143.7 | 0.603 |
| Median | -0.493 | -4.829 | 10.672 | -0.942 | 0.295 | 0.914 | -0.781 | 118.8 | 0.345 | 0.917 | 408.8 | 0.595 |
| Max | 1.482 | 1129.487 | 438.617 | 0.999 | 13.512 | 5.070 | 1.000 | 4816664 | 4.155 | 1.000 | 7618704 | 1.157 |
| Min | -3.671 | -665.716 | -634.969 | -1.000 | 0.001 | 0.056 | -1.000 | 0.125 | 0.018 | -1.000 | 4.691 | 0.303 |
| St.dev. | 0.498 | 21.388 | 30.090 | 0.576 | 0.696 | 0.411 | 0.746 | 26933 | 0.312 | 0.829 | 35943 | 0.100 |
| Skewness | -0.289 | 3.163 | -0.730 | 1.845 | 2.981 | 0.966 | 0.717 | 146.619 | 1.578 | -0.799 | 192.372 | 0.465 |
| Kurtosis | 3.345 | 228.48 | 24.512 | 4.946 | 19.225 | 4.950 | 1.820 | 23952 | 7.067 | 1.782 | 40398 | 3.358 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -1.556 | -43.141 | -56.167 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.396 | 36.127 | 54.482 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.150 | 0.939 | 0.998 | – | – | – | – | – | – | – | – | – |

*Notes*: The first two rows of the Table reports the point estimates for the beta coefficients and for the implied VAR statisitcs, both actual and bias-corrected. CORR is the correlation statistics, VR is the variance ratio and SD is the standard error of the residual in each equation under the null of, i.e. UIP. The remainder of the Table reports the empirical distribution obtained by 50.000 simulations. Jarque-Bera refers to the $p$-value of the Jarque-Bera normality test.

Table 6: Three Parity Conditions at Long Horizon. Conditioning Broad Money (M3) Growth Differential

| | Coefficients | | | Additional Statistics | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | UIP | | | PPP | | | RIP | | |
| | $\beta^{uip}$ | $\beta^{ppp}$ | $\beta^{rip}$ | CORR | VR | SD | CORR | VR | SD | CORR | VR | SD |
| *UK - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.371 | -2.489 | 6.864 | -0.973 | 0.145 | 1.268 | -0.939 | 7.026 | 0.479 | 0.988 | 48.311 | 0.792 |
| Bias corrected | -0.303 | -2.619 | 9.082 | -0.956 | 0.100 | 1.063 | -0.884 | 8.768 | 0.337 | 0.972 | 87.305 | 0.729 |
| Mean | -0.273 | -2.309 | 10.123 | -0.587 | 0.182 | 0.967 | -0.420 | 32.031 | 0.322 | 0.874 | 194.993 | 0.682 |
| Median | -0.266 | -2.160 | 9.973 | -0.906 | 0.099 | 0.937 | -0.704 | 12.834 | 0.275 | 0.943 | 123.575 | 0.677 |
| Max | 1.166 | 17.780 | 31.203 | 0.997 | 5.727 | 3.641 | 0.999 | 1750.511 | 2.675 | 0.999 | 2935.2 | 1.163 |
| Min | -2.380 | -28.241 | -26.453 | -0.998 | 0.002 | 0.083 | -0.998 | 0.043 | 0.032 | -0.925 | 3.154 | 0.378 |
| St.dev. | 0.299 | 3.288 | 3.902 | 0.609 | 0.231 | 0.295 | 0.610 | 60.040 | 0.216 | 0.196 | 205.5 | 0.093 |
| Skewness | -0.163 | -0.328 | -0.775 | 1.563 | 3.472 | 0.713 | 1.128 | 6.271 | 1.388 | -3.701 | 2.764 | 0.282 |
| Kurtosis | 3.550 | 4.445 | 8.597 | 3.924 | 29.699 | 4.243 | 2.863 | 73.623 | 6.163 | 20.380 | 14.553 | 3.071 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -0.880 | -9.294 | 4.175 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.299 | 3.825 | 17.648 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.007 | 0.785 | 0.010 | – | – | – | – | – | – | – | – | – |
| *Japan - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.719 | -22.863 | 30.403 | -0.979 | 0.540 | 1.387 | -0.969 | 556.586 | 0.615 | 0.947 | 1030.650 | 0.780 |
| Bias corrected | -0.427 | -20.215 | 42.138 | -0.935 | 0.208 | 0.949 | -0.912 | 491.128 | 0.314 | 0.868 | 2358.519 | 0.649 |
| Mean | -0.530 | -2.904 | 1.705 | -0.534 | 0.808 | 1.125 | -0.216 | 1414.170 | 0.556 | 0.179 | 0.681 | 1824.566 |
| Median | -0.468 | -2.990 | 7.185 | -0.939 | 0.324 | 0.975 | -0.634 | 104.590 | 0.386 | 0.833 | 0.658 | 310.846 |
| Max | 3.246 | 1537.130 | 587.119 | 1.000 | 96.956 | 24.968 | 1.000 | 2794618 | 23.679 | 1.000 | 2.714 | 1648784 |
| Min | -9.849 | -528.250 | -663.727 | -1.000 | 0.001 | 0.027 | -1.000 | 0.071 | 0.019 | -1.000 | 0.316 | 0.754 |
| St.dev. | 0.702 | 25.803 | 28.811 | 0.699 | 1.500 | 0.717 | 0.801 | 20908.7 | 0.561 | 0.897 | 0.156 | 11988 |
| Skewness | -0.755 | 5.521 | -0.604 | 1.311 | 11.895 | 3.150 | 0.429 | 94.083 | 4.768 | -0.387 | 1.283 | 69.068 |
| Kurtosis | 5.681 | 329.454 | 29.528 | 3.012 | 457.173 | 42.491 | 1.426 | 10971.710 | 85.882 | 1.249 | 7.228 | 7837.8 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -2.081 | -48.775 | -55.566 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.679 | 42.116 | 51.724 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.194 | 0.930 | 0.998 | – | – | – | – | – | – | – | – | – |

