EUROPEAN CENTRAL BANK

EUROSYSTEM

WORKING PAPER SERIES

NO 1404 / DECEMBER 2011

A VAR ANALYSIS FOR THE UNCOVERED INTEREST PARITY AND THE EX-ANTE PURCHASING POWER PARITY

THE ROLE OF MACROECONOMIC AND FINANCIAL INFORMATION

by Corrado Macchiarelli

EUROPEAN CENTRAL BANK

EUROSYSTEM

WORKING PAPER SERIES

NO 1404 / DECEMBER 2011

A VAR ANALYSIS FOR THE UNCOVERED INTEREST PARITY AND THE EX-ANTE PURCHASING POWER PARITY

THE ROLE OF MACROECONOMIC AND FINANCIAL INFORMATION [1]

by Corrado Macchiarelli [2]

In 2011 all ECB publications feature a motif taken from the €100 banknote.

NOTE: This Working Paper should not be reported as representing the views of the European Central Bank (ECB). The views expressed are those of the author and do not necessarily reflect those of the ECB.

1 The author is grateful to Fabio Bagliano, Claudio Morana for constructive comments and discussions. He is also thankful to Andrea Beltratti, Carlo Favero, Agostino Consolo, Giulio Nicoletti and Philippe de Rougemont for further discussion. The paper also benefited from comments provided by participants at the Rimini Time Series Workshop, together with participants at the GdRe Symposium on Money, Banking and Finance and participants at an internal seminar organized by the Directorate General Economics of the European Central Bank. The author also acknowledges comments from an anonymous referee.

2 European Central Bank, Kaiserstrasse 29, D-60311, Frankfurt am Main, Germany; e-mail: corrado.macchiarelli@ecb.europa.eu

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Internet
http://www.ecb.europa.eu

Fax
+49 69 1344 6000

Information on all of the papers published in the ECB Working Paper Series can be found on the ECB's website, http://www.ecb.europa.eu/pub/scientific/wps/date/html/index.en.html

ISSN 1725-2806 (online)

CONTENTS

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 parities conditions at a long horizon. Particularly, test results are derived by rewriting the UIP, the EXPPP and the RIP as a set of cross-equation restrictions in the VAR (see also Campbell and Shiller, 1987; Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007; King and Kurmann, 2002). Consistent with the idea of some form of proportionality among the above three parities, we find a "forward premium" bias in both the UIP - as it is normally found in empirical analysis (e.g. Fama, 1987) - and the *ex ante* PPP. The latter result is new in the literature and stems from testing the PPP in *expectational* terms, thus assuming agents to bear on the uncertainty of future exchange rate changes and inflation dynamics. The overall results confirm the UIP to be currency-based (see also Bekaert *et al.*, 2007) and the EXPPP to be horizon-dependent (see also Lothian and Taylor, 1996; Taylor, 2002). Moreover, we find (weak) evidence that conditioning the VAR on variables having a strong forward-looking component (i.e. share prices) helps recover a unitary coefficient in the UIP equation.

JEL Classifications: E31, E43, E44, F31, C58.

Keywords: ex ante PPP, UIP, RIP, international parity conditions.

Non-Technical Summary

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. Despite the empirical support in favour of the two parities is generally mixed, recent analysis (Juselius, 1991; 1992; 1995; Johansen and Juselius, 1992; Pesaran *et al.*, 2000; Cheng, 1999; Throop, 1993; Zhou and Mahadavi, 1996; Hunter, 1991; Macchiarelli, 2011) have found evidence in favour of a PPP-UIP joint relation, in line with a goods vs. capital *general* equilibrium framework.
Following the same root, in this paper we jointly test the uncovered interest parity and the purchasing power parity, by introducing a third parity condition: the real interest parity (RIP). Here, the PPP is taken in *ex ante* terms (EXPPP) in order to test it over the same horizon as the UIP.
With respect to previous approaches, *inter alia* Pesaran *et al.* (2000), Juselius and MacDonald (2004), Macchiarelli (2011), in this paper the originality of the contribution can be gouged under two perspectives: first, a *present value* model (see also Campbell and Shiller, 1987; Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007; King and Kurmann, 2002) is employed in order to test the three parity conditions (UIP, EXPPP and RIP); secondly, the latter framework is augmented in order to account for macroeconomic and financial information, considered relevant for the formation of agents' expectations.
Focusing on UK and Japan vs. US data, our results support the idea of the UIP to be currency-dependent (Bekaert *et al.*, 2007) rather than horizon-based, whereas for the EXPPP the evidence goes in the opposite direction (see also Lothian and Taylor, 1996; Taylor, 2002). Moreover, the findings are consistent with the existence of some form of proportionality across the three parities; i.e. at the horizon considered (10-year maturities) the EXPPP is always found to hold, whilst UIP deviations are generally more likely for the UK vs. US data rather than for the Japan vs. US data, reconciling - in the former case - with observed RIP failures. Finally, the results of augmenting the framework with macroeconomic and financial variables invite further exploration, as the findings, albeit not supporting standard theoretical predictions, do support the view that, for accurate verification, the modeling of expectations is a central issue.
So formulated, the analysis in this article has two important implications. First, it shows that, at the long horizon, RIP deviations mainly stem from UIP (and not EXPPP) deviations. Such a finding has important implications for economists interested in international finance, and in particular, for those involved in jointly testing the above parity conditions, as it reinforces the view that, not only, proportionality do exist amongst the UIP, EXPPP and RIP, but also that nominal returns differentials explain real returns differentials alone, given that expected inflation differentials hold in purchasing

power parity. Secondly, for policy makers, it suggests that, as a way for pinning down UIP deviations, it can be important to consider variables embedding a strong forward looking component as a proxy for private sector expectations.

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). During that time, the faith in PPP essentially derived from the prevailing theory according to which price movements were dominated by monetary factors, given 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 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

[1]On the empirical ground, valid statistical results were achieved when the PPP started to be tested as a long run condition. Contributions such as Edison (1987), Lothian and Taylor (1996) and Taylor (2002) found the PPP to 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 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).

rational expectations assumption (see Section 2.2) being usually the explanation provided for such findings.[3] If, on the one side, the evidence in favour of the UIP at long horizons is recognized in the attempt of getting rid of short run exchange rate movements, 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 run UIP - rather than its short-term version - to be affected by market frictions, as it is very unlikely trading desks to keep capital in long-term contracts (Chaboud and Wright, 2005; Bekaert *et al.*, 2007, Burnside, 2011).

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; Macchiarelli, 2011) 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).

In order to test the PPP and the UIP jointly, we introduce a third parity condition: the real interest parity (RIP) (Cumby and Obstfeld, 1980; Mishkin, 1982; Jore *et al.*, 1993; Marston, 1997; Campbell *et al.*, 2007). Here, the PPP is taken in *ex ante* terms (EXPPP) in order to test it over the same horizon as the UIP. Particularly, the relation between the above three parities (UIP, EXPPP and RIP) is revisited by developing some joint coefficient tests obtained from a set of cross-equation restrictions in a VAR framework (Campbell and Shiller, 1987; Bekaert and Hodrick, 2001; and Bekaert *et al.*, 2007; King and Kurmann, 2002). The analysis focuses on the US dollar, the British sterling and the Japanese yen interest rates, exchange rates and changes in prices.

The results support the idea of the UIP to be currency-dependent (Bekaert *et al.*, 2007), whereas the EXPPP is found to be horizon-based (see also Lothian and Taylor, 1996; Taylor, 2002). The findings are moreover consistent with the existence of some form of proportionality across the three parities: at the horizon considered (10-year maturities) the EXPPP is always found to hold, whilst UIP deviations are generally more likely for the UK vs. US data than for the Japan vs. US data, reconciling - in the former case - with observed RIP failures. In light of the above, we also find the existence of a "forward premium" bias in both the UIP, as normally found in empirical studies, and

[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 the UIP would suggest. The "carry trade" consists in borrowing low-interest rate currencies and investing in high interest rate currencies, by exploiting such an anomaly (see Diez de los Rios and Sentana, 2007; Burnside, 2011).

the *ex ante* PPP; stemming the latter result from testing the purchasing power parity in *expectational* terms.

