

Another look at long-horizon uncovered interest parity

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Abstract

Long-horizon uncovered interest parity during the post-Bretton Woods era in the G7 countries is analyzed in this paper. The main difference with previous studies relies in the use of cointegration methods due to the non-stationary behavior of the variables involved in the sample. The consideration of structural breaks becomes a key element for this relationship to hold, which are related to monetary policy changes, the exchange rate evolution and the existence of a time-varying risk premium on long-term bonds. It is also addressed the robustness of the obtained results to recent developments in the Eurozone.

Keywords: Long-horizon regressions, Uncovered Interest Parity, Cointegration, Structural Breaks, Risk Premium.

JEL: C1, E4, E5, F3.

1 Introduction

Uncovered interest parity (UIP) is a cornerstone of modern models of international finance and exchange rate determination. It establishes a simple relationship between the interest rate differential of two countries and the expected currency depreciation. Although there is more favourable support in recent studies¹, an empirical failure has usually been found when

*The authors acknowledge Seminar participants and Jesús Crespo-Cuaresma for their helpful comments, D. Andrews for providing his code. Marcos Sanso Navarro also acknowledges Gobierno of Aragón and Caja de Ahorros de la Inmaculada for financial support. Corresponding author: Marcos Sanso-Navarro. Departamento de Análisis Económico. Gran Vía, 2. 50005 Zaragoza (Spain). Tel: (+34) 976761000 (Ext 4728) Fax: (+34) 976761996. e-mail: marcosn@unizar.es

¹See Baillie and Bollerslev (2000), Bekaert and Hodrick (2001), Chaboud and Wright (2003) and Bekaert et al (2005).

tested with real data².

In this paper it is analyzed over long time horizons in the G7 countries during the post-Bretton Woods era. Related works are those in Flood and Taylor (1997), Alexius (2001), Berk and Knot (2001), Zhang (2004) and Chinn and Meredith (2004), who generally argue there is more evidence for the UIP when it is considered over wide periods of time.

As pointed in Campbell et al (1997), long-horizon regressions generally find favourable and significant results where short-horizon ones find any. Examples of this fact are those in Fama and French (1988) and Campbell and Shiller (1988) for equity returns predictability, Mishkin (1992) for the Fisher Effect testing and Mark (1995) and Rapach and Wohar (2002) who analyze the relationship between the deviation of long-horizon exchange rate from the fundamental value with their returns. Valkanov (2003), Mark and Sul (2004) and the references therein deal with this issue.

It is demonstrated in Valkanov (2003) that, in addition to incorrect testing, the use of long-horizon variables might lead to inconsistent estimators and coefficients of determination that do not converge in probability to one. This is directly related to the non-stationary behavior the variables involved in this kind of regressions usually exhibit. These results are similar to those in Granger and Newbold (1974) and Phillips (1986, 1991), where the analogy relies in finding a spurious correlation between persistent variables when they are, in fact, statistically independent.

Following this arguments, I question previous results about long-horizon UIP. First of all, the order of integration of the variables used when testing this theoretical relationship over long time periods will be analyzed. Zhang (2004) is the only work among those mentioned above that cares about the possible non-stationarity of the expected currency depreciation and the long-term interest rates differentials. It is found that they behave as non-stationary but GMM estimation methods are used. This latter work also questions why UIP tests gain power over horizon and relates it with the frequency bands of the data concluding they are more interesting than horizon in this respect.

²See surveys in Froot and Thaler (1990) and Engel (1996).

It is also found in my data evidence that long-term variables behave as non-stationary, so long-horizon UIP is tested using the residual-based cointegration methods in Engle and Granger (1987). I consider the results in Chinn and Meredith (2004) and Zhang (2004) as doubtful because evidence of cointegration between the variables is found in only one of the six countries (the United Kingdom). Since the relationship is analyzed over an extensive time period, we can think in some shifts occurring along it and take into account the possible presence of structural breaks. This is done in the spirit of Gregory and Hansen (1996) and an extension of their test to the case of two shifts is proposed. It is precisely the introduction of two shifts what becomes a key element for obtaining some additional evidence of cointegration and the long-horizon UIP relationship to hold.

For all countries, structural breaks are located around similar dates corresponding to changes in the monetary policy of the Federal Reserve and in the evolution of the U.S. Dollar exchange rate. The parameter allowed to vary is the constant term in the UIP testing regression, usually related to the risk premium. For this reason, I justify breakpoints found with the evolution of the risk premium inherent in the long-term bonds. It is calculated following the method proposed in Crespo-Cuaresma et al (2005), which is based on the expectations hypothesis of the term structure.

As pointed before, monetary policy changes and exchange rates evolution seem to influence UIP. At the end of the sample an important event related with both has occurred: the introduction of the common monetary policy and currency in the Eurozone. Cointegration methods allowing for breaks do not consider observations at the end of the sample where its effects could be present. In addition, the two countries for which the presence of cointegration with shifts is found are France and Germany. For these reasons, a robustness check of our results is implemented using the end-of-sample cointegration breakdown tests in Andrews and Kim (2003).

The essay is structured as follows. Section 2 defines the concept of UIP, describes the usual ways of testing it and summarizes recent empirical findings. Special emphasis is made on long-horizon testing and the problems inherent to this kind of regressions. Section 3

describes and analyzes the data used and ‘conventional’ UIP tests are applied. Section 4 applies cointegration methods that takes into account the presence of structural breaks to long-horizon UIP testing. Section 5 calculates the risk premium inherent to long-term interest rates and relates the breaks found with changes in its evolution. Section 6 implements a robustness check of our inferences to the possible presence of shifts at the end of the sample and section 7 concludes.

2 Long-horizon UIP. Background and empirical evidence

Uncovered interest parity (UIP) establishes a relationship between the expected currency depreciation and the interest rates differential of the corresponding countries. If it holds, under the assumption of risk neutral investors and rational expectations, the market forecast is taken implicitly into account in the international differences of the k-period interest rates. The basic regression to test this relationship is as follows:

$$\Delta s_{t+k} = \alpha + \beta(i_{t,k} - i_{t,k}^*) + \eta_{t+k} \quad (1)$$

Δs_{t+k} is the currency depreciation (change in the logarithm of the spot price of foreign currency in terms of the domestic one) over k-periods and $(i_{t,k} - i_{t,k}^*)$ is the k-period national interest rate less the corresponding foreign one. η_{t+k} is an error term.