*Notes*: The first two rows of the Table reports the point estimates for the beta coefficients and for the implied VAR statisitcs, both actual and bias-corrected. CORR is the correlation statistics, VR is the variance ratio and SD is the standard error of the residual in each equation under the null of, i.e. UIP. The remainder of the Table reports the empirical distribution obtained by 50.000 simulations. Jarque-Bera refers to the $p$-value of the Jarque-Bera normality test.

Table 7: Three Parity Conditions at Long Horizon. Conditioning on Reserve Assets Growth Differential

| | Coefficients | | | Additional Statistics | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | UIP | | | PPP | | | RIP | | |
| | $\beta^{uip}$ | $\beta^{ppp}$ | $\beta^{rip}$ | CORR | VR | SD | CORR | VR | SD | CORR | VR | SD |
| *UK - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.126 | -0.298 | 2.387 | -0.904 | 0.019 | 2.081 | -0.895 | 0.111 | 1.007 | 0.996 | 5.740 | 1.080 |
| Bias corrected | -0.111 | -0.346 | 3.169 | -0.865 | 0.016 | 1.755 | -0.850 | 0.166 | 0.675 | 0.994 | 10.155 | 1.086 |
| Mean | -0.103 | -0.477 | 4.199 | -0.303 | 0.082 | 1.600 | -0.280 | 1.758 | 0.592 | 0.988 | 21.599 | 1.043 |
| Median | -0.102 | -0.379 | 3.828 | -0.797 | 0.040 | 1.549 | -0.762 | 0.634 | 0.514 | 0.991 | 14.936 | 1.040 |
| Max | 1.266 | 5.552 | 26.368 | 0.999 | 1.930 | 4.958 | 1.000 | 225.1 | 3.747 | 0.999 | 1357 | 1.576 |
| Min | -1.380 | -12.124 | 1.110 | -0.998 | 0.001 | 0.132 | -1.000 | 0.002 | 0.046 | 0.079 | 1.235 | 0.308 |
| St.dev. | 0.254 | 1.129 | 1.730 | 0.785 | 0.115 | 0.471 | 0.790 | 4.484 | 0.394 | 0.013 | 27.066 | 0.116 |
| Skewness | -0.019 | -0.934 | 2.149 | 0.657 | 3.395 | 0.707 | 0.628 | 16.595 | 1.266 | -15.546 | 10.989 | 0.159 |
| Kurtosis | 3.419 | 6.835 | 12.542 | 1.661 | 21.424 | 4.174 | 1.624 | 523.704 | 5.511 | 740.655 | 278.61 | 3.195 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -0.609 | -2.977 | 2.077 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.393 | 1.451 | 8.554 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.001 | 0.315 | 0.000 | – | – | – | – | – | – | – | – | – |
| *Japan - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.885 | -15.890 | 17.598 | -0.986 | 0.806 | 1.512 | -0.988 | 258.441 | 0.762 | 0.983 | 320.658 | 0.756 |
| Bias corrected | -0.548 | -13.859 | 24.246 | -0.961 | 0.325 | 1.038 | -0.968 | 205.157 | 0.406 | 0.965 | 631.532 | 0.642 |
| Mean | -0.684 | -4.308 | 3.998 | -0.629 | 1.086 | 1.262 | -0.329 | 1567.639 | 0.670 | 0.364 | 1774.788 | 0.673 |
| Median | -0.599 | -4.547 | 8.089 | -0.963 | 0.435 | 1.073 | -0.845 | 105.387 | 0.461 | 0.953 | 238.128 | 0.649 |
| Max | 4.132 | 818.529 | 858.573 | 1.000 | 67.681 | 19.156 | 1.000 | 963727 | 16.999 | 1.000 | 4677161 | 2.933 |
| Min | -8.228 | -734.946 | -467.079 | -1.000 | 0.002 | 0.034 | -1.000 | 0.026 | 0.025 | -1.000 | 1.941 | 0.289 |
| St.dev. | 0.765 | 27.141 | 27.716 | 0.644 | 1.882 | 0.827 | 0.794 | 14237.950 | 0.680 | 0.854 | 23132.220 | 0.157 |
| Skewness | -0.836 | 0.557 | -0.514 | 1.681 | 6.211 | 2.799 | 0.690 | 34.885 | 3.584 | -0.809 | 168.319 | 1.403 |
| Kurtosis | 5.127 | 75.60 | 35.92 | 4.165 | 96.734 | 25.247 | 1.699 | 1640.436 | 34.836 | 1.756 | 33465.970 | 8.890 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -2.444 | -52.989 | -53.731 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.582 | 44.655 | 52.126 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.284 | 0.937 | 0.999 | – | – | – | – | – | – | – | – | – |

