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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. Moreover, the consideration of structural breaks becomes a key element for this relationship to hold. These shifts are identified as sharp changes in the time-varying risk premium as a consequence of turning points in monetary policy and exchange rates regimes. Finally, the robustness of the obtained results to recent developments in the Eurozone is checked.

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

JEL: C1, E4, E5, F3.

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1 Introduction

Uncovered interest parity (UIP) is a cornerstone of modern models of international finance and exchange rate determination. It establishes a relationship between interest rate differentials of two countries and expected currency depreciations. From this point of view, changes in interest rate differentials should be a good predictor of the variations in nominal exchange rates. In spite of its theoretical importance, it is well known that empirical failure has usually been found when tested with real data. Surveys on the issue can be found in Froot and Thaler (1990), Engel (1996) and Chinn (2006). Even more, the empirical evidence is counterintuitive in the sense that some studies find that nominal exchange rates exhibit a negative correlation with nominal interest differentials, especially when the analysis is based on the use of short periods of time.

Due to this lack of evidence in favour of the UIP relationship, recent papers on this issue have studied it by adopting a long-horizon perspective. A reasonable argument for using this long-horizon strategy can be found in Campbell et al (1997), where it is demonstrated that long-horizon regressions generally find favorable and significant results where short-horizon regressions find none. Various examples can illustrate this: 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 deal with the relationship between the deviation from their fundamental value and their returns in the case of long-horizon exchange rates.

In this regard, we consider the contribution of Valkanov (2003) as being of the highest interest since it is shown there that the use of long-horizon regressions may lead to inconsistent estimators, incorrect tests and coefficients of determination that do not converge in probability to one. This author concludes that the variables included in this long-horizons regressions usually exhibit a non-stationary pattern of behavior and, therefore, economists may find spurious results, an issue very well studied in Granger and Newbold (1974), Phillips (1986, 1991) or Ferson et al (2003).

In our view, the results reflected in Valkanov (2003) have not yet been appropriately

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taken into consideration so far as the long-horizon UIP analysis is concerned. In particular, we think that most earlier papers on this issue have not considered the time series properties of the variables included in the long-horizon UIP relationship. The possible non-stationarity of these variables is very important in the sense that it determines the most adequate method to estimate the UIP relationship. If we can show that the variables are better characterized as being I(1) rather than I(0), then the use of cointegration techniques is more appropriate than the employment of OLS or GMM estimators, as applied in Chinn and Meredith (2004) or Zhang (2005), respectively.

However, we recognize that the mere use of standard cointegration techniques may not be enough in order to appropriately explain the UIP hypothesis and, consequently, some additional sources of misspecification in the UIP relationship should be also considered. Given that the empirical failure of the UIP has been related to the existence of a time-varying risk premium, this paper also explores whether the presence of a changing risk premium can help us to explain the puzzle about the lack of evidence in favour of the UIP. If our hypothesis is true and the changes in the risk premium play an important role in this scenario, then we should expect the standard cointegration statistics not to be useful. By contrast, some new statistics that test for the no-cointegration null hypothesis in the presence of changes in the intercept of the model may be necessary. This task has been partially carried out in Gregory and Hansen (1996), where the methodology of Engle and Granger (1987) is extended to the case of a single shift in the parameters of the model. If this single change is not enough to capture the evolution of the risk premium, the extension of these statistics to the case of multiple breaks will be carried out.

The rest of the paper is organized as follows. Section 2 defines the concept of UIP and the methods that have been commonly employed for testing this hypothesis. Special emphasis is made on long-horizon UIP and the problems inherent to the regressions using variables measured over wide time periods. Section 3 describes the data used and tests for the integration order of the long-horizon variables involved. This section further analyzes long-horizon UIP by way of the use of cointegration methods that allow for the presence of multiple structural breaks, whose existence is clearly justified by a time-varying risk premium on long-term gov-
ernment bond yields. As these changes are related to variations in the monetary policy, it seems sensible to study whether the recent adoption of an European common monetary policy has modified the parameters of the long-horizon UIP. This analysis is the aim of section 4. Section 5 closes the paper with a review of the most important conclusions.