Finally, the baseline framework is augmented with a set of macroeconomic and financial variables, entering the VAR information set as exogenous. These latter results invite further exploration, as the findings, albeit not supporting standard theoretical predictions, do support the view that, for accurate verification, the modeling of expectations is a central issue.

So formulated, the analysis in this article has two important implications. First, it shows that, at the long horizon, RIP deviations mainly stem from UIP (and not EXPPP) deviations. Such a finding has important implications for economists interested in international finance, and in particular, for those involved in jointly testing the above parity conditions, as it reinforces the view that, not only, proportionality do exist amongst the UIP, EXPPP and RIP, but also that nominal returns differentials explain real returns differentials alone, given that expected inflation differentials hold in purchasing power parity. Secondly, for policy makers, it suggests that, as a way for pinning down UIP deviations, it can be important to consider variables embedding a strong forward looking component as a proxy for private sector expectations, while markets do not generally seem to react to variables such as cross-countries differences in industrial production (i.e. understood as a broad measure for the output gap) or differences in foreign reserve assets.

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

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

where $i_{t,l}$ represents the yield of a bond with maturity l at time t in the home country, and $f_{t,t+l}$ is the forward value of the (log) home vs. foreign spot nominal exchange rate, s, expiring l-periods ahead. The above expression is regardless of investors preferences (*unbiasedness* hypothesis).[4] Assuming individuals to be risk-adverse makes the forward rate to differ from the expected future spot rate,

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

$E_t s_{t+l}$, 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,

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

where $v_{t,t+l}$ is an *ex ante* risk premium. Substituting f into the CIP gives the standard uncovered interest parity condition,

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

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 zero 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:

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

under the null that $\beta^{uip} = 1$. In the regression, the time-varying premium (if any) enters the residual term ϵ_{t+l}^{uip}, i.e. $\epsilon_{t+l}^{uip} = \epsilon(v_{t+l})$.

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), asserting that for any good i:

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

where $p_t(i)$ and $p^*_t(i)$ describe the (log) current price for the good i in the home and in the foreign economy respectively, and s is the (log) 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.[5] If the LOP (at least theoretically) holds for every good i, the

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

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 (log) current price levels in both the foreign and the home country. However, many empirical tests do not compare identical basket of goods, but use different countries CPIs (consumer price indices) or WPIs (wholesale price indices).[6] 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 Δ is the difference operator. This relation predicts the relative inflation rate across countries to be necessarily compensated by changes in the nominal exchange rate.[7]

At this stage it should be noted that the PPP and the UIP are fundamentally different, being the former backward looking and the latter forward looking (Mishkin, 1982). In order to bring the PPP to the same horizon of the UIP (l periods), the former is reformulated in *ex ante* terms (EXPPP) (see also Macchiarelli, 2011), as

$$E_t \Delta s_{t+l} = E_t \Delta p_{t+l} - E_t \Delta p_{t+l}^* + o_{t+l}.$$

The above formulation is also augmented with a term, o_{t+l}, imposing a departure of the real exchange rate (RER) from the PPP equilibrium.[8] Iterating forward:

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

where the RER deviations term (o_{t+l}) is captured by the residual $\epsilon_{t+l}^{exppp} = \epsilon(o_{t+l})$. As before, if markets are efficients, equation (2) ensures commodity speculators to keep *expected* deviations from PPP in line under $\beta^{exppp} = 1$ (Roll, 1979).

[6] The PPP has indeed no reason to hold unless the two countries share identical consumption bundles. As underlined by Froot and Rogoff (1994), 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".

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

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

2.3 Uncovered Interest Parity and the Ex Ante Purchasing Power Parity

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

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

where real rates are defined according to the Fisher's (1907) parity condition, $r_{t+l} = i_{t,l} - E_t\Delta p_{t+l}$. So formulated, the RIP states that, in integrated financial markets, assets with identical liquidity and risk should command the same expected return regardless of market location.

According to Marston (1997), the real interest parity holds as soon as capital and goods markets are in equilibrium.[9] In other words, any couple in between the UIP, the EXPPP and the RIP should "naturally" imply the third relation. Adding and subtracting the term $E_t\Delta s_{t+l}$ in the above expression clarifies how the RIP becomes a relation conditional on the joint validity of the UIP and the EXPPP. Re-arranging and adding deviation terms we get in fact :

$$r_{t+l} - r^*_{t+l} + \xi_{t+l} =$$

$$= E_t\Delta s_{t+l} - (E_t\Delta p_{t+l} - E_t\Delta p^*_{t+l}) - o_{t+l} - \left[E_t\Delta s_{t+l} - (i_{t,l} - i^*_{t,l}) + v_{t+l}\right].$$

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 do exist, the RIP would necessarily allow for an erratic component ξ_{t+l}, which - by definition - should equal $(-v_{t+l} - o_{t+l})$.

Iterating forward, the RIP can be tested with the following regression:

$$\left(i_{t,l} - i^*_{t,l}\right) = \alpha^{rip}_l + \beta^{rip}_l \left[\frac{1}{l}\sum_{j=1}^{l}\left(E_t\Delta p_{t+j} - E_t\Delta p^*_{t+j}\right)\right] + \epsilon^{rip}_{t+l} \tag{3}$$

or

$$\overbrace{\frac{1}{l}\sum_{j=1}^{l} E_t\Delta s_{t+j} - \alpha^{exppp}_l - \beta^{exppp}_l \left[\frac{1}{l}\sum_{j=1}^{l}\left(E_t\Delta p_{t+j} - E_t\Delta p^*_{t+j}\right)\right] - \epsilon^{exppp}_{t+l}}^{EXPPP} = \tag{3'}$$

$$= \overbrace{\frac{1}{l}\sum_{j=1}^{l} E_t\Delta s_{t+j} - \alpha^{uip}_l - \beta^{uip}_l (i_{t,l} - i^*_{t,l}) - \epsilon^{uip}_{t+l}}^{UIP},$$

where, efficient markets simply imply the joint UIP-EXPPP restriction that $\beta^{rip} = 1$. The above formulation implies proportionality of the type $\beta^{rip} = \frac{\beta^{exppp}}{\beta^{uip}}$ among the tested coefficients, which

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

follows from the equality between equations (3) and (3').

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 introduce further uncertainty given *ex post* exchange rate and/or inflation forecast errors (see Marston, 1997).

In order to estimate equations (1), (2) and (3), we consider a 3-dimensional stationary VAR:

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

In the VAR i and i^*_t are 10 years constant maturity Treasury bond rates in the home and foreign economy respectively, s_t is the home vs. foreign spot nominal exchange rate (monthly average, denominated in US dollars) and Δp_t and Δp^*_t are the *cpi*-inflation levels in both the home and foreign economy. Further in the paper we consider some exogenous 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 exogenous macroeconomic/financial variables are transformed in logarithmic month-on-month changes. The only variables which are not transformed are interest rates. 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. The sample covers the period from 1975-1 to 2008-6. The series for the long term interest rate for Japan starts in 1989-1.

Table 1 reports some descriptive statistics for the variables. Over the whole 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 Δp themselves (see Bekaert *et al.*, 2007). For sake of completeness, in the last three columns of Table 1 some descriptive statistics for RIP, PPP and UIP deviations are also reported. Those represent changes in the log real exchange rate, the real interest rate differential and the UIP premium, obtained as the difference between the exchange rate changes and the interest rate spread, the exchange rate changes and the inflation spread, and the interest rate spread and the inflation spread, respectively. The results point out that such deviations are negative in the case of the UIP and the PPP, while positive for the RIP. Nonetheless deviations are derived under the assumption of a one-to-one adjustment between, e.g., exchange rate changes and the interest rate spread, thus leading to deviation terms which are generally lower than the one obtained with standard regression results. In addition, the presence of expectational errors may clearly inflate second moments.