*Notes*: The first two rows of the Table reports the point estimates for the beta coefficients and for the implied VAR statisitcs, both actual and bias-corrected. CORR is the correlation statistics, VR is the variance ratio and SD is the standard error of the residual in each equation under the null of, i.e. UIP. The remainder of the Table reports the empirical distribution obtained by 50.000 simulations. Jarque-Bera refers to the $p$-value of the Jarque-Bera normality test.

Table 8: Three Parity Conditions at Long Horizon. Conditioning on Share Prices Return Differential

| | Coefficients | | | Additional Statistics | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | UIP | | | PPP | | | RIP | | |
| | $\beta^{uip}$ | $\beta^{ppp}$ | $\beta^{rip}$ | CORR | VR | SD | CORR | VR | SD | CORR | VR | SD |
| *UK - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.667 | -1.375 | 2.061 | -0.997 | 0.448 | 3.128 | -0.994 | 1.914 | 2.159 | 0.996 | 4.277 | 0.975 |
| Bias corrected | -0.593 | -1.686 | 2.844 | -0.995 | 0.355 | 2.473 | -0.989 | 2.905 | 1.463 | 0.994 | 8.184 | 1.014 |
| Mean | -0.551 | -2.142 | 3.900 | -0.939 | 0.384 | 2.190 | -0.921 | 7.285 | 1.219 | 0.987 | 19.461 | 0.977 |
| Median | -0.546 | -1.918 | 3.510 | -0.990 | 0.305 | 2.120 | -0.980 | 3.870 | 1.136 | 0.991 | 12.572 | 0.976 |
| Max | 0.899 | 2.935 | 31.753 | 0.997 | 4.445 | 7.825 | 0.999 | 1032 | 8.067 | 1.000 | 2662 | 1.478 |
| Min | -2.107 | -26.031 | -0.300 | -1.000 | 0.001 | 0.207 | -1.000 | 0.004 | 0.058 | -0.011 | 0.857 | 0.219 |
| St.dev. | 0.271 | 1.427 | 1.702 | 0.240 | 0.329 | 0.595 | 0.249 | 16.988 | 0.559 | 0.018 | 37.451 | 0.112 |
| Skewness | -0.122 | -1.803 | 2.363 | 6.234 | 1.804 | 0.837 | 5.892 | 24.772 | 1.110 | -18.057 | 32.812 | 0.010 |
| Kurtosis | 3.522 | 11.979 | 15.465 | 43.264 | 8.846 | 4.670 | 39.392 | 1078.653 | 5.923 | 695.989 | 1792.22 | 3.643 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -1.105 | -5.614 | 1.891 | – | – | – | – | – | – | – | – | – |
| 97.5% | -0.030 | -0.070 | 8.261 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.025 | 0.800 | 0.000 | – | – | – | – | – | – | – | – | – |
| *Japan - US System* | | | | | | | | | | | | |
| Not bias corrected | -0.903 | -15.148 | 16.487 | -0.988 | 0.836 | 1.578 | -0.988 | 234.908 | 0.805 | 0.983 | 281.140 | 0.779 |
| Bias corrected | -0.582 | -14.237 | 23.569 | -0.967 | 0.362 | 1.099 | -0.968 | 216.348 | 0.444 | 0.964 | 597.207 | 0.664 |
| Mean | -0.715 | -4.513 | 4.116 | -0.665 | 1.090 | 1.328 | -0.343 | 1686.349 | 0.705 | 0.383 | 1705.611 | 0.696 |
| Median | -0.643 | -4.841 | 8.219 | -0.969 | 0.475 | 1.146 | -0.866 | 111.160 | 0.503 | 0.955 | 239.578 | 0.673 |
| Max | 2.366 | 910.175 | 486.742 | 1.000 | 54.370 | 19.208 | 1.000 | 2369952 | 16.264 | 1.000 | 2063715 | 2.944 |
| Min | -7.363 | -450.049 | -841.571 | -1.000 | 0.001 | 0.034 | -1.000 | 0.069 | 0.022 | -1.000 | 1.172 | 0.150 |
| St.dev. | 0.740 | 27.330 | 27.634 | 0.616 | 1.738 | 0.827 | 0.791 | 21184.530 | 0.681 | 0.845 | 13743 | 0.157 |
| Skewness | -0.714 | 1.801 | -1.457 | 1.848 | 5.079 | 2.419 | 0.725 | 70.273 | 3.131 | -0.858 | 97.935 | 1.158 |
| Kurtosis | 4.548 | 92.067 | 36.672 | 4.791 | 61.181 | 19.305 | 1.748 | 6722.818 | 26.748 | 1.844 | 13049.970 | 6.736 |
| Jarque-Bera | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| 2.5% | -2.372 | -53.215 | -53.898 | – | – | – | – | – | – | – | – | – |
| 97.5% | 0.540 | 45.757 | 51.542 | – | – | – | – | – | – | – | – | – |
| $H_0 : \beta = 1$ $p$-value | 0.282 | 0.940 | 0.998 | – | – | – | – | – | – | – | – | – |