Testing the null hypothesis that UIP holds is equivalent to test that the slope parameter in (1) is equal to unity. Another possibility is to extend this hypothesis with the fact that the constant parameter is equal to zero³. However, relaxing the assumption of risk-neutral investors this last parameter may reflect the (time-invariant) risk-premium demanded on foreign assets.

Another specification of equation (1) is obtained when replacing the interest rate differential by the forward discount (difference between the forward and spot exchange rates). It is based on arbitrage arguments and can be interpreted as, under the same assumptions as

³In this case, it is a joint test of UIP and rational expectations.

before, the forward discount to be an unbiased estimator of the subsequent exchange rate depreciation. A very large literature has tested the ‘unbiasedness hypothesis’ and find the coefficient β to be considerably less than unity. In the survey of Froot and Thaler (1990) an average coefficient across 75 studies equal to -0.88 is reported⁴. This difference is difficult to explain, but some common interpretations are those related with the possible existence of a time-varying risk premium, expectational errors or ‘peso’ problems. Engel (1996) contains another survey with similar conclusions.

Recent studies are more optimistic about the empirical performance of UIP. Baillie and Bollerslev (2000) argue this anomaly to be an statistical phenomenon as a consequence of the persistence of the forward premium and the small sample sizes used and Bekaert and Hodrick (2001) also focus on the finite sample properties of the tests. Chaboud and Wright (2003), using intradaily data, find that UIP works better in the very short run as the risk premium shrinks to zero. Bekaert et al (2005) conclude that evidence against UIP is mixed and currency, not horizon, dependent.

Among the studies supporting more favourable (or, at least, not as negative) evidence for UIP are those that analyze it over long time horizons. Flood and Taylor (1997) used IFS medium-term government bonds and found a coefficient on the interest differential of 0.596 (standard error 0.195), rejecting both the zero and one null hypotheses for the slope parameter. Alexius (2001) works with quarterly IFS long-term interest rates from 1957:1 to 1997:4 and, once dealing with the possible problems inherent to the data, finds substantial evidence in favor to the unbiasedness hypothesis. Chinn and Meredith (2004) criticize that the sample in Alexius (2001) covers periods of fixed exchange rates and extensive controls and consider the 10-year government bonds and ‘constant maturity’ 5 and 10-year yields during 1980:1-2000:4. They find more support for the UIP and give an interpretation from a macroeconomic point of view extending the model in McCallum (1994). Zhang (2004) makes a similar analysis for both monthly and quarterly data and link the favourable evidence to the low and medium

⁴McCallum (1994) argues that the UIP relationship is distinct from, and more important than, the unbiasedness of forward exchange rates as predictors of future spot rates. That is, negative estimations of the slope parameter implies rejection of unbiasedness but not necessarily UIP.

frequency bands⁵ of the data instead of the horizon. Note that in these studies no distinction about the evidence among countries is made.

Up to date, empirical success of UIP over wide time periods can be related to the problems practitioners face when working with variables measured during long horizons. In the context of stock returns predictability, Campbell et al (1997) were the first who explicitly posed the question of why long-horizon regressions have more power to reject the null of a zero slope parameter. They pointed that estimated slope coefficients, t-statistics and determination coefficients increased with the time period over which returns were calculated. Since then, several studies have appeared trying to explain why it happens and making proposals in order to apply adequate techniques given the characteristics usually displayed by long-horizon variables as those in Valkanov (2003), Mark and Sul (2004) and the references therein.

Using asymptotic arguments and a ‘local-to-unity’ framework, Valkanov (2003) explains the tendency of long-horizon regressions towards finding ‘significant’ results where previous short-term approaches find any. The explanation relies on the persistent and possible non-stationary behavior of the variables involved in this kind of regressions. In addition to incorrect testing due to serial correlation in the error terms, it is found that inconsistent estimators and a coefficient of determination that does not converge to one in probability can be obtained. Results about the possibility of finding an spurious correlation between two persistent variables can also be found in Granger and Newbold (1974), Phillips (1986, 1991) and Ferson et al (2003).

It can be concluded that estimation and testing using long-horizon variables cannot be carried out applying ‘conventional’ regression methods as Ordinary Least Squares. Estimated parameters using this latter method will be super-consistent if the variables, although non-stationary, are cointegrated. However, even if this latter circumstance applies it is needed to correct for biases and distributional divergence in the t-statistics when working with small samples⁶.

⁵Following Baxter (1994) this two frequency bands are, respectively, a ‘trend’ component that shows fluctuations in the data which exceed 32 quarters in length and a ‘business cycle’ one for those lasting between 6 and 32 quarters.

⁶This issue will be described with more detail in section 5.

We can think in previous long-horizon UIP tests to also present this problem. Zhang (2004) is the only work among those mentioned before that cares about the order of integration of both long-term exchange rates depreciation and interest rates differentials. He finds they behave as non-stationary but still applies UIP testing in (1) with the variables in levels as in Chinn and Meredith (2004). The main difference with this latter work is that estimation is carried out using the Generalized Method of Moments (GMM) due to the possible correlation of the error term with the explanatory variable.

Given the evidence of a possible non-stationary behavior of both the endogenous and explanatory variables in long-horizon UIP testing and the problems inherent to this kind of regressions it seems more appropriate to use cointegration methods. In the case we find those variables to be cointegrated we could trust previous evidence about long-horizon UIP but, on the contrary, we should expect them to suffer a spurious regression problem and question their results in the case cointegration is not present.

3 Data description and analysis

The data analyzed in this paper correspond to the G7 countries, has quarterly frequency and covers the period 1973:1 to 2004:3. Exchange rates and long-term government bond yields are obtained from the IMF International Financial Statistics CD-Rom and the short-term (3-month) interest rates from the OECD Main Economic Indicators database.