Following a standard approach, 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 system and $K = 1$ for the Japan vs. US system (Table 3). Using a VAR framework inevitably introduces estimation error, possibly worsened by the overlapping observation problem which induces moving average errors. When estimating the VAR, we use heteroskedasticity consistent standard errors and also correct for MA terms up to the $l-1$ order, using a Newey-West window as in Chinn and Meredith (2004; 2005).[10]

For each system we then reformulate y_t in the standard companion form, defining $z_t^{'} = (y_t, y_{t-1}, ..., y_{t+1-K})$. 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 Σ.

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. $E(z_{t+j}|z_t) = A^j z_t$, consistent with the idea of *present value* models (see Campbell and Shiller, 1987; King and Kurmann, 2002).[11]

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 (see also Bekaert and Hodrick, 2001; Bekaert *et al.*, 2007):

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

[10] We avoid using the Hansen-Hodrick (1980) estimator as this has the tendency to produce non-positive-definite variance-covariance matrices.

[11] 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).

which, using *geometric series* results, for the 120-months horizon is:[12]

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

with $C = A(I - A^{120})(I - A)^{-1}$. Using a *present value* model assumes that the interest rate spread is a linear function of the present discounted value of expected future exchange rate changes, where the discount factor equals A. In this way, expectations of future exchange rate changes are conditional on the full public information set, I_t, which includes Δs and $i - i^*$ themselves and generally exceeds the information set available to the econometrician, J_t (see Campbell and Shiller, 1987). The usual problem of deriving restrictions in a VAR framework, thus assuming that agents use the same information set as the VAR, does not apply here. In fact, for the motivations outlined above, also a simple threevariate VAR can well capture the dynamics of interest (*ibid.*).[13]

Under the same reasoning, the EXPPP imposes the restrictions:

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

and so too does the real interest parity:

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

3.2 Implied VAR Statistics

The set of restrictions (5)-(7) allows the estimation of the slope coefficients implied from the VAR 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 Ψ is the unconditional variance of z_t, computed as $vec(\Psi) = (I - A \otimes A)^{-1}vec(\Sigma)$. The numerator in equation (8) is the covariance between the expected future rate of appreciation and the interest rate differential, whereas the denominator represents the variance of the interest rate spread. Analogously, for the EXPPP the implied slope coefficients for the 120-month horizon is:[14]

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

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

[13] In the next Session we nonetheless assess the validity of our findings by conditioning the VAR on a set of exogenous macroeconomic and financial variables.

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

and similarly for the RIP:

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

To further characterize UIP-EXPPP-RIP deviations we compute three distinct statistics for each condition, following Bekaert *et al.* (2007). The tests are performed following the same set of restrictions as in equations (5)-(7), with C and Ψ fully capturing exchange rates changes, interest rates and inflation spread dynamics in the VAR.

Under the UIP, the expected exchange rate change should be perfectly correlated with the interest rate differential, and they are expected to display equal variability. Hence, a correlation (CORR) and variance ratio (VR) statistics are computed for the UIP as

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

and

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

The standard deviation (SD) of the residual from the UIP equation is also measured as

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

where the residual is derived from the UIP under the null, i.e. $\epsilon_{t+l}^{uip} = \frac{1}{l}\left(E_t \sum_{j=1}^{l} \Delta s_{t+j}\right) - (i_{t,l} - i_{t,l}^*)$. In a similar fashion, correlation, variance ratio and standard deviation statistics are computed for the EXPPP and the RIP conditions.

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).[15] This might turn to be a crucial point when discriminating across round proximate test results.

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

In order to bias-correct VAR coefficients,[16] 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 starting values, 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 by 1000 data points. Those initial conditions are then discarded when the estimation is performed. 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 correction is performed by adding back the biases to the original VAR estimates. This set of bias-corrected parameters is then used to construct the point estimates for the betas and the statistics described in Section 3.2, representing the *alternative* of violation of the parity conditions hypotheses. Importantly, bias correction and the computation of the VAR implied coefficients and statistics are performed at two different stages; i.e. bias correction and the computation of the betas as non-linear functions of the VAR coefficients do not interfere with each others.

In order to obtain the empirical distribution of the aforementioned VAR coefficients and statistics, the Montecarlo procedure is repeated so to simulate again 50.000 data samples from the original VAR residuals. At each bootstrap draw, bias correction is implemented on the parameters of interest (where biases are taken from the previous step), and new implied VAR coefficients and statistics are obtained. Relevant quantiles at the 2.5% and 97.2% 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.[17] 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 estimates for the betas 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.[18] If this is not surprising in the context of the UIP - given the existence of the well documented "forward premium" puzzle - it is surprising under the EXPPP hypothesis. Nonetheless,

[16] Bias correction can increase mean square error in the case of a *purely* non-stationary VAR (see Abadir *at al.*, 1999) which is not the case here.

[17] 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 (upwardly biased in the UK - US system and downwardly biased in the Japan - US system).

[18] 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.

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. Also in light of the estimated signs, the beta coefficients are consistent with the existence of proportionality of the type $\beta^{rip} = \frac{\beta^{exppp}}{\beta^{uip}}$ among the three conditions (see Section 2).

The dimension of the correlation coefficients gives further insight on the validity of the UIP, the EXPPP and the RIP hypotheses. As the statistics inherits its sign from the implied slope coefficient, in both the UK vs. US and the Japan vs. US systems correlation among the numerator and the denominator in equations (8), UIP, and (9), EXPPP, is negative, while it is positive in the RIP, equation (10). Interestingly, the correlation statistics is broadly the same among the three parities (about 0.9 in absolute value). For the Japan vs. US system this is consistent with the idea that, for the RIP to hold, the UIP and the EXPPP correlation statistics should be sensitively close to 1 (on the purchasing power parity see also Gokey, 1994). Alternatively, for the UK. vs. US system, one might think at more substantial deviations to occur in the UIP case rather than in the EXPPP case.[19]

In both systems, 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. Such finding also suggests that the variance of the interest rate spread is generally higher than the variance of the expected nominal exchange rate changes. 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 compared to the expected exchange rate changes and the interest rate spread; thus overall confirming goods prices to be less volatile than interest rates.

The standard deviation of the errors (SD), capturing the variability of the residuals / deviation terms in each equation (see Section 2.1), is in all cases close (or higher) than unity, being consistent with a time-varying *risk premium* explanation which is, on average higher, for the UIP hypothesis. In this respect, higher deviations in the case of the UIP, compared to the EXPPP, corroborate the idea that UIP departures 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, we reject the null of beta equal to one for the UIP with p-value (0.001).[20]

---

[19] In fact CORR is high but, as explained later, the UIP and the RIP still do not hold at 1%, possibly reconciling with a *risk premium* explanation.

[20] In light of the distribution for beta UIP, a two sided test is used here.

This is in line with the findings in Bekaert *et al.* (2007) against the uncovered interest parity using the USD vs. British sterling data at different horizons (3 and 60 months respectively; see also Macchiarelli, 2011). The probability of having a unitary coefficient in the EXPPP is not-rejected instead with p-value (0.570). For the RIP, the test moreover confirms the rejection of $\beta^{rip} = 1$ with p-value (0.000).[21] For the Japan - US system the results report a decisive non-rejection of the three parities at the horizon considered.

Overall, the findings 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 the UIP, the EXPPP and the RIP should "naturally" imply the third relation (see also Marston, 1997). For the EXPPP and the RIP, the estimated confidence intervals (2.5% and 97.5% fractiles) are nonetheless quite large - as price inflation is projected in the VAR - 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), in this Section we assess the role of expectations mispecifications by augmenting the information set to include macroeconomic and financial variables, some expected to feature forward looking properties. Such an extension also assesses the robustness of our findings. We condition the VAR on a set of exogenous regressors (h_t), including: (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" spreads. For each currency pair, the VAR is conditional on one exogenous regressor at a time, where, for sake of simplicity, regressors are let to affect the VAR only contemporaneously.