2 Long-horizon UIP. Background and empirical evidence

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

\[ \Delta s_{t+k} = \alpha + \beta (i_{t,k} - i^*_t) + \eta_t \]  

where \( \Delta s_{t+k} \) 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 \((i_{t,k} - i^*_t)\) is the k-period national interest rate less the corresponding foreign one. \( \eta_t \) is an error term.

Testing for the null hypothesis that UIP holds is equivalent to testing for the null hypothesis that the slope parameter in (1) is equal to 1. We can alternatively consider the joint null hypothesis that the intercept of the model is equal to 0 and the slope parameter is 1. However, if the assumption of risk-neutral investors is relaxed, this last parameter may reflect the (time-invariant) risk premium demanded on foreign assets.

An alternative specification of equation (1) is obtained when the interest rate differential is replaced by the ‘forward discount’ (difference between the forward and spot exchange rates). Based on arbitrage arguments, and under the same assumptions as before, this variable can be interpreted as an unbiased estimator of the subsequent exchange rate depreciation. A very large literature has tested the ‘unbiasedness hypothesis’, finding that the coefficient \( \beta \) is considerably lower than 1. In the survey carried out by Froot and Thaler (1990) an average coefficient across 75 studies equal to \(-0.88\) is reported\(^2\). This difference is difficult to explain,

\(^2\)McCallum (1994) argues that the UIP relationship is distinct from, and more important than, the unbi-
but some common interpretations are those related to the possible existence of a time-varying risk premium, heterogeneity or errors in expectations and ‘peso’ problems. Engel (1996) presents another survey with similar contents and interpretations and Chinn (2006) is the most recent review of the literature.

Recent studies are more optimistic about the empirical performance of UIP. Baillie and Bollerslev (2000), for example, argue that the rejection of the UIP hypothesis is a statistical anomaly as a consequence of the persistent ‘forward premium’ and the small sample sizes used. Bekaert and Hodrick (2001) focus also on the finite sample properties of the tests. Chaboud and Wright (2005), using intradaily data, find that UIP works better in the very short run, where the risk premium shrinks to zero. Finally, 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 we will focus in this paper on those that analyze it over long time horizons. Flood and Taylor (1997) used IFS medium-term government bonds and found a coefficient for the interest differential of 0.596 (standard error 0.195), rejecting both the 0 and 1 null hypothesis for the slope parameter. Alexius (2001) worked with quarterly IFS long-term interest rates from 1957:Q1 to 1997:Q4 and, having dealt with the possible problems inherent to the data, found substantial evidence supporting the unbiasedness hypothesis. Chinn and Meredith (2004) criticize the fact that the sample used in Alexius (2001) covers periods of fixed exchange rates and extensive controls and consider 10-year government bonds and ‘constant maturity’ 5 and 10-year yields during 1973:Q1-2000:Q4. They find more support for the UIP and give an interpretation from a macroeconomic point of view extending the model in McCallum (1994). Zhang (2005) makes a similar analysis for both monthly and quarterly data and links the favorable evidence to the low and medium frequency bands of the data instead of the horizon. In all these studies no distinction about the evidence among countries is made.

Up to date, the empirical success of UIP over wide time periods can be related to the asedness of forward exchange rates as predictors of future spot rates. This means that the existence of negative estimations of the slope parameter implies the rejection of the ‘unbiasedness hypothesis’ but not necessarily UIP.

3Following Baxter (1994) these 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.
problems that practitioners face when working with variables referred to 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 hypothesis of no predictability. They pointed that estimated slope coefficients, t-statistics and coefficients of determination increased with the time period over which returns were calculated. Since then, several studies have appeared, such as Valkanov (2003), Mark and Sul (2004) and the references therein, seeking to explain why this happens and proposing adequate techniques given the characteristics usually displayed by long-horizon variables.

Using asymptotic arguments and a ‘local-to-unity’ framework, Valkanov (2003) explains the reasons for long-horizon regressions’ tendency to find significant and favorable results where short-term approaches find none. The explanation relies on the persistent and possibly non-stationary behavior of the variables involved. It is found that inconsistent estimators, incorrect tests and coefficients of determination that do not converge to one in probability can be obtained. These facts are related to the chance of finding a spurious regression between two persistent variables, a phenomenon highlighted by Granger and Newbold (1974), Phillips (1986, 1991) and Ferson et al (2003).