Given the specification in (1) and since we are considering rational expectations, the endogenous variable is the annualized change in the logarithm of the exchange rate (national currency in terms of U.S. Dollars) over k periods⁷. The regressor is equal to the difference between national and U.S. yields (expressed in percent per annum) for a k -period horizon at time t .

Tables I and II show the resulting statistics obtained when applying a battery of unit root tests discussed in Ng and Perron (2001) to both short ($k = 1$) and long ($k = 40$) horizon

⁷Japanese data needs interpolation for the long-term depreciation in 1985:1 and 1985:2 as a consequence of the annualization.

variables. The deterministic components considered include both a constant and a trend⁸. It can be observed for the short-term variables that the unit root null is more generally rejected for the case of the series corresponding to the annualized exchange rates than to those for the 3-month interest rates differentials. This gives a first idea of the initial suspect that this latter variable has a considerable degree of persistence and can behave as non-stationary.

Results in line with the findings in Zhang (2004) are obtained when these tests are applied to the long-horizon variables. It is mainly observed that both the endogenous and explanatory variables used in long-horizon UIP testing to behave as integrated of order one during the sample considered. The main exceptions are those regarding the long-term interest rates differentials for Canada and France, for which can be found some little rejection of the null of a unit root. Following this results, we conclude that variables involved in long-horizon UIP testing are highly persistent and generally behave as non-stationary in the sample. For that reason, we will proceed as if they were integrated of first order and analyze the relationship using cointegration methods.

The use of the IFS data in long-horizon UIP testing has been criticized in Alexius (2001) due to the different maturities of long-term government bond yields and the existence of coupon payments. For that reason, she made some data manipulations to make them homogeneous. We prefer not to make any correction and work with raw data. That is why we are implementing a ‘conventional’ UIP analysis in order to establish a bridge between previous studies and this one. Our closest reference is Chinn and Meredith (2004) due to the similarity in the sample period, data frequency and countries considered. Their results are very similar to those obtained below.

Short-horizon UIP tests regress a stationary and volatile endogenous variable (Δs_t) on a highly persistent one ($i_t - i_t^*$) by OLS methods. Common findings have been obtained when implementing this kind of analysis to our dataset as can be concluded from the results in Table III. With the exception of the Italian Lira, all estimated slope parameters are negative, standard errors are high and the adjusted coefficients of determination low (if not negative).

⁸The consideration of only a constant as the deterministic component does not change the main conclusions drawn.

The null that the slope parameter is equal to unity is rejected for all cases, except Italy. If it can be accepted for any significance level it is as a consequence of the high standard errors.

Table IV shows the estimation results obtained from a ‘conventional’ long-horizon UIP analysis. It can be observed that now, again with the Italian exception, all estimated slope parameters are positive pointing to a more favourable evidence regarding UIP. Due to the overlapping observations, a MA component can be present so Newey and West (1987) corrected standard errors are used. Coefficients of determination are greater, although not much for most cases. Resulting Durbin-Watson (1971) statistics are considerably low, pointing to the presence of serial correlation in the residuals. The null hypothesis that the slope parameter is equal to unity can only be rejected for the cases of the Italian Lira and the British Pound.

Summarizing, long-horizon variables involved in UIP testing behave as unit-root non-stationary in the sample considered. For that reason, it seems more convenient to test this relationship using cointegration methods. Since results in Tables III and IV are similar to those obtained in Chinn and Meredith (2004) we will avoid the problem of data appropriateness and use them as a comparison benchmark.

4 Cointegration analysis of long-horizon UIP

If both endogenous and explanatory variables in (1) are persistent a spurious regression problem may be underlying previous long-horizon UIP tests. This will not be the case if those variables are cointegrated because OLS estimations of the slope parameters are super-consistent (converge at a rate T , instead of $T^{1/2}$ for the true value⁹).

Long-horizon UIP will only be reliable for those countries whose exchange rate depreciation and interest rates differential are cointegrated. Once determined in which of those considered cointegration holds inferences about the parameters will be made.

The way I am testing for the presence of cointegration in this here is by the use of the residual-based tests, first developed in Engle and Granger (1987). Following this latter au-

⁹However, they will not be efficient and biased in finite samples away from the null. For that reason, other estimation methods than OLS will be needed.

thors, two $I(1)$ variables are cointegrated if a linear combination of them has a stationary distribution. The null hypothesis is the absence of cointegration among the variables. These tests are implemented as follows:

Let $y_t = (y_{1t}, y'_{2t})$ be the observed data. y_{1t} is a scalar and y_{2t} a m -dimensional vector. The cointegration model is given by

$$y_{1t} = \mu + \gamma y_{2t} + e_t, \quad t = 1, \dots, T \quad (2)$$

where y_{1t} is $I(1)$, e_t is $I(0)$, μ a parameter and γ a $(1 \times m)$ vector of parameters.

The relationship in (2) is estimated by OLS and a unit root test is applied to the regression residuals. If the null of a unit root can be rejected it can also be rejected the null of no cointegration between the variables, and viceversa.

Resulting ADF residual-based test statistics for equation (1) using the variables measured over 10 years are those reported in the upper panel of Table V. It can be observed that the null of no-cointegration can only be rejected for the case of the British Pound. Then, the United Kingdom is the only country among those analyzed for which the long-horizon UIP can be satisfied. This result is at odds with those obtained from the ‘conventional’ approach in the previous section where it was one of the two countries for which this relationship was rejected.

However, we can think of cointegration as a relationship maintained over some (fairly long) period of time and then shifting to a new one. In this case, standard residual-based tests are not appropriate since they presume the cointegrating vector to be time invariant under the alternative. The consideration of the presence of breaks makes sense in the context we are analyzing since the period spans around three decades where policy changes took place and could have exerted some kind of influence.

The extension of the method in Engle and Granger (1987) to the case where structural breaks are present has been proposed in Gregory and Hansen (1996). They develop cointegration tests allowing for a regime shift at an unknown date. The way changes are modeled

is through the use of a dummy variable that takes a unitary value after the break date:

$$\varphi_{t\tau} = \begin{cases} 0 & \text{if } t \leq [\tau \cdot T] \\ 1 & \text{if } t > [\tau \cdot T] \end{cases} \quad (3)$$

where the (unknown) parameter $\tau \in (0, 1)$ denotes the relative timing of the change point and $[\cdot]$ is the integer function.