The inclusion of variables besides the one predicted by standard economic theories 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). In this setting, exogenous regressors are primarily aimed at capturing variables signaling cross-country macroeconomic developments and international imbalances agents may use when formulating their expectations. For instance, the industrial production growth differential can be seen as a broad measure for the output gap, whereas share prices may proxy agents' perceptions about future cyclical economic developments.

In order to solve the VAR forward, we write the VAR augmented with exogenous regressors (VAR-

[21] A one-sided test is considered here, as the distribution is strongly skewed.

X) in companion form and partition exogenous regressors as $z_t = Az_{t-1} + Bh_t + e_t$. As before, A is $3K$-dimensional companion matrix, while B is a $3K$ vector which is non-zero only in the first 3 rows. Taking expectations over $E\left(z_{t+j} - B\sum_{q=0}^{j-1} A^q h_{t+j-q}|z_t\right) = A^j z_t$ allows the usual formulation to apply (Section 3.3). The results are reported 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 the RIP hypothesis is found only for the UK - US system, in what follows we mainly focus on the results for this latter pair of countries. Overall, however, the results are not sensitively affected by the inclusion of exogenous regressors, as a "forward premium" bias persists in all cases. In particular, considering productivity growth differentials and foreign reserve assets for the UK vs. US, the p-value for non-rejections of the null $\beta^{uip} = 1$ and $\beta^{rip} = 1$ (in absolute value) does not increase above 1% (see Table 5 and 7). Broad money (M3) growth differential helps instead reduce the coefficient bias in the RIP equation for the UK - US system (skewness is reduced), albeit there is not clear evidence of non-rejection of $\beta^{rip} = 1$ at a conventional significance level (see Table 7). Finally, conditioning on share price return differentials helps center the distribution for the UIP in the UK - US system over a mean value of about -0.6, yet allowing not to reject the null of $\beta^{uip} = -1$ at least at the 2.5% level. Albeit weak, 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. This somewhat reconciles with theories pointing to the importance of foreign exchange rate *premia* in explaining UIP deviations (e.g. Fama, 1987).

5 Conclusions

In this paper, we revisited the relation between the uncovered interest parity (UIP), the *ex ante* purchasing power parity (EXPPP) and the real interest parity (RIP) using a VAR framework for the UK vs. US and the Japan vs. US data. The evidence is based on developing some joint coefficient-based tests obtained by rewriting the above relations as a set of cross-equation restrictions in the VAR (Campbell and Shiller, 1987; Bekaert and Hodrick, 2001; Bekaert *et al.*, 2007; King and Kurmann, 2002).

The results point to the existence of a "forward premium" bias in both the UIP and the EXPPP equations. A "forward premium puzzle" in the EXPPP case is new in the literature and stems from testing the PPP in *expectational* terms, thus assuming agents to bear on the uncertainty of future exchange rate changes and prices dynamics.

The overall results are 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). In addition, the failure of the 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 the UIP, the EXPPP and the RIP should necessarily imply the third relation (Marston, 1997). Said that, 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.

To better characterize our results, we finally augmented the baseline VAR framework with exogenous regressors. We find the results not to be sensitively affected by the inclusion of exogenous variables, as a "forward premium" bias persists in all cases. Nonetheless, conditioning on share prices return differentials yields a better fitting of the UIP relation in the UK - US system, thus providing support to the role of foreign exchange rate *premia* in explaining UIP deviations (e.g. Fama, 1987). This has important implications for policy making as it suggests that variables proxying private sector expectations may play some role in explaining UIP misalignments.

Overall, these results invite further exploration, as albeit standard theoretical predictions are not corroborated, we support the view that, for accurate verification, the modeling of expectations is a central issue. All in all, 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.

References

[1] Abadir K. M. *et al.* (1999), The Influence of VAR Dimensions on Estimator Biases, *Econometrica*, **67**(1), 163-181.

[2] Bekaert G. *et al.* (1997), On biases in tests of the expectations hypothesis of the term structure of interest rates, *Journal of Financial Economics*, **44**, 309-348.

[3] Bekaert G. *et al.* (2007), Uncovered Interest Rate Parity and the Term Structure, *Journal of International Money and Finance*, **26**, 1038-1069.

[4] Bekaert, G., R.J. Hodrick (2001), Expectations Hypotheses Tests, *Journal of Finance*, **56**, 1357-1394.

[5] Burnside C. (2011), Carry trades and risks, NBER Working Paper 17278.

[6] Campbell J.Y., Shiller (1987), Cointegration and test of present value models, *Journal of Political Economy*, **95**, 1062-88.

[7] Campbell R.A.J. *et al.* (2007), Irving Fisher, Expectational Errors and the UIP Puzzle, CEPR Discussion Paper, No. 6294.

[8] Chaboud A.P., Wright J.H. (2005), Uncovered Interest Parity: It works but not for long, *Journal of International Economics*, **66**(2), 349-362.

[9] Cheng B.S. (1999), Beyond the purchasing power parity: testing for cointegration and causality between exchange rates, prices and interest rates, *Journal of International Money and Finance*, **18**, 911-924.

[10] Cheung Y-W. (1993), Exchange rate premiums, *Journal of International Money and Finance*, **12**, 182-194.

[11] Cheung Y-W., Lai K.S. (1993), Long run purchasing power parity during the recent float, *Journal of International Economics*, **34**, 181-192.

[12] Chinn M.D., Meredith G. (2004), Monetary policy and long horizon uncovered interest parity, IMF Staff Papers, **51**(3), 409-430.

[13] Chinn M.D., Meredith G. (2005), Testing Uncovered Interest Parity at Short and Long Horizons during the Post-Bretton Woods Era, NBER Working Paper, 1107.

[14] Cumby R.E., Obstfeld M. (1980), Exchange-Rate Expectations and Nominal Interest Differentials: A Test of the Fisher Hypothesis, NBER Working Paper Series, No. 537.

[15] Diez de los Rios A., Santena E. (2007), Testing Uncovered Interest Parity: A Continuous Time Approach, CEMFI Working Paper No. 0714.

[16] Dornbusch R. (1976), Expectations and Exchange Rate Dynamics, *The Journal of Political Economy*, **84**(6), 1161-1176.

[17] Edison H.J. (1978), Purchasing Power Parity in the Long Run: A Test of the Dollar/Pound Exchange Rate (1890-1978), *Journal of Money, Credit and Banking*, **19**(3), 376-387.

[18] Evans M.D.D. (1998), Expected Inflation and Inflation Risk Premia, *The Journal of Finance*, **53**(1), 187-218.

[19] Fama E.F. (1984), Forward and Spot Exchange Rates, *Journal of Monetary Economics*, **14**, 319-338.

[20] Fisher I. (1907), The Rate of Interest, New York, MacMillan, 278 and ff. in Campbell et al (2007).

[21] Fraga A. (1986), Price Uncertainty and the Exchange Rate Risk Premium, *Journal of International Economics*, 20, 179-185.

[22] Froot A.K., Rogoff K. (1994), Perspectives on PPP and Long-Run Real Exchange Rates, NBER Working Paper, No. 4952.

[23] Gokey T.C. (1994), What explains the risk premium in foreign exchange returns?, *Journal of International Money and Finance*, **13**(6), 729-738.

[24] Green P. (1992), Is Currency Trading Profitable? Exploiting Deviations from Uncovered Interest Parity, *Financial Analyst Journal*, **48**(4), 82-86.

[25] Hai W. *et al.* (1997), Understanding Spot and Forward Exchange Rate Regressions, *Journal of Applied Econometrics*, **12**, 715-734.