*Notes*: The first two rows of the Table reports the point estimates for the beta coefficients and for the implied VAR statisitcs, both actual and bias-corrected. CORR is the correlation statistics, VR is the variance ratio and SD is the standard error of the residual in each equation under the null of, i.e. UIP. The remainder of the Table reports the empirical distribution obtained by 50.000 simulations. Jarque-Bera refers to the $p$-value of the Jarque-Bera normality test.

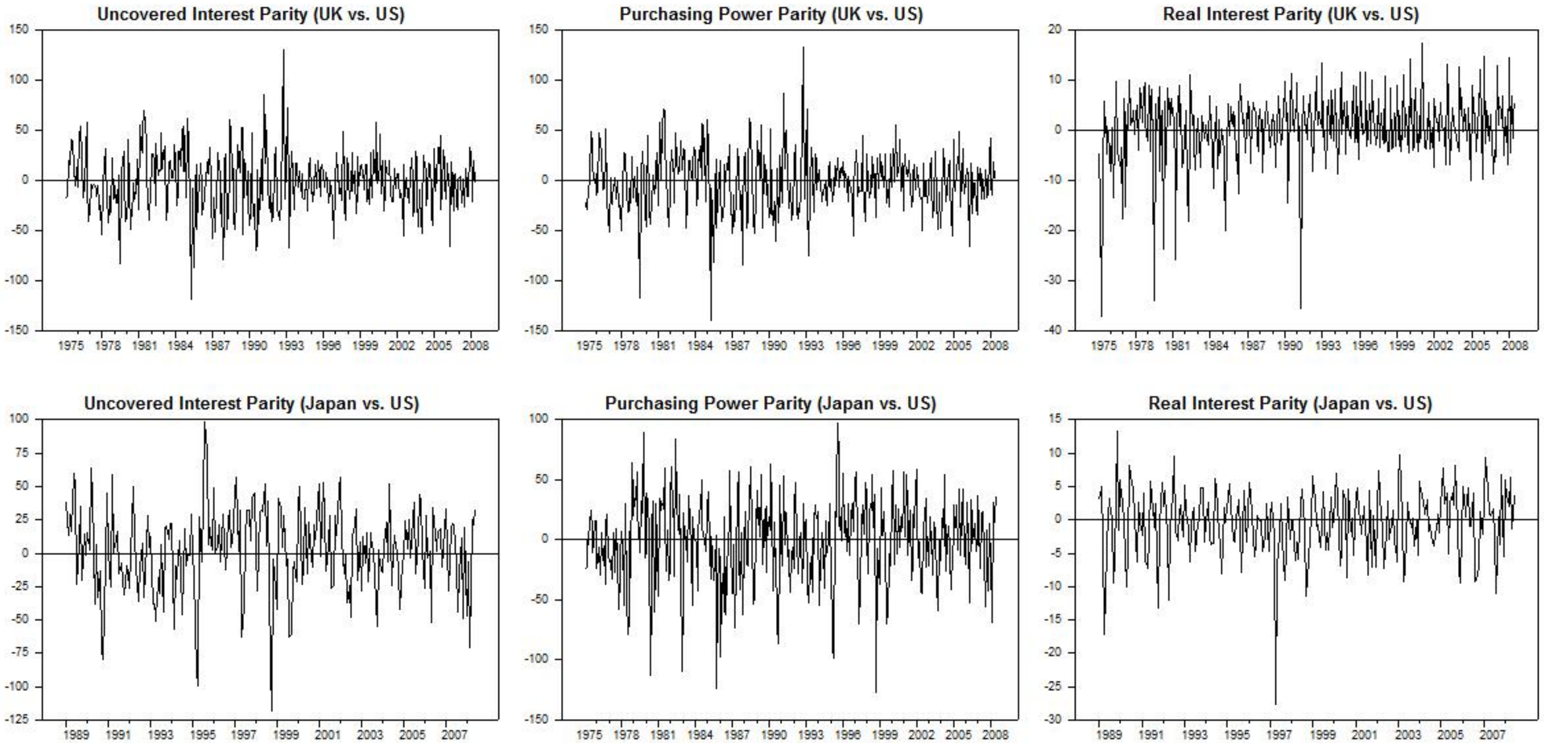


Figure 1: Deviations from International Parity Conditions

*Notes*: Uncovered interest parity (left panel), relative purchasing power parity (middle panel) and real interest parity (right panel). All parities are computed as month-on-month changes. In each UIP and PPP parity, the first part of the sample is dominated by the big swings in the spot nominal exchange rate of the mid and late 70s.