A first way a structural change can occur is in the form of a level shift in the cointegration relationship. It is denoted as model C and is equivalent to a change in the constant in (2). This implies the equilibrium relationship to shift in a parallel way. A second possibility is a regime shift in which both constant and slope parameter change (model C/S).

$$\text{Model C} : y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \gamma y_{2t} + e_t \quad t = 1, \dots, T \quad (4)$$

$$\text{Model C/S} : y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \gamma_1 y_{2t} + \gamma_2 y_{2t} \varphi_{t\tau} + e_t \quad t = 1, \dots, T \quad (5)$$

Standard methods to test the null of no cointegration with breaks are also based on the analysis of the residuals. That is, the candidate cointegration relationship is estimated via OLS and a unit root test is applied to the regression errors. This is implemented for all possible dates where the break can occur ($\tau \in (0.15, 0.85)$, for example) and the smallest value of the resulting statistic is considered. Estimated breakpoint is that date for which this value is the minimum:

$$ADF^* = \inf_{\tau \in (0.15, 0.85)} ADF(\tau) \quad (6)$$

The way I am modeling structural change is that of model C in (4) where a change in the constant term takes place. This parameter reflects the risk premium in UIP testing and the consideration of an invariant one over 30 years can be considered a thought assumption. It can be expected that policy and macroeconomic developments along this period to change agent's perception of risk. I am not using complicated models as those in (5) because their interpretation is more difficult and less intuitive. Resulting test statistics and estimated break dates are those displayed in the second panel of Table V. The introduction of a shift in the

constant term does not allow to reject the null of no-cointegration for any of the currencies in our dataset. However, it can be observed that there are some similarities in the location of the break dates. The breakpoint is 1976:3 for Canada and Germany, at the beginning of 1988 for France and Japan, 1979:1 for Italy and 1982:4 for United Kingdom. Then, the introduction of a shift in the constant term does not improve the weak evidence of cointegration between the variables used in long-horizon UIP in our sample.

We can think in the period analyzed to be wide enough to allow for the presence of an additional break . For that reason, I propose to extend the method in Gregory and Hansen (1996) to the case where two breaks are present in the cointegration relationship. They occur at the unknown break dates $\tau_1 \cdot T$ and $\tau_2 \cdot T$ and are also modeled by the use of dummy variables:

$$\begin{aligned}\varphi_{t\tau}^1 &= \begin{cases} 0 & \text{if } t \leq [\tau_1 \cdot T] \\ 1 & \text{if } t > [\tau_1 \cdot T] \end{cases} \\ \varphi_{t\tau}^2 &= \begin{cases} 0 & \text{if } t \leq [\tau_2 \cdot T] \\ 1 & \text{if } t > [\tau_2 \cdot T] \end{cases}\end{aligned}\tag{7}$$

Since it is the case relevant for this analysis, I only focus in two level shifts¹⁰. Then, the formulation of model C with two breaks is as follows:

$$\text{Model C (2 breaks)} : y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau}^1 + \mu_3 \varphi_{t\tau}^2 + \alpha y_{2t} + e_t \quad t = 1, \dots, T \tag{8}$$

and the test statistic is calculated as:

$$ADF^{**} = \inf_{\substack{\tau_1, \tau_2 \in (0.15, 0.85) \\ \tau_1 < \tau_2}} ADF(\tau_1, \tau_2) \tag{9}$$

Following Theorem in page 109 of Gregory and Hansen (1996) the limiting distribution of

¹⁰Similar arguments apply for the case of two regime shifts. Interpretation of the results in this case for our analysis will be even more difficult.

this test under the null will be given by:

$$ADF^{**} \xrightarrow{d} \inf_{\substack{\tau_1, \tau_2 \in (0.15, 0.85) \\ \tau_1 < \tau_2}} \frac{\int_0^1 W_\tau dW_\tau}{\int_0^1 W_\tau^2} \quad (10)$$

where

$$W_\tau(r) = W_1(r) - \int_0^1 W_1 W_{2\tau}^\top \left[\int_0^1 W_{2\tau} W_{2\tau}^\top \right]^{-1} W_{2\tau}(r) \quad (11)$$

and

$$W_{2\tau}(r) = [1, \varphi_{t\tau}^1, \varphi_{t\tau}^2, W_2^\top(r)]^\top \quad (12)$$

As in the original work, critical values are generated via Montecarlo simulations. Using 10,000 replications, they are found ($Crt(T, p, m)$) for different sample sizes ($T=50, 100, 150, 200, 250$ and 300), significance levels (p) and number of regressors (m). Once obtained, it is estimated a response surface for each m and p combination. The regressions applied are of the following form:

$$Crt(T, p, m) = \Psi_0 + \Psi_1 T^{-1} + error \quad (13)$$

Asymptotic critical values are the OLS estimations of Ψ_0 . Once generated, we can compare them with the resulting test statistics displayed in the third panel of Table V.

The introduction of a second shift in the constant term leads to the rejection of the null of no cointegration for France and Germany. Estimated break dates in both cases are very close to each other: the first break is located at 1981:1 in France and 1983:2 in Germany and the second one in 1987:2 and 1986:4, respectively. These breaks are also similar to those found in the rest of the countries where no evidence of cointegration is found, pointing to the existence of a common cause of this shifts.

Break dates are located around the periods 1979-1983 and 1985-1987. Since we are considering the exchange rate of each national currency against the U.S. Dollar the most immediate interpretation is that they reflect changes determined by the evolution in the monetary policy of the Federal Reserve and the exchange rates. The first interval correspond to the period when Paul Volcker became the Chairman of the Board of Governors and implemented changes

in the monetary aggregates targeting in order to mitigate inflationary pressures. It also coincides with a period between the late 70s and 1985 when there was a sharp depreciation of the U.S. Dollar. The second of these periods begins around 1987, when Alan Greenspan substituted his predecessor and the evolution of the exchange rate changed towards a strong appreciation that lasted until the mid 90s.