[26] Hunter J. (1991), Test of Cointegration Exogeneity for PPP and Uncovered Interest Parity in the United Kingdom, *Journal of Policy Modeling*, **14**(4), 453-463.

[27] Johansen S., Juselius K. (1992), Testing Structural Hypothesis in a Multivariate Cointegration Analysis of the PPP and the UIP for UK, *Journal of Econometrics*, **53**, 211-244.

[28] Jore A.S. *et al.* (1993), Testing for Purchasing Power Parity and Interest Rate Parity on Norwegian Data, Central Bureau of Statistics Working Paper, Oslo.

[29] Juselius K. (1991), Long Run Relations in a Well Defined Statistical Model for the Data Generating Process. Cointegration Analysis of the PPP and UIP Relations, in Gruber J. (edited by),

Econometric Decision Models: New Methods of Modeling and Applications, Springer Verlag, 336-357.

[30] Juselius K. (1992), On the Empirical Verification of the Purchasing Power Parity and the Uncovered Interest Parity, Nationaløkonomisk Tidskirift, No. 130, 57-66.

[31] Juselius K. (1995), Do Purchasing Power Parity and Uncovered Interest Parity hold in the long run? An example of likelihood inference in a multivariate time-series model, *Journal of Econometrics*, **69**, 211-240.

[32] Juselius, K., MacDonald, R. (2004), International Parity Relationships Between the USA and Japan, *Japan and the World Economy*, **16**, 17-34.

[33] Kendall M. G. (1954), Note on the bias in the estimation of autocorrelation, *Biometrika*, **41**, 403-404.

[34] King R. G. and Kurmann A. (2002), Expectations and the term structure of interest rates: evidence and implications, *Federal Reserve Bank of Richmond Economic Quarterly*, **88**(4), 49-95.

[35] Lothian J.R., Taylor M.P. (1996), Real Exchange Rate Behaviour: The Recent Float from the Perspective of the Past Two Centuries, *The Journal of Political Economy*, **103**(3), 488-509.

[36] MacDonald R. (1993), Long-Run Purchasing Power Parity: Is it for Real?, *The Review of Economics and Statistics*, **75**(4), 690-695.

[37] MacDonald R. (1998), What Do We Really Know About Real Exchange Rates?, OeNB Working Paper, No. 28.

[38] MacDonald R., Nagayasu J. (1999), The Long Run Relationship Between Real Exchange Rates and Real Interest Rate Differentials: A Panel Study, IMF Working Paper, No. 99/37.

[39] MacDonald R., Taylor M.P. (1991), Exchange Rate Economics: A Survey, IMF Working Paper, No. 91/62.

[40] Macchiarelli C. (2011), Bond Market Co-Movements, Expected Inflation and the Equilibrium Real Exchange Rate, ECB Working Paper 1405.

[41] Mariott F.H.C., Pope J.A. (1954), Bias in the estimation of autocorrelations, *Biometrika*, **41**(3-4), 390-402.

[42] Mark N.C., Wu Y. (1998), Rethinking Deviations from Uncovered Interest Parity: The Role of Covariance Risk and Noise, *The Economic Journal*, **108**, 1686-1706.

[43] Marston R. (1997), Test of three parity conditions: distinguishing risk premia and systematic forecast errors, *Journal of International Money and Finance*, **16**(2), 285-303.

[44] Mishkin F.S. (1982), Are Real Interest Rates equal across Countries? An empirical Investigation of International Parity Conditions, NBER Working Paper, No. 1048.

[45] Pesaran M.H. *et al.* (2000), Structural Analysis of vector error correction models with exogenous I(1) variables, *Journal of Econometrics*, **97**, 293-343.

[46] Rogoff K. (1996), The Purchasing Power Parity Puzzle, *Journal of Economic Literature*, **2**(34), 647-668.

[47] Sarno L., Taylor M.P. (2002), Purchasing Power Parity and the Real Exchange Rate, *IMF Staff Papers*, **1**(49), 65-105.

[48] Roll R. (1979), Violations of purchasing power parity and their implications for efficient international commodity markets, in *International Finance and Trade*, Eds. M. Sarnat and G.P. Szego, pp. 133-176, Ballinger Publishing, Cambridge, MA in Marston R. (1997), Test of three parity conditions: distinguishing risk premia and systematic forecast errors, *Journal of International Money and Finance*, **16**(2), 285-303.

[49] Taylor A.M. (2002), A Century of Purchasing-Power Parity, *The Review of Economics and Statistics*, **84**(1), 139-150.

[50] Taylor M.P. *et al.* (2001), Nonlinear Mean-Reversion in Real Exchange Rates: Toward a Solution to Purchasing Power Parity Puzzles, *International Economic Review*, **42**(4), 1015-1042.

[51] Throop A.W, (1993), A Generalized Uncovered Interest Parity Model of Exchange Rates, Economic Review, Federal Reserve Bank of San Francisco, No. 2, 3-16.

[52] Zhou S., Mahdavi S. (1996), Simple vs. Generalized Interest Rate and Purchasing Power Parity Models of Exchange Rates, *The Quarterly Review of Economics and Finance*, **36**(2), 197-218.

Table 1: Summary and Descriptive Statistics

	Δs_t	$i_t - i_t$	$\Delta p_t - \Delta p_t^*$	RIP	PPP	UIP
UK - US System						
Mean	0.561	1.396	1.060	0.321	-0.499	-0.820
Variance of sample mean	1.458	0.089	0.361	0.343	1.499	1.464
Maximum	132.185	8.300	42.386	17.542	133.447	130.126
Minimum	-120.312	-2.0300	-17.852	-37.336	-140.274	-119.953
AC(1)	0.330	0.973	0.252	0.165	0.301	0.335
AC(2)	-0.018	0.930	0.096	-0.011	-0.042	-0.012
AC(3)	0.040	0.896	0.043	-0.061	0.019	0.045
AC(4)	0.055	0.876	0.062	-0.038	0.027	0.059
Japan - US System						
Mean	-3.086	-2.959	-2.374	-0.569	-0.712	2.215
Variance of sample mean	1.665	0.058	0.325	0.331	1.721	2.123
Maximum	95.588	-0.858	25.024	13.337	97.537	98.863
Minimum	-123.181	-4.969	-24.844	-27.727	-128.346	-119.530
AC(1)	0.297	0.965	0.081	0.147	0.297	0.272
AC(2)	0.040	0.919	-0.200	-0.186	0.033	0.081
AC(3)	0.060	0.885	-0.094	-0.214	0.053	0.047
AC(4)	0.033	0.857	0.016	-0.040	0.028	-0.067

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. The last three columns represent RIP, PPP and UIP deviations, computed as the difference between the month-on-month changes appreciation rate and the interest rate spread (UIP) and month-on-month inflation spread (PPP). The RIP is analogously obtained as the difference between the interest rate spread and the inflation spread. In all cases differences are considered as home vs. foreign, i.e. UK and Japan vs. US. 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.

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

	Δs_t	$i_t - i_t^*$	$\Delta p_t - \Delta p_t^*$
UK - US System			
Δs_{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)
Δs_{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)
Δs_{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			
Δs_{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	
	β^{uip}	β^{exppp}	β^{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 statistics, 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		
	β^{uip}	β^{exppp}	β^{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 statistics, 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 on Broad Money (M3) Growth Differential

	Coefficients			Additional Statistics								
					UIP			PPP			RIP	
	β^{uip}	β^{exppp}	β^{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 statistics, 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	
	β^{uip}	β^{exppp}	β^{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 statistics, 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	
	β^{uip}	β^{exppp}	β^{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 statistics, 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.