Therefore, estimation and testing using long-horizon variables cannot be carried out applying ‘conventional’ regression methods, such as Ordinary Least Squares, given that spurious results can be found. Previous papers that are based on the use of long-horizon UIP tests can be affected by this problem. Zhang (2005) 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, concluding that these variables behave as non-stationary. In spite of this result, the UIP relationship is estimated by way of the Generalized Method of Moments (GMM) due to the possible correlation of the error term with the explanatory variable. However, we should take into account that the GMM estimation does not solve the problems mentioned above, in the sense that it ignores the presence of unit roots in the variables, whilst the application of cointegration techniques seems to be more appropriate in this scenario.

Against this background, the aim of the next section is, first, to analyze whether the long-horizon variables involved in UIP testing for the G7 countries during the post-Bretton Woods
era can be characterized as I(1) variables. If we cannot provide evidence against the unit root hypothesis, then we should conclude that the cointegration analysis is an appropriate methodology to study the UIP hypothesis although, it is possible that we should admit the presence of changes in the parameters of this relationship that can reflect the variations on risk premiums.

3 Data description and analysis

As pointed before, the aim of this paper is to analyze whether there is evidence in favour of the long-horizon UIP hypothesis for the G7 countries. The sample size covers the period 1973:Q1 to 2004:Q3 and exchange rates and long-term (around 10-years) government bond yields have been obtained from the IMF International Financial Statistics CD-ROM. In line with the specification stated in (1), the endogenous variable is the annualized\(^4\) change in the logarithm of the exchange rate (national currency in terms of U.S. Dollars) over 40 periods (10 years), whilst we use as explanatory variable the difference between the national and the U.S. yields (both expressed in percent per annum) for a 40-period horizon at time \(t\).

Following the results reported in Valkanov (2003), this long-horizon analysis must take into account that the variables may exhibit a great amount of persistence. Consequently, we suspect that the variables included to be better characterized as being I(1) than I(0). Since the appropriate consideration of the integration order of the variables is the key point of our proposal, we first test for the null hypothesis for the two variables considered. We dispose of a great range of statistics devoted to that end. A first option is the use of the Augmented Dickey-Fuller tests (Dickey and Fuller (1979) and Said and Dickey (1984)) or those presented in Phillips and Perron (1988). However, we should take into consideration the recent paper of Ng and Perron (2001), which compares the performance of a wide range of unit root statistics. On the basis of their results, these authors proposes the ADF\(^{GLS}\) statistic, which is based on the very popular ADF test. It can be obtained from the estimation of the following model:

\(^4\)Japanese data needs interpolation for the long-term depreciation in 1985:1 and 1985:2 as a consequence of this annualization.
\[ y_t = \delta_t + \rho y_{t-1} + \sum_{i=1}^{\ell} \phi_i \Delta y_{t-i} + \varepsilon_t \]  

(2)

where \( y_t \) is the variable of interest, \( \delta_t \) reflects the deterministic elements\(^5\), and subsequently calculating the pseudo \( t \)-ratio for testing whether the parameter \( \rho \) is 1. The differences between this and the simple ADF lie in the fact that \( \text{ADF}^{\text{GDS}} \) is based on the use of GLS detrending methods and on the determination of the value of the lag truncation parameter (\( \ell \)) via the use of an information criterion, called MIC, also proposed in Ng and Perron (2001). These authors also propose extensions of the Phillips-Perron statistics, which are based on an appropriate selection of both the lag truncation parameter and the estimation of the deterministic elements, as well as on the use of the autoregressive spectral density estimator at frequency zero. These statistics are commonly known as \( MZ_t \) and \( MZ_a \). The three statistics have a good performance, in the sense of showing an adequate empirical size and good power, so we will report the values of the \( \text{ADF}^{\text{GDS}} \), \( MZ_t \) and \( MZ_a \) for testing the unit root null hypothesis.

Table 1 presents the resulting statistics from the application of the above-mentioned unit root tests. They confirm our initial hypothesis and the presence of a unit root cannot be rejected for the annualized change in the logarithm of the exchange rate, even when the liberal 10% significance level is used. The results are slightly less robust for the interest rate differentials, in the sense that we can offer somewhat limited evidence against the unit root null hypothesis, although only for the cases of Canada and France. In any event, this evidence is so weak that we can admit that these two interest rate differentials show, at least, a great amount of persistence and we consider to be more sensible to treat these two variables as being first order integrated rather than being I(0). Consequently, and in line with our initial point of view, the long-horizon UIP should be analyzed by applying those techniques that admit the presence of integrated variables, as is the case of the cointegration methods, rather than by those techniques based on the use of non-integrated variables, as most earlier papers do.