Following previous analysis, it can be concluded that long-horizon UIP does not hold in Canada, Italy and Japan for the sample considered since there seems to be an absence of a long-run relationship between long-horizon exchange rate depreciation and interest rates differentials¹¹. Cointegration is only found in three of the countries analyzed: France, Germany and United Kingdom, for which we are going to test the parameter restrictions implied by the UIP.

As pointed before, standard OLS estimation leads to biased estimated parameters and test statistics that do not converge to a normal distribution when working with small samples. This occurs even if the variables are cointegrated and the ‘Fully-Modified’ OLS estimation method of Phillips and Hansen (1990), which introduces a correction using semiparametric methods, is more appropriated to make inferences. Estimation results are those in Table VI.

With the exception of the constant term, all estimated parameters are statistically significant. Furthermore, the null hypothesis of a unitary slope cannot be rejected for any of the three countries and the estimated time-varying constant parameters are of the same sign and similar magnitude in the two cases they are introduced. The adjusted coefficients of determination have increased with respect to the ‘conventional’ long-horizon tests. Conclusions drawn using the cointegration approach are also very different to those from the analysis in section 4. In this latter case, long-horizon UIP could not be rejected in four of the six countries. Due to the documented persistence in regressors it can be expected this evidence, as that in similar studies, to be spurious. Once cointegration methods are used and appropriate estimation techniques applied the conclusions regarding the British Pound is reversed and

¹¹Also note that the introduction of an additional third level shift does not change these conclusions. In addition, the possible consideration of a model with a change in regime does not lead to find any evidence in favour of cointegration.

no evidence with respect the other countries is found. Only when two shifts in the constant term are introduced the results coincide with those drawn from the previous methodology for the case of France and Germany, disappearing the favourable results for the Canadian and Japanese ones. The disfavourable evidence for Italy still persists.

5 Identifying breaks with changes in the risk premium

Existing evidence among the G7 countries regarding long-horizon UIP relies in a great extent on the assumption of a changing constant parameter, what is usually related to the existence of a time-varying risk premium.

In this section the risk premium of long-term interest rates for each of the countries is going to be estimated following the method proposed in Crespo-Cuaresma et al (2005), based on the Rational Expectations Hypothesis of the Term Structure (REHTS). This will allow to determine if the results in the previous section are casual or, on the contrary, they are really related to the existence of changes in the evolution of the risk premium.

The REHTS states that the yield to maturity of a k -period bond can be decomposed into the one-period yields and the risk premium. Following the notation used along this paper it can be formulated as:

$$i_{t,k} = \frac{1}{k} \sum_{i=0}^{k-1} E_t(i_{t+i,1}) + \phi(k,t) \quad (14)$$

where $E_t(\cdot)$ is the conditional expectation operator using the info available at period t , $i_{t,k}$ is the yield to maturity of a k -period bond and $\phi(k,t)$ is the average risk premium on a k -period bond until it matures.

As it has been assumed before, there are rational expectations so we are considering agents have perfect foresight with respect to the one-period yields. Due to the quarterly frequency of our data the short-term interest rates correspond to a 3-month period and the long-term one is the 10-year government bond yield.

Resulting long-term interest rates risk premia are plotted in Figure I. It can be observed they follow a time-varying pattern, where the most abrupt changes take place around the break

dates found in the previous section. For almost all series, there is a peak around 1981-1982 and a valley in 1986-1987. The main differences between countries are in the late 70s.

It is in the estimated series of the risk premium for the United States where the existence of three different regimes (two break dates) is more evident. It must also be noted that the less variable serie that can better be approximated by a constant value is that corresponding to the United Kingdom, for which cointegration between exchange rate depreciation and interest rates differential was found without the need of imposing any level shifts.

Summarizing, allowing for the existence of breaks in the UIP relationship when testing it over long-horizons is reasonable in light of the estimated risk-premia series using the REHTS hypothesis, given their variability and the location of the main changes.

6 Robustness check. End of sample cointegration stability

As can be derived from (6) and (9), the method in Gregory and Hansen (1996) left some observations without being breakpoint candidates both at the beginning and the end of the sample. However, some shifts can occur in those dates affecting the estimated cointegration relationship.

The two countries for which the long-horizon UIP holds with a double change in the constant are France and Germany. Breaks also seem to be determined by monetary policy and exchange rates developments. An important event related with this two latter aspects has taken place in the period analyzed for this two countries that is ignored by the method in Gregory and Hansen (1996): the introduction of the common monetary policy and currency in the European Monetary Union. For this reason, it is proposed in this section a robustness check to determine if there has been any change in the relationships at the end of the sample.

The way I am implementing it is through the use of the method developed in Andrews and Kim (2003) to detect possible cointegration breakdowns over short periods of time. The types of breaks it considers are both a change in the cointegration vector or a shift in the errors

distribution from being $I(0)$ to $I(1)$. The test statistics are constructed from the following model:

$$y_t = \begin{cases} x_t' \beta_0 + u_t & \text{for } t = 1, \dots, T \\ x_t' \beta_t + u_t & \text{for } t = T + 1, \dots, T + m \end{cases} \quad (15)$$

y_t is the endogenous variable, x_t are the explanatory ones and u_t denotes the error term. β_0 and β_t are the parameter vectors. $y_t, u_t \in R$ and $x_t, \beta_0, \beta_t \in R^k$.