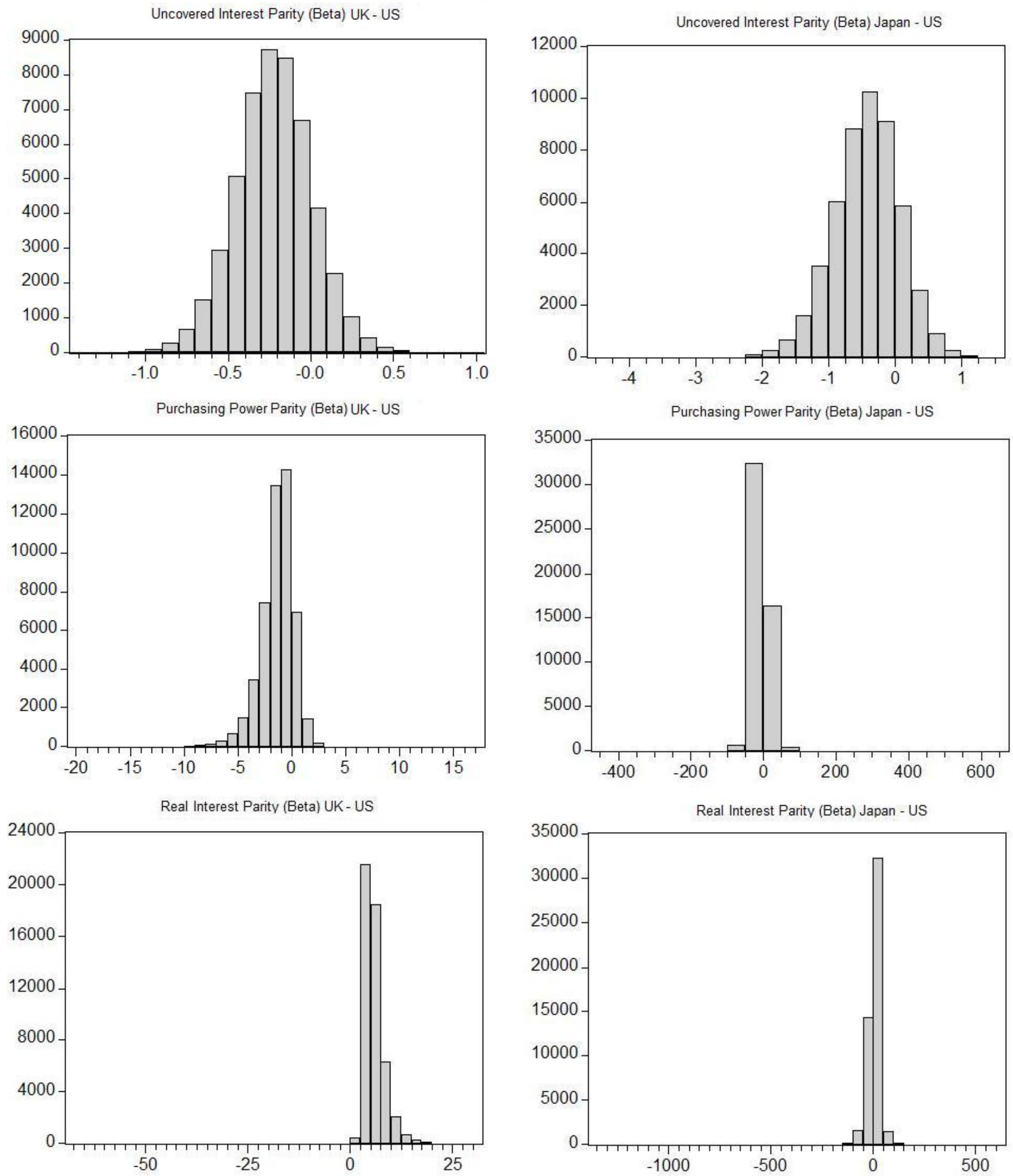


Figure 2: Histograms of the Simulated Beta Coefficients

*Notes*: Uncovered interest parity (upper panel), relative purchasing power parity (mid panel) and real interest parity (lower panel, below). The distribution is based on 50.000 replications.