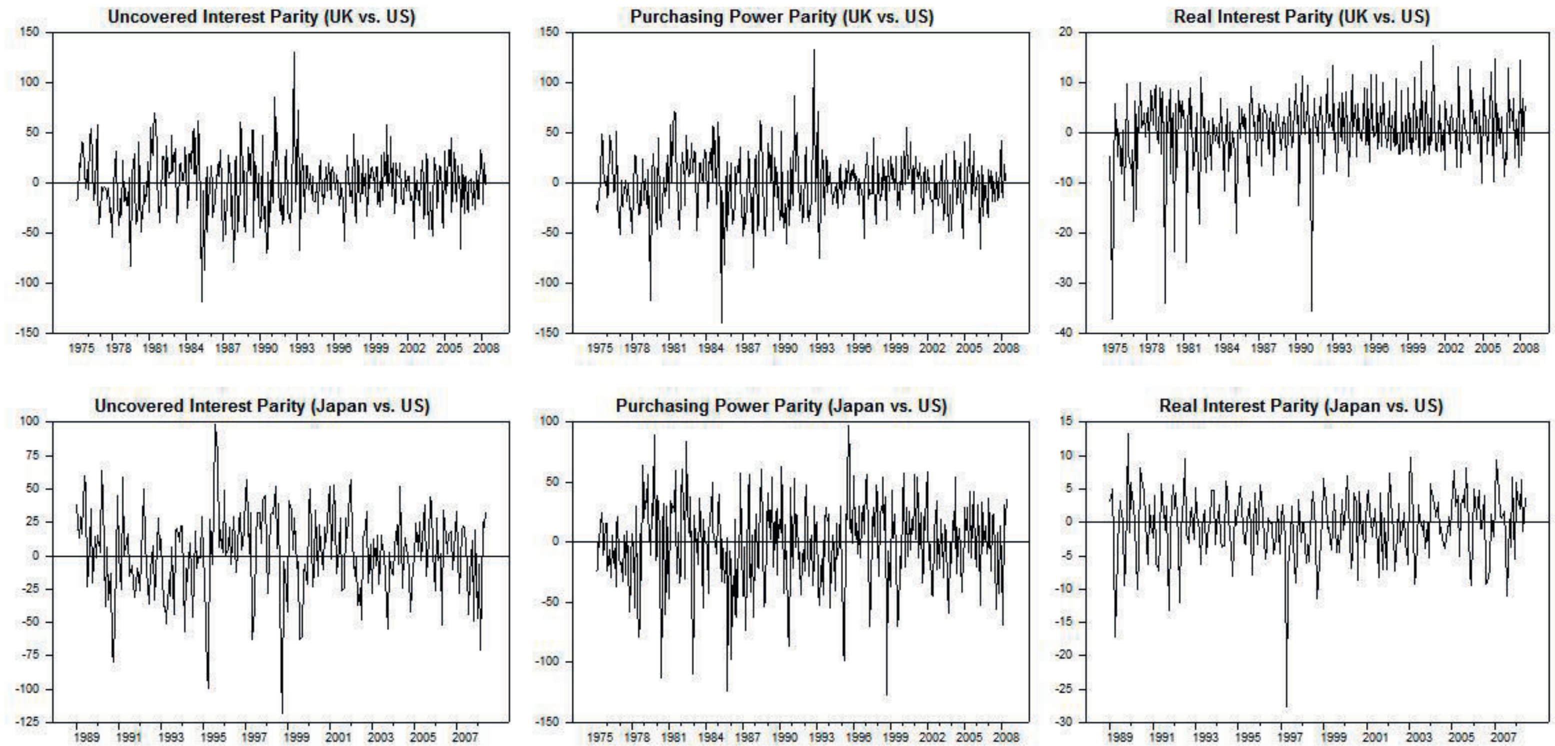

Figure 1: Deviations from International Parity Conditions

Notes: Uncovered interest parity (left panel), relative purchasing power parity (middle panel) and real interest parity (right panel). For each currency pairs, the chart represents RIP, PPP and UIP deviations, computed as the difference between the month-on-month changes appreciation rate and the interest rate spread (UIP) and month-on-month inflation spread (PPP). The RIP is analogously obtained as the difference between the interest rate spread and the inflation spread. In all cases differences are considered as home vs. foreign, i.e. UK and Japan vs. US. 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.

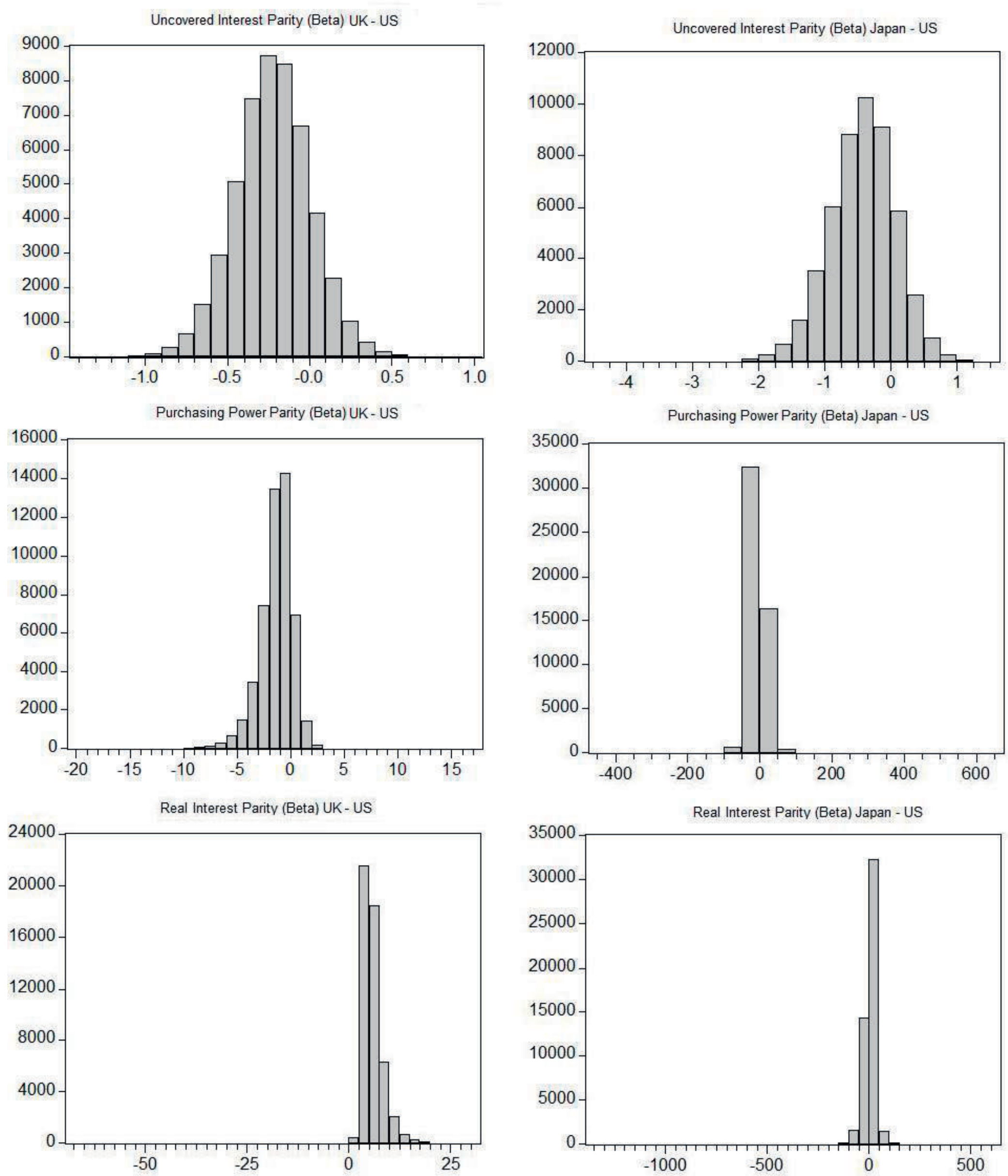

Figure 2: Histograms of the Simulated Beta Coefficients

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.