The null and alternative hypotheses can be formulated as follows:

$$\begin{aligned} H_0 &: \begin{cases} \beta_t = \beta_0 \text{ for all } T + 1, \dots, T + m \text{ and} \\ \{u_t : t = 1, \dots, T + m\} \text{ are stationary and ergodic} \end{cases} \\ H_1 &: \begin{cases} \beta_t \neq \beta_0 \text{ for some } t = T + 1, \dots, T + m \text{ and/or} \\ \text{the distribution of } \{u_{T+1}, \dots, u_{T+m}\} \text{ differs from} \\ \text{the distribution of } \{u_1, \dots, u_m\} \end{cases} \end{aligned} \quad (16)$$

Testing for a change in the cointegration vector is made through a test statistic that is a quadratic form in the ‘post-breakdown’ residuals using a ‘pre-break’ estimator.

$$P_b = \sum_{t=T+1}^{T+m} (y_t - x_t' \hat{\beta}_{1-(T+\lceil m/2 \rceil)})^2 \quad (17)$$

$\lceil m/2 \rceil$ denotes the smallest integer that is greater than or equal to $m/2$ and $\hat{\beta}_{1-(T+\lceil m/2 \rceil)}$ is the OLS estimator obtained using the observations $t = 1, \dots, T + \lceil m/2 \rceil$

The null hypothesis is rejected if the test statistic exceeds a critical value determined using a parametric subsampling method. Consider $\{P_{b,j} = P_j(\hat{\beta}_j) : j = 1, \dots, T - m + 1\}$, being

$$P_{b,j} = \sum_{t=j}^{T+m} (y_t - x_t' \hat{\beta}_{(j)})^2 \quad (18)$$

where $\hat{\beta}_{(j)}$ is the estimator of β using $t = 1, \dots, T$ with $t \neq j, \dots, j + m - 1$.

Its empirical distribution function is:

$$\hat{F}_{P_b, T}(x) = \frac{1}{T - m + 1} \sum_{t=1}^{T-m+1} 1(P_{b,j} \leq x) \quad (19)$$

so the critical value at the α significance level is equal to the $1 - \alpha$ sample quantile

$$\hat{q}_{P_b, 1-\alpha} = \inf \left\{ x \in R : \hat{F}_{P_b, T}(x) \geq 1 - \alpha \right\} \quad (20)$$

In order to test for a change in the error distribution from an $I(0)$ process to a non-stationary one in the last m observations it is used a locally best invariant test which has the same spirit as (17) and consists in a sum of squares of reverse partial sum of residuals after the breakpoint:

$$R_b = \sum_{t=T+1}^{T+m} \left(\sum_{s=t}^{T+m} (y_s - x'_s \hat{\beta}_{1-(T+\lceil m/2 \rceil)}) \right)^2 \quad (21)$$

The test is implemented in the same way as the previous one.

In the application of these tests it has been considered as possible break dates not only those ignored by the Gregory and Hansen (1996) method (which include the observations corresponding to 2001:2 onwards) but also all observations constructed using data after 1999:1, when the European common monetary policy began its functioning.

Results are those in the four graphs at the bottom of Figure II where p-values of each test for France and Germany are plotted and compared with a dotted line representing the 5% significance level. At each date, it is reported the p-value of testing the null of a well-specified cointegration regression model for all dates against it to be only correct for the observations before. The null cannot be rejected in any of the four cases even at the 10% significance level. Then, it can be concluded the introduction of the European common monetary policy and currency has not had any influence in the relationships determined before, so the resulting inferences drawn from them are still valid in this latter period.

The same exercise has been done for the observations trimmed at the beginning of the

sample, p-values are those at the top of Figure II, from which similar conclusions are derived.

7 Conclusions

Existing studies testing the uncovered interest parity over wide periods of time do not care about the non-stationary behavior of the variables involved. I analyzed here long-horizon UIP using cointegration methods for the G7 countries in the post-Bretton Woods era. Once determined their order of integration, cointegration between the exchange rate depreciation and interest rate differential is only found for one of the six countries: the United Kingdom. This finding makes me to question previous results that support a stronger evidence in favour of this relationship when it is considered over a long period of time. Due to the amplitude of the sample the existence of up to two structural breaks is introduced following the method in Gregory and Hansen (1996). This allows to find additional evidence of cointegration for France and Germany. This shifts are justified by the presence of a time-varying risk premium and, for all the countries where cointegration is present, the null of UIP cannot be rejected. Finally, the robustness of our inferences to possible breakdowns in the cointegration relationships at the end of the sample due to recent monetary developments in the Eurozone has been assessed.

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Table I. Unit root test for the variables involved in short-horizon UIP testing (3-month. $k = 1$ for quarterly data). 1973:Q1-2004:Q2

Country		Canada		France		Germany		Italy		Japan		United Kingdom	
Variable		$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$
c.v.(5%)	Test												
-17.30	Z_a	-115.81	-11.91	-37.38	-14.91	-32.83	-16.16	-84.29	-6.68	-46.91	-21.32	-88.69	-13.81
-17.30	MZ_a	-62.15	-11.24	-9.72	-14.02	-7.20	-15.82	-49.46	-6.32	-16.83	-20.62	-51.23	-12.85
-2.91	MZ_t	-5.50	-2.36	-2.18	-2.65	-1.86	-2.80	-4.97	-1.76	-2.89	-3.20	-5.06	-2.53
0.17	MSB	0.09	0.21	0.22	0.19	0.26	0.18	0.10	0.28	0.17	0.16	0.10	0.20
5.48	ERS Pt	1.80	9.11	10.44	6.41	14.38	6.36	1.82	16.15	5.93	4.42	1.76	7.36
5.48	Mod Pt	1.80	8.15	9.47	6.51	12.71	5.85	1.85	14.43	5.51	4.46	1.79	7.10
-2.91	DF – GLS	-10.25	-2.50	-3.10	-2.81	-2.65	-2.46	-8.47	-1.86	-3.88	-3.34	-8.76	-2.72

Note : Tests implemented are those discussed in Ng and Perron (2001).

Table II. Unit root test for the variables involved in long-horizon UIP testing (10-years. $k = 40$ for quarterly data). 1973:Q1-2004:Q2

Country		Canada		France		Germany		Italy		Japan		United Kingdom	
Variable		$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$	$\Delta^k s_t$	$i_{t,k} - i_{t,k}^*$
c.v. (5%)	Test												
-17.30	Z_a	-2.63	-18.74	-3.31	-18.20	-4.66	-2.52	-3.40	-5.90	-3.49	-5.01	-7.05	-6.17
-17.30	MZ_a	-2.58	-16.69	-3.23	-16.22	-4.53	-2.35	-3.29	-5.55	-3.29	-4.86	-6.70	-5.91
-2.91	MZ_t	-1.06	-2.89	-1.24	-2.84	-1.50	-1.02	-1.28	-1.66	-1.21	-1.55	-1.83	-1.71
0.17	MSB	0.41	0.17	0.38	0.17	0.33	0.43	0.39	0.30	0.37	0.32	0.27	0.29
5.48	ERS Pt	31.91	5.37	27.25	5.76	19.69	38.81	28.00	18.00	28.19	18.34	13.98	15.56
5.48	Mod Pt	32.49	5.46	27.60	5.64	20.08	35.95	27.69	16.41	26.23	18.71	13.60	15.40
-2.91	DF – GLS	-1.08	-3.24	-1.27	-3.19	-1.54	-1.10	-1.32	-1.76	-1.29	-1.60	-1.92	-1.79

Note : Tests implemented are those discussed in Ng and Perron (2001).