Table 2 reports the results obtained when the cointegration analysis is applied. To that

\(^5\)We will always include in our specification an intercept and a deterministic trend.
end, and following Engle and Granger (1987), we have estimated by OLS the following model:

$$\Delta s_{t+40} = \alpha + \beta (i_{t,40} - i^{*}_{t,40}) + \eta_t, \quad t = 1, ..., T$$  \hspace{1cm} (3)

and then we have tested for the presence of a unit root in the perturbations by way of the standard ADF statistic\(^\text{6}\). The results of Table 2 show that the evidence against the no-cointegration null hypothesis is very weak and we can only admit the possible existence of a long-horizon UIP relationship for the United Kingdom. Consequently, we conclude that the evidence in favour of the UIP relationship is quite limited, contrary what the economic theory states. Moreover, this also leads us to admit that financial markets are not related, what is counterintuitive if we analyze the real performance of these markets where we can observe a great degree of correlation. Thus, we should look for some explanations for this lack of evidence in favour of the UIP relationship.

3.1 Estimating the time-varying risk premium of long-term bonds

The presence of a time-varying risk premium is one of the most common explanations for the empirical rejection of the UIP hypothesis. We should take into account that the existence of this variation in the risk premium may be associated with changes in the intercept of equation (1) and it is well known nowadays that the no consideration of changes in the parameters of the model causes the no-cointegration tests to be biased towards the non rejection of their null hypothesis\(^\text{7}\). Thus, it seems to be sensible to explore whether changes in the risk premium may influence the analysis of the long-horizon UIP.

Therefore, we need an estimation of the risk premia of long-term interest rates for each country. To that end, we use the Rational Expectations Hypothesis of the Term Structure (REHTS), as in Crespo-Cuaresma et al (2005), which allows us to determine the risk premium perceived by the agents on long-term investments and whether the assumption of a constant one is reasonable for the sample period analyzed.

The REHTS states that the yield to maturity of a \(k\)-period bond is equal to the sum of

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\(^6\)The use of the Z-type tests proposed in Phillips and Ouliaris (1990) or the most recent statistics proposed in Perron and Rodriguez (2001) does not alter the conclusions reported here.

\(^7\)See Perron (1989) or Gregory and Hansen (1996).
the composition of the one-period yields plus the risk premium, which can be represented as follows:

\[ i_{t,40} = \frac{1}{40} \sum_{j=0}^{39} E_t(i_{t+j,1}) + \phi(40,t) \]  

(4)

where \( E_t(\cdot) \) is the conditional expectation operator using the information available at period \( t \), \( i_{t+j,1} \) is the 3-months interest rate \( j \) periods ahead (obtained from the OECD Main Economic Indicators), \( i_{t,40} \) is the yield to maturity of a government bond (10 years) and \( \phi(40,t) \) is the average risk premium until it matures. Note that we are still considering the existence of rational expectations and, therefore, we assume that agents will have perfect foresight with respect to the one-period interest rates and realized values will be used.

The resulting long-term government bond yields risk premiums are plotted in Figure 1. It may immediately be observed that they follow a time-varying pattern of behavior. More interestingly, we can also see that the most abrupt changes take place around similar dates. For almost all series, there are two large positive shifts from zero: the first change appears at the beginning of the 1980s, whilst the second occurs in the second part of that decade, with this pattern of behavior being especially clear for the case of the USA.

It is very interesting to note that the risk premium for the United Kingdom does not exhibit a great amount of variation or, at least, this variation is much lower than the observed for the rest of the countries. This fact could help us to explain why this is only country for which cointegration has been found.

Thus, this simple analysis confirms our initial intuition that the risk premium can exhibit a time-varying pattern of behavior. Given that the intercept of the UIP relationship is clearly related to it, then it seems to be sensible to admit that this parameter may vary across the time, with this suggesting the use of those statistics that test for the non-cointegration null hypothesis in the presence of changes in the parameters of the model specification. This is the aim of the next subsection.