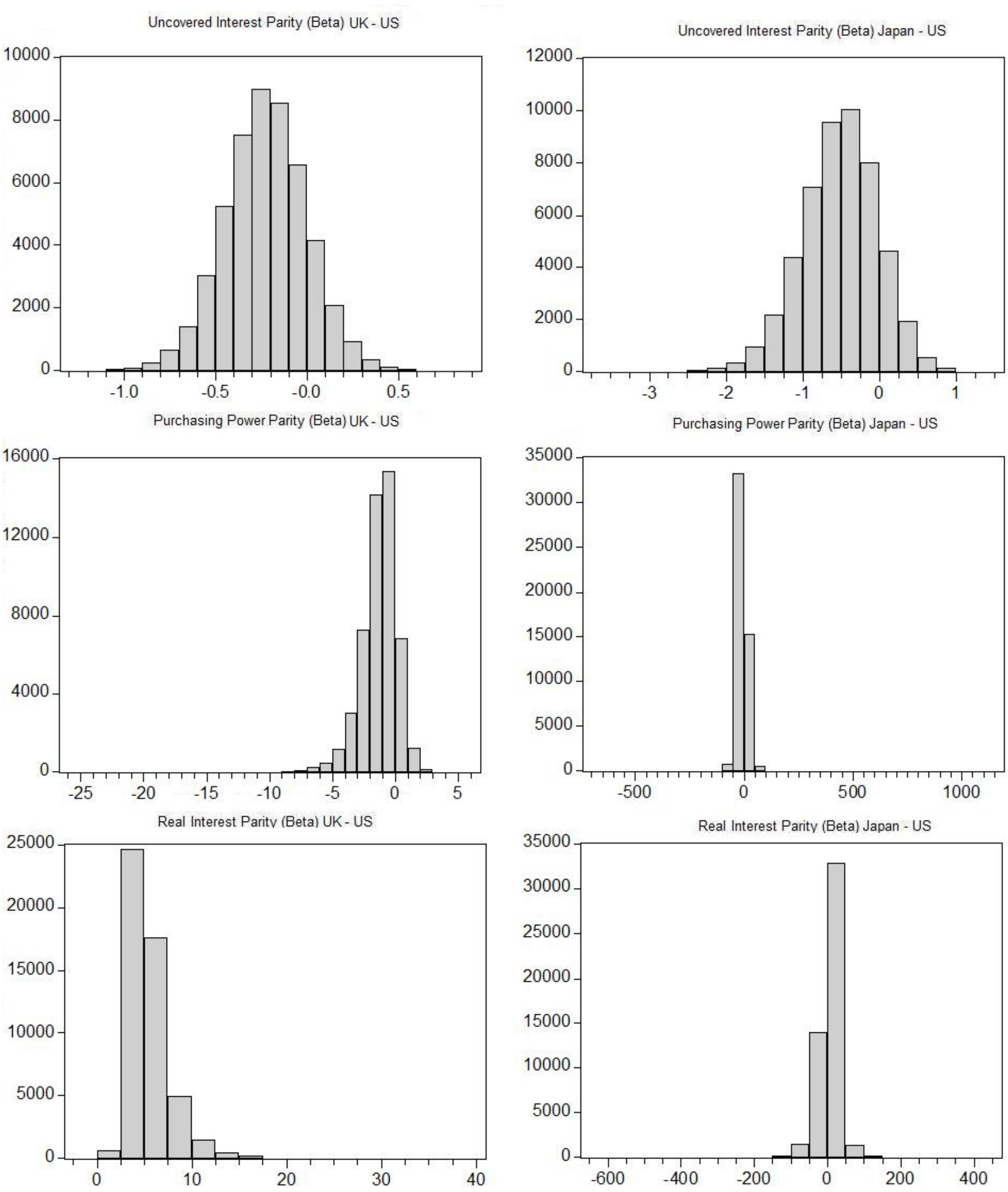


Figure 3: Histograms of the Simulated Beta Coefficients. Conditioning on Productivity Growth Differential

*Notes*: Uncovered interest parity (upper panel), relative purchasing power parity (mid panel) and real interest parity (lower panel, below). The distribution is based on 50.000 replications.