Table III. Short-horizon ($k = 1$) UIP regressions. 1973:Q1-2004:Q2

	α	β	$H_0: \beta = 1$	$\bar{R}^2$	DW
Canada	0.02 [†] (0.01)	-0.44 (0.51)	-2.82*	-0.00	1.92
France	0.04 (0.03)	-1.21 (1.11)	-1.99**	0.00	1.81
Germany	-0.00 (0.02)	-0.73 (0.89)	-1.94***	-0.00	1.92
Italy	0.02 (0.04)	0.64 (1.15)	-0.31	-0.00	1.63
Japan	-0.10 [†] (0.03)	-3.30 [†] (0.79)	-5.44*	0.09	2.09
United Kingdom	0.05 (0.03)	-1.81 (1.48)	-1.90***	0.02	1.75

Note : Newey and West (1987) standard errors in parentheses.

[†] and [‡] means statistical significance at the 1 and 5 % level.

*, ** and *** denotes rejection at the 1, 5 and 10 % significance level

Table IV. Long-horizon ($k = 40$) UIP regressions. 1983:Q1-2004:Q2

	α	β	$H_0: \beta = 1$	$\bar{R}^2$	DW
Canada	0.00 (0.00)	0.65 [†] (0.29)	-1.21	0.07	0.11
France	-0.01 (0.01)	1.31 (0.94)	0.33	0.09	0.08
Germany	-0.00 (0.00)	1.35 [†] (0.31)	1.13	0.48	0.19
Italy	0.04 [†] (0.01)	-0.25 (0.24)	-5.21*	0.02	0.08
Japan	-0.04 [†] (0.01)	0.94 (0.57)	-0.10	0.11	0.10
United Kingdom	-0.00 (0.00)	0.58 [†] (0.13)	-3.24*	0.45	0.25

Note : Newey and West (1987) standard errors in parentheses.

[†] and [‡] means statistical significance at the 1 and 5 % level.

*, ** and *** denotes rejection at the 1, 5 and 10 % significance level

Table V. Residual-based cointegration tests. Long-horizon UIP. 1973:Q1-2004:Q2

		Canada	France	Germany	Italy	Japan	United Kingdom
Engle and Granger (1987)							
c.v. (5%):	ADF	-3.17	-2.45	-2.19	-2.91	-2.01	-1.91
							-3.39**
Gregory and Hansen (1996)							
1 level shift							
c.v. (5%):	ADF*	-4.83	-3.52	-2.29	-3.24	-3.21	-3.53
	date		1976:3	1988:1	1976:3	1979:1	1988:2
							1982:4
Gregory and Hansen (1996)							
2 level shifts							
c.v. (5%):	ADF**	-5.44	-4.00	-5.79**	-5.50**	-4.37	-4.82
	date 1		1979:4	1981:1	1983:2	1981:4	1982:2
	date 2		1981:4	1987:1	1986:3	1987:4	1986:2
							1985:1

Note : *, ** and *** denotes rejection at the 1, 5 and 10 % significance level

Table VI. Cointegration relationship estimation. Long-horizon UIP.

Fully modified OLS. 1973:Q1-2004:2

Country	t_1	t_2	μ_1	μ_2	μ_3	γ	$H_0: \gamma = 1$	$\bar{R}^2$
United Kingdom			-0.01 (0.01)			0.69 [†] (0.27)	-1.15	0.59
France	1981:1	1987:1	0.02 (0.01)	-0.09 [†] (0.01)	0.07 [†] (0.01)	1.34 [†] (0.48)	0.71	0.28
Germany	1983:2	1986:3	-0.01 (0.01)	-0.07 [†] (0.01)	0.09 [†] (0.02)	0.64 [†] (0.28)	-1.29	0.46

Note : Standard errors in parentheses.[†] and [‡] means statistical significance at the 1 and 5 % level. *, ** and *** denotes rejection at the 1, 5 and 10 % significance level