### 3.2 Cointegration and the presence of structural breaks

As we have already mentioned, the intercept in equation (3) reflects the risk premium demanded on foreign assets. We have offered evidence in favour of this variable to exhibit a
time-varying pattern of behavior. In these circumstances, we cannot consider the standard residual-based statistics to be appropriate to test the no-cointegration null hypothesis and, therefore, it is necessary to adapt these statistics to the presence of changes in the parameters of the model. In this regard, we should take into account the proposal of Gregory and Hansen (1996) where the Engle-Granger methodology is extended to that case where a single structural break is present in the model specification. These authors develop some new statistics that allow for a level or regime shift at an unknown date. These changes are modeled by the inclusion of a dummy variable in the model:

\[
\varphi_{t\tau} = \begin{cases} 
0 & \text{if } t \leq [\tau \cdot T] \\
1 & \text{if } t > [\tau \cdot T] 
\end{cases}
\] (5)

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

If we focus on that case where the break exclusively affect the intercept then model (3) can be respecified as:

\[
\Delta s_{t+40} = \alpha + \alpha_1 \cdot \varphi_{t\tau} + \beta (i_{t,40} - i^*_{t,40}) + \varepsilon_t, \quad t = 1, \ldots, T
\] (6)

The Gregory-Hansen method to test for the non-cointegration null hypothesis in the presence of a break is also based on the analysis of the residuals. First, the candidate cointegration relationship is estimated via OLS and, later, a unit root test is applied to the residuals. This is implemented for all possible dates where the break can occur (imposing the restriction that \( \tau \epsilon (0.15, 0.85) \), for example) and the smallest value of the resulting statistics is considered. The breakpoint is considered to be that date for which the unit root statistic is minimum:

\[
ADF_1^{1} = \inf_{\tau \epsilon (0.15, 0.85)} ADF(\tau)
\] (7)

\[
\hat{\tau} = \arg \min_{\tau \epsilon (0.15, 0.85)} ADF(\tau)
\] (8)

However, the results of the preceding subsection suggest the presence of more than one change in the risk premiums. Thus, we need to adapt the Gregory-Hansen method to the
presence of a second break. To that end, let us consider that the two level shifts occur at the unknown dates $\tau_1 \cdot T$ and $\tau_2 \cdot T$ and create the following dummy variables

$$\varphi^1_{tT} = \begin{cases} 0 & \text{if } t \leq [\tau_1 \cdot T] \\ 1 & \text{if } t > [\tau_1 \cdot T] \end{cases}$$

$$\varphi^2_{tT} = \begin{cases} 0 & \text{if } t \leq [\tau_2 \cdot T] \\ 1 & \text{if } t > [\tau_2 \cdot T] \end{cases}$$

The Gregory-Hansen equation should be consequently transformed as follows:

$$\Delta s_{t+40} = \alpha + \alpha_1 \cdot \varphi^1_{tT} + \alpha_2 \cdot \varphi^2_{tT} + \beta(i_{t,40} - i^*_{t,40}) + \xi_t, \quad t = 1, \ldots, T$$

and, then, the new test statistic is defined:

$$ADF^2 = \inf_{\tau_1, \tau_2 \in (0.15, 0.85), \tau_1 < \tau_2} ADF(\tau_1, \tau_2)$$

$$\hat{\tau}_1, \hat{\tau}_2 = \arg \min_{\tau_1, \tau_2 \in (0.15, 0.85), \tau_1 < \tau_2} ADF(\tau_1, \tau_2)$$

If simply follow the Theorem stated in page 109 of Gregory and Hansen (1996), then the limiting distribution of this new test is given by:

$$ADF^2 \overset{d}{\longrightarrow} \inf_{\tau_1, \tau_2 \in (0.15, 0.85)} \frac{\int_0^1 W_\tau dW_\tau}{\int_0^1 W^2_\tau}$$

where

$$W_\tau(r) = W_1(r) - \int_0^1 W_1 W^\top_2 r \left[ \int_0^1 W_2 W^\top_2 r \right]^{-1} W_2(r)$$

and

$$W_2(r) = [1, \varphi^1_{1r}, \varphi^2_{1r}, W_2^\top(r)]^\top$$

Following Gregory and Hansen (1996), we have obtained the critical values via Monte Carlo simulations. Using 10,000 replications, the asymptotic critical value for a 5% nominal significance level is $-5.44$ when the model specification includes a single regressor. This is the
value that we will use in our empirical application.