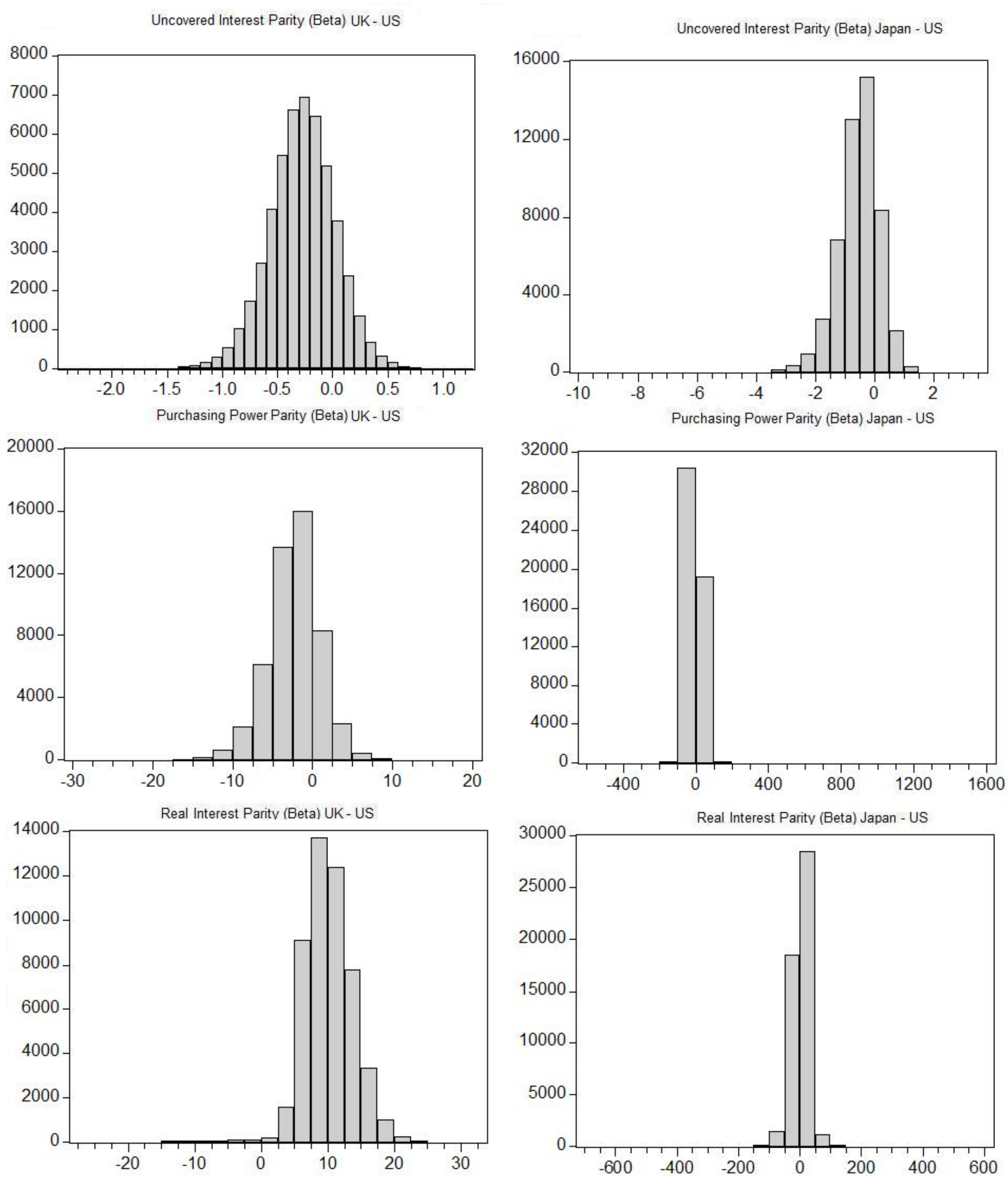


Figure 4: Histograms of the Simulated Beta Coefficients. Conditioning on Broad Money (M3) Growth Differential

*Notes*: Uncovered interest parity (upper panel), relative purchasing power parity (mid panel) and real interest parity (lower panel, below). The distribution is based on 50.000 replications.