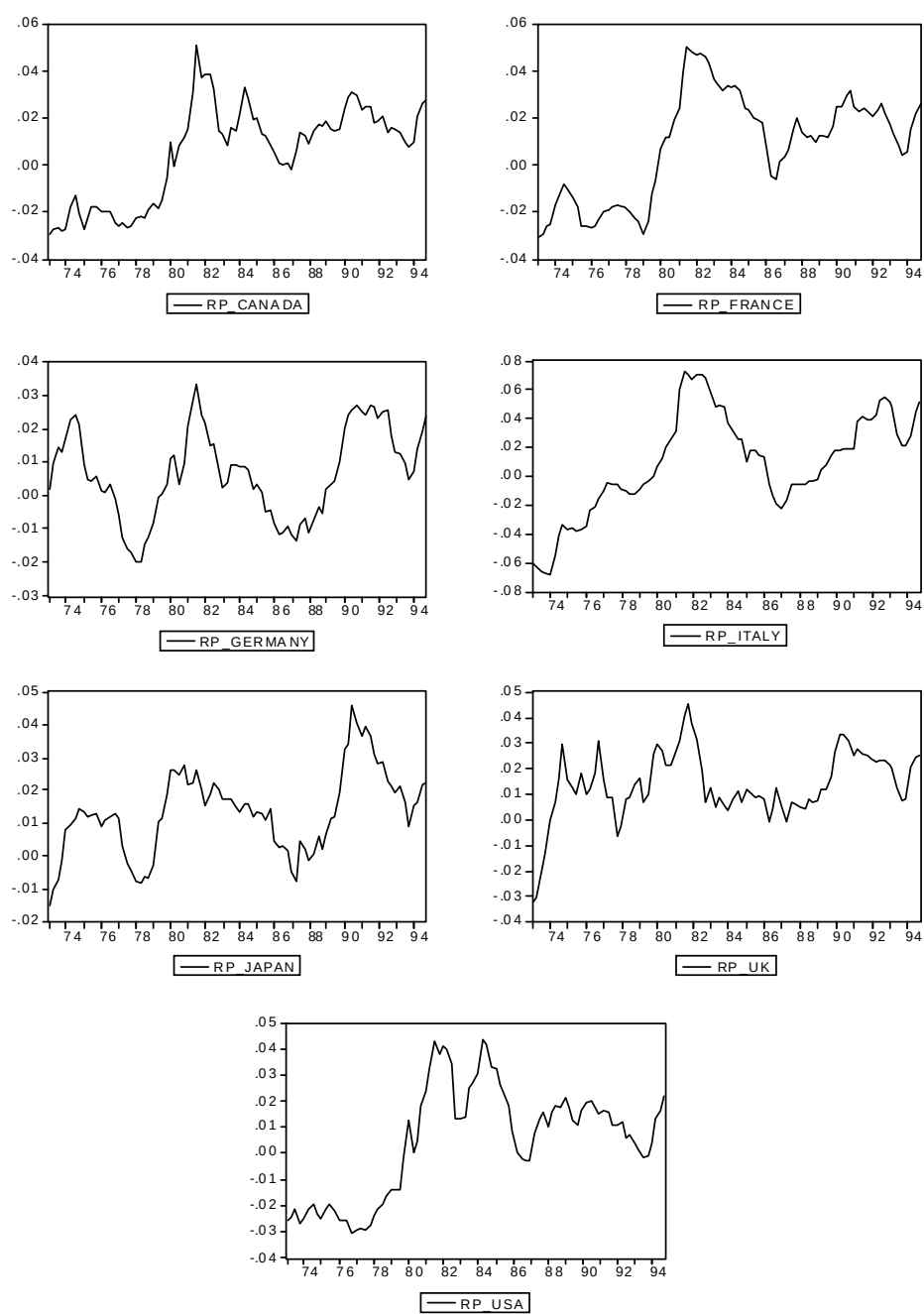


Figure I. Estimated average long-term risk premia

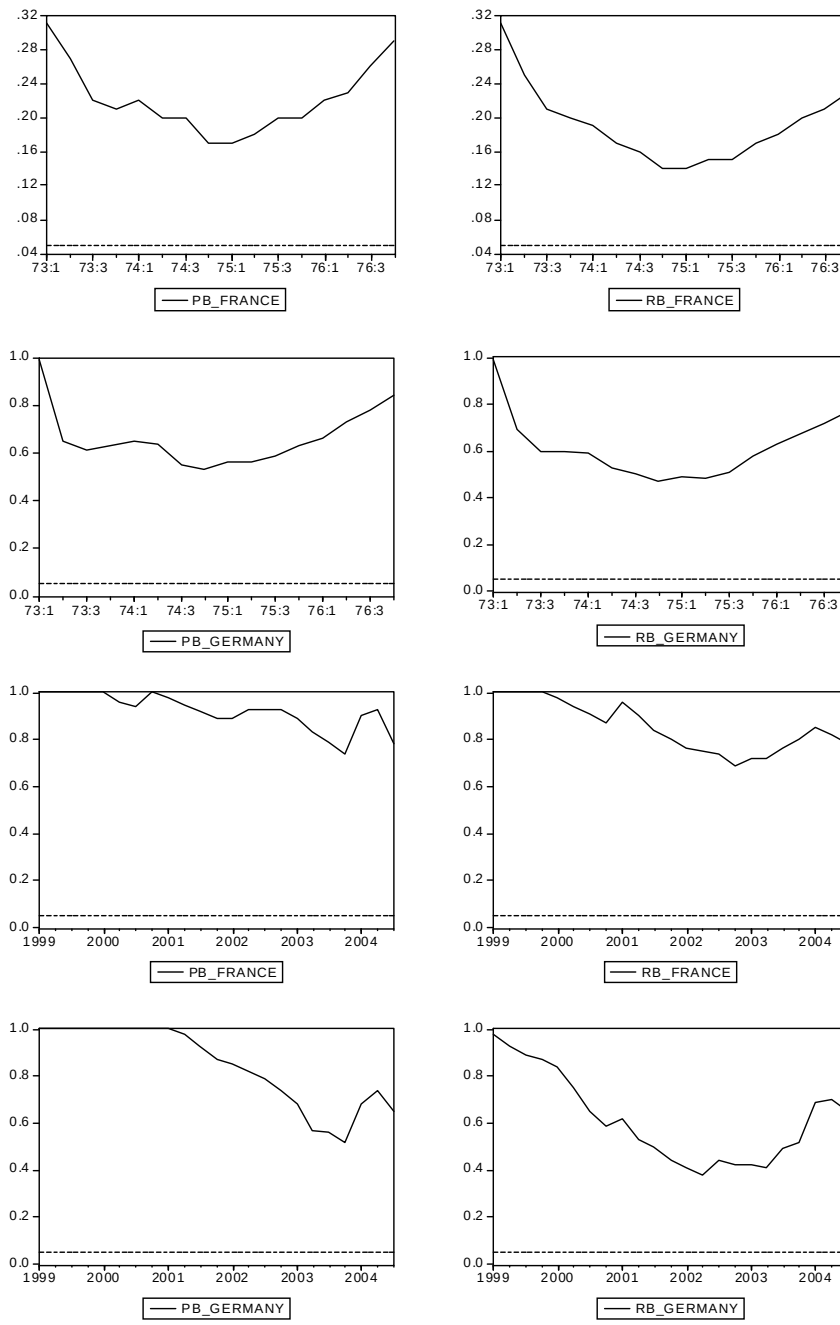


Figure II. p-values Andrews and Kim (2003) cointegration breakdown tests