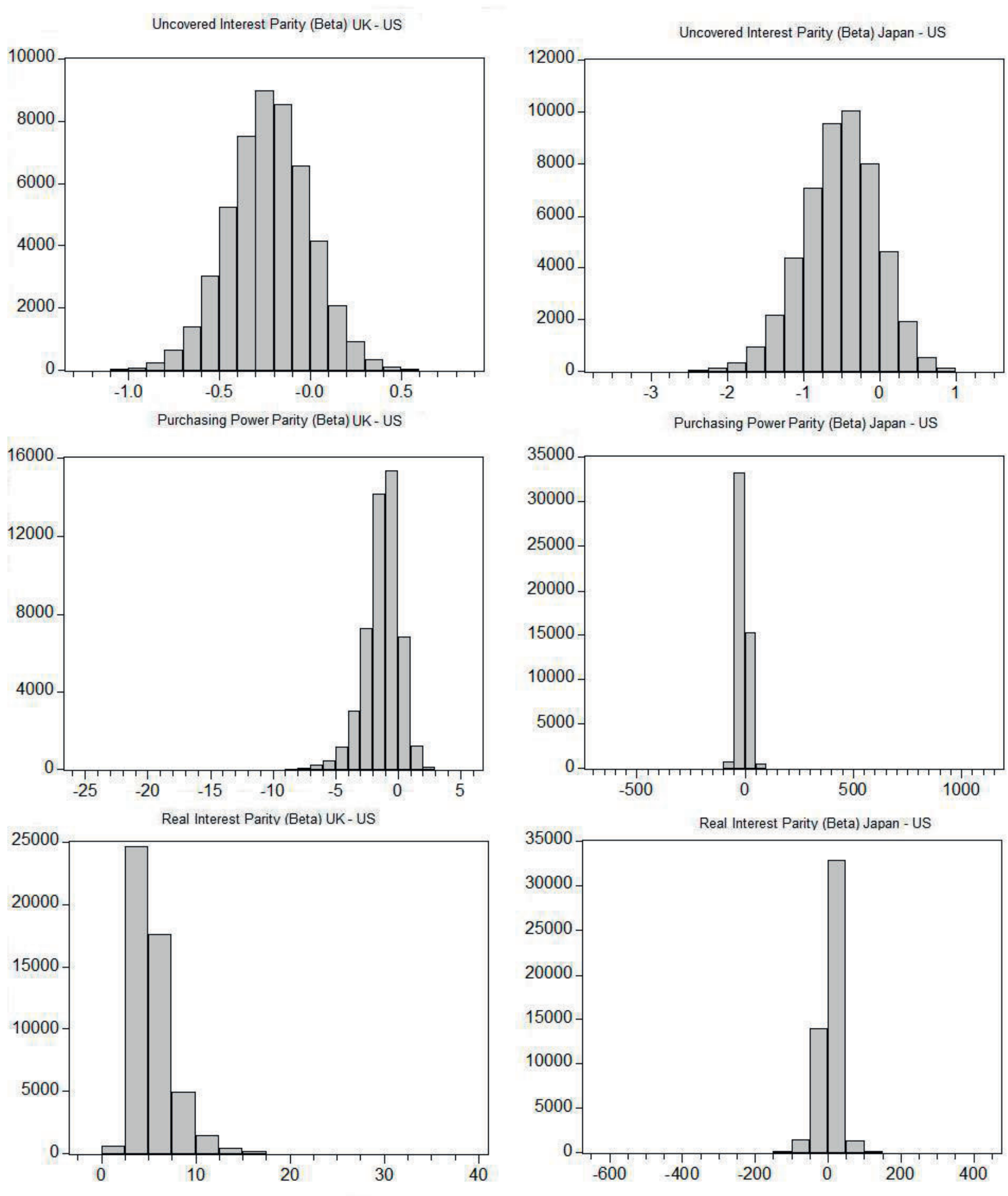

Figure 3: Histograms of the Simulated Beta Coefficients. Conditioning on Productivity Growth Differential

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.

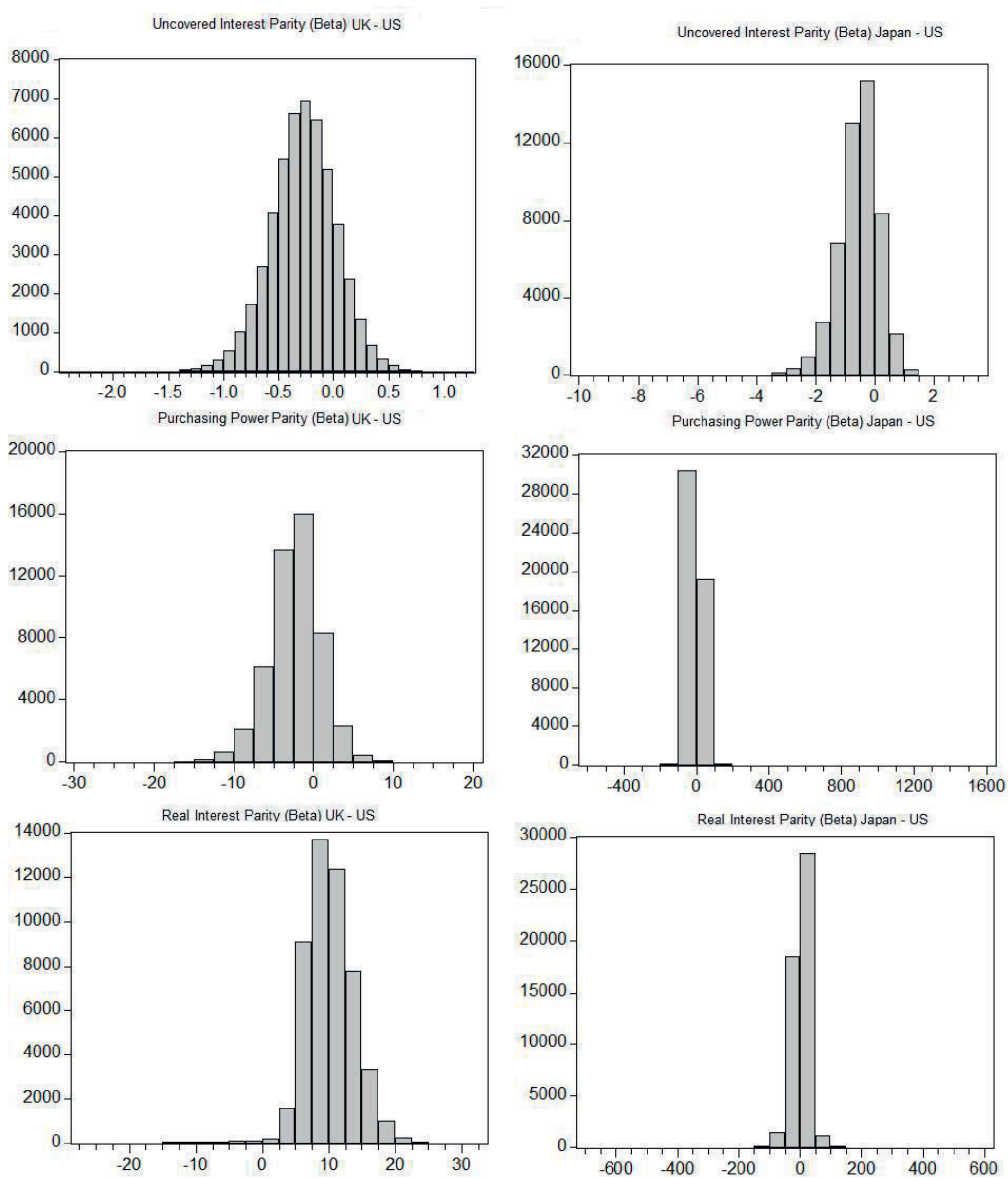

Figure 4: Histograms of the Simulated Beta Coefficients. Conditioning on Broad Money (M3) Growth Differential

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.