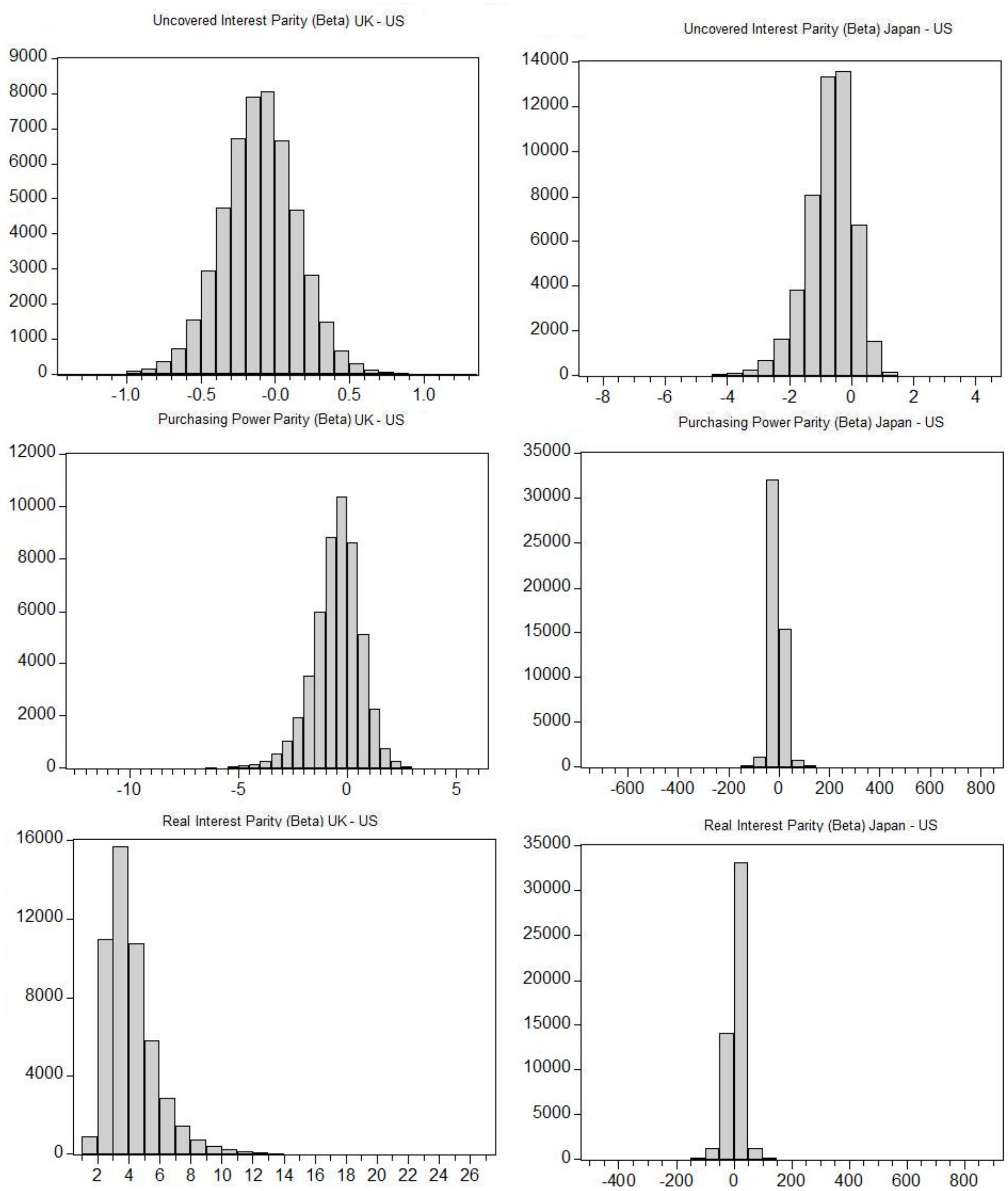


Figure 5: Histograms of the Simulated Beta Coefficients. Conditioning on Reserve Assets Growth Differential

*Notes*: Uncovered interest parity (upper panel), relative purchasing power parity (mid panel) and real interest parity (lower panel, below). The distribution is based on 50.000 replications.

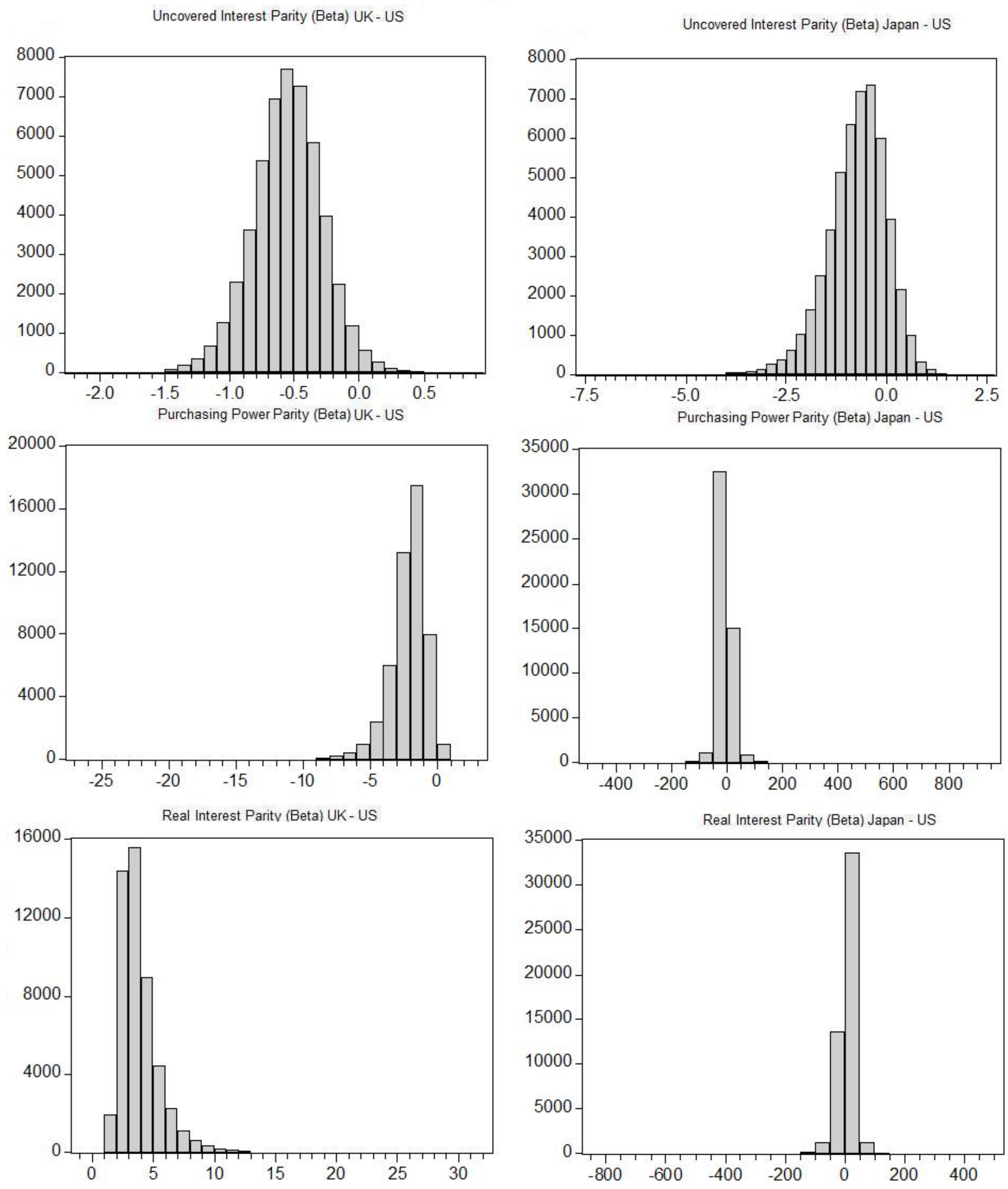


Figure 6: Histograms of the Simulated Beta Coefficients. Conditioning on Share Prices Return Differential

*Notes*: Uncovered interest parity (upper panel), relative purchasing power parity (mid panel) and real interest parity (lower panel, below). The distribution is based on 50.000 replications.