Once we have statistics that can test for the no-cointegration null hypothesis in the presence of changes in the parameters of the model, we can apply them to analyze whether the long-horizon UIP holds. The statistics obtained are reported in Table 2. As we can see, we can now reject the no-cointegration null hypothesis for the cases of France and Germany, which confirms the importance of considering the presence of changes in the intercept. Moreover, the estimated break dates in both cases are clearly related: the first change is located at 1981:1 in France and 1983:2 in Germany, whilst the second appears in 1987:2 and 1986:4, respectively. Then, if we take into account that the most abrupt changes in the risk premiums obtained take place around these periods of time, it is quite straightforward to connect the breaks estimated to those in the risk premiums. Given that we are considering the exchange rate of each national currency against the US Dollar and the interest rates differentials with respect to the US interest rate, we can interpret this result through the changes in the monetary policy of the Federal Reserve and the Dollar evolution that occurred at those periods of time. We should recall that the economic decisions taken in these periods of time also play an important role in, for example, the explanation of the deviations of the long-run PPP behavior, as is shown in Gadea et al (2004).

The location of the periods where the shifts occur for the models of Japan, Italy and Canada are clearly related to those for France and Germany. We cannot admit the presence of a UIP relationship in these cases but, at least, we can observe some degree of connection between the international financial markets in the sense that changes in the risk premium are affecting to all of them in a similar way.

3.3 Total versus Partial UIP

The above results allow us to conclude that there is no evidence in favour of a long-horizon UIP relationship for Canada, Italy and Japan. By contrast, we can find it for the United Kingdom, Germany and France, in these latter cases when two shifts in the intercept of the

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8 This Table reports the results of the case where the intercept of the model exhibits two changes. We have also considered the introduction of an additional third level or the consideration of changes in the slope of the model. However the results obtained do not modify those reported in this Table.
model are included. However, this does not imply that variations in interest differentials are totally transmitted to the exchange rate, as UIP states. By contrast, we could observe the existence of a partial UIP effect if the parameter $\beta < 1$ in equation (1), instead of the presence of a total effect, represented by that case where $\beta = 1$. To solve this question, we should test for the $\beta = 1$ null hypothesis in those cases where the non-cointegration null hypothesis is rejected.

We may note here that standard t-statistics are not appropriate to this aim; rather, we employ alternative methods that allow inferences about the parameters of the model in a cointegrated relationship. We have chosen the ‘Fully-Modified’ OLS estimation proposed by Phillips and Hansen (1990). Results are those in Table 3 which allow us to conclude that, first, we cannot reject that $\beta = 1$, what can be interpreted as evidence in favour of a total long-horizon UIP effect for these countries. We also see that the changes in the intercept of the models of Germany and France reject the non-significance null hypothesis, confirming their importance to explain the relationship between the evolution of exchange rates and interest rate differentials. Moreover, the estimations of these effects are similar in magnitude, although they have opposite signs: the shift that appears at the beginning of the 1980’s implies a decrease in the intercept, whilst that occurring in the mid 1980’s involves an increase in this value. These breaks reflect the variation in the risk premium demanded by French and German agents on home assets over U.S. investments.

4 Have the European Common Monetary Policy affected the long-horizon UIP relationship?

As can be observed in (7) and (11), the method in Gregory and Hansen (1996) and its extension as proposed here keep some observations without being breakpoint candidates, both at the beginning and the end of the sample. This restriction is important in our cases given that the two countries for which the long-horizon UIP holds with a double shift in the constant are France and Germany. We have also observed that the estimation of the periods where the breaks may have occurred seems to be closely in connexion with changes in monetary policy. An important event related to these aspects has taken place at the end of the sample considered in
these countries: the introduction of the common monetary policy and the common currency in the European Monetary Union. For this reason, it will be of interest to check the robustness of our results and explore whether there has been any change in the relationships due to the creation of the Euro.