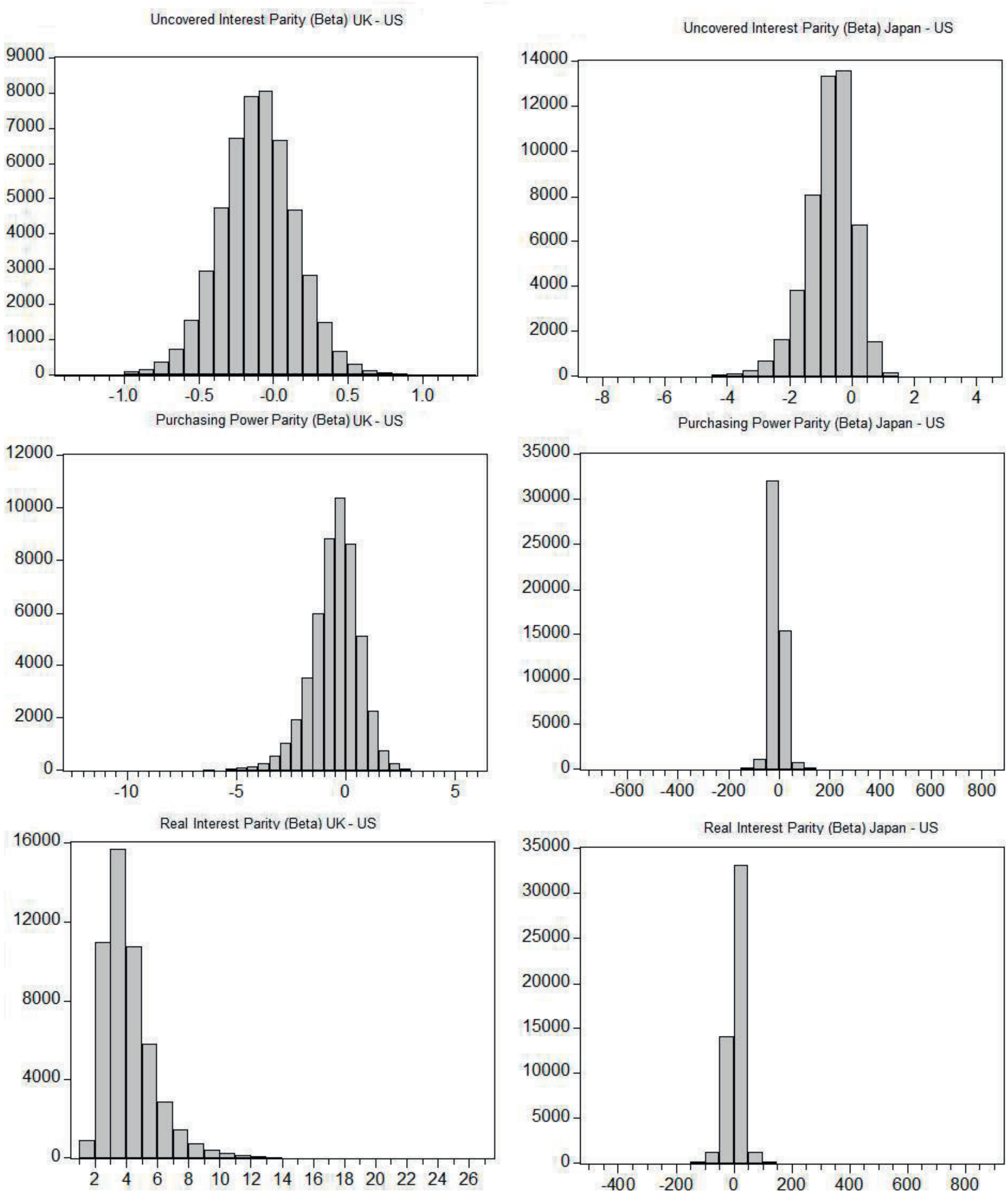

Figure 5: Histograms of the Simulated Beta Coefficients. Conditioning on Reserve Assets Growth Differential

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.

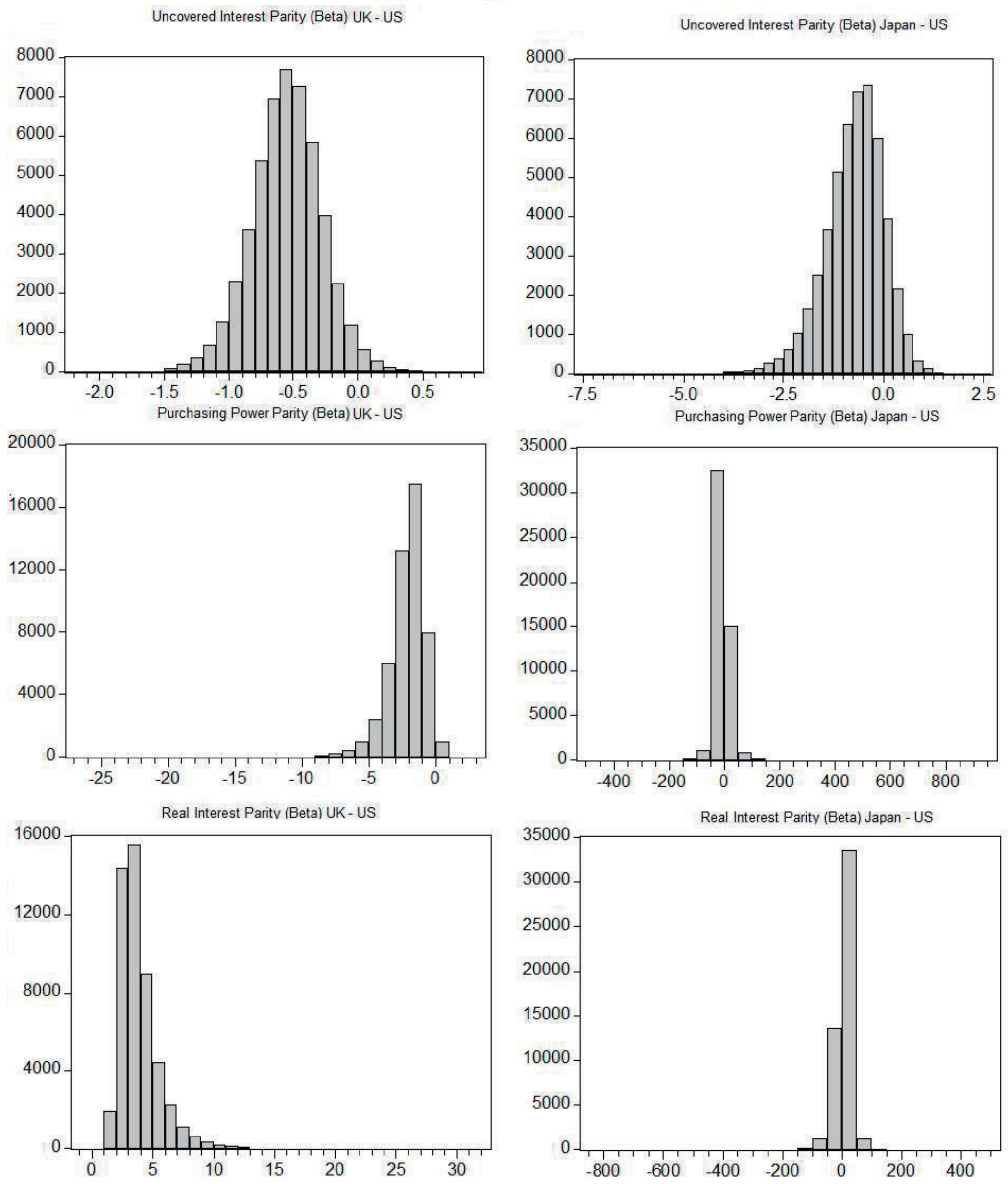

Figure 6: Histograms of the Simulated Beta Coefficients. Conditioning on Share Prices Return Differential

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.