To that end, we can use the methodology recently developed in Andrews and Kim (2003). This method detects possible cointegration breakdowns over short periods of time (small number of observations). The types of breaks it considers are both a change in the cointegration vector and a shift in the error distribution from \(I(0)\) to \(I(1)\). The test statistics are constructed from the following model:

\[
y_t = \begin{cases} 
  x_t'\beta_0 + u_t & \text{for } t = 1, \ldots, T^* \\
  x_t'\beta_1 + u_t & \text{for } t = T^* + 1, \ldots, T^* + m 
\end{cases}
\]  

(16)

where \(y_t\) is the endogenous variable \((\Delta s_{t+40})\), \(x_t\) are the explanatory variables \((\varphi_{t^1}, \varphi_{t^2}\) and \((i_{t,40} - i_{t,40}^*)\)) and \(u_t\) denotes the error term. \(\beta_0\) and \(\beta_1\) are the parameter vectors. \(y_t, u_t \in R\) and \(x_t, \beta_0, \beta_1 \in R^3\). \(T^*\) is the breakpoint and \(T = T^* + m\) is the total number of observations.

The null and alternative hypotheses can be stated as:

\[
H_0 : \begin{cases} 
  \beta_1 = \beta_0 & \text{for all } t = T^* + 1, \ldots, T^* + m \text{ and} \\
  \{u_t : t = 1, \ldots, T^* + m\} \text{ are stationary and ergodic} 
\end{cases}
\]

(17)

\[
H_1 : \begin{cases} 
  \beta_1 \neq \beta_0 & \text{for some } t = T^* + 1, \ldots, T^* + m \text{ and/or} \\
  \text{the distribution of } \{u_{T^*+1}, \ldots, u_{T^*+m}\} \text{ differs from} \\
  \text{the distribution of } \{u_1, \ldots, u_{T^*}\} 
\end{cases}
\]

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

\[
P_b = \sum_{t=T^*+1}^{T^*+m} (y_t - x_t'\hat{\beta}_{(T^*+\lfloor m/2 \rfloor)})^2
\]

(18)

where \(\lfloor m/2 \rfloor\) denotes the smallest integer that is greater than or equal to \(m/2\) and
\( \hat{\beta}_{1-(T^*+[m/2])} \) is the OLS estimator obtained using the observations \( t = 1, \ldots, T^* + [m/2] \).

The null hypothesis is rejected whenever the test statistic exceeds a critical value determined using a parametric subsampling method. Consider \( \{P_{b,j} = P_j(\hat{\beta}_j) : j = 1, \ldots, T^* - m + 1\} \), where \( P_{b,j} \) is defined as:

\[
P_{b,j} = \sum_{t=j}^{T^*+m} (y_t - x_t'\hat{\beta}(j))^2
\]

(19)

with \( \hat{\beta}(j) \) being the estimator of \( \beta \) using \( t = 1, \ldots, T^* \) with \( t \neq j, \ldots, j + m - 1 \).

The empirical distribution function of the statistic is given by the following equation:

\[
\hat{F}_{P_{b,T^*}}(x) = \frac{1}{T^* - m + 1} \sum_{t=1}^{T^* - m + 1} 1(P_{b,j} \leq x)
\]

(20)

and the critical value at the \( \alpha \) significance level is equal to the \( 1 - \alpha \) sample quantile

\[
\hat{q}_{P_{b,1-\alpha}} = \inf \left\{ x \in R : \hat{F}_{P_{b,T^*}}(x) \geq 1 - \alpha \right\}
\]

(21)

In order to test for a change in the error distribution from a stationary process to an integrated one in the last \( m \) observations, a locally best invariant test is used which has the same spirit as (18) but consists of a sum of squares of reverse partial sum of residuals after the breakpoint:

\[
R_b = \sum_{t=T^*+1}^{T^*+m} \left( \sum_{s=t}^{T^*+m} (y_s - x_s'\hat{\beta}_{1-(T^*+[m/2])}) \right)^2
\]

(22)

The test is implemented in the same way to the procedure previously stated.

We have considered as possible break dates not only those ignored by the Gregory-Hansen method (which include the observations corresponding to 2001:2 onwards) but also all observations whose variables have been constructed using data after 1999:1, when the European common monetary policy began to act.

The results that we have obtained are presented in the four graphs reported at the bottom of Figure 2. 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, the graph reflects 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 preceding observations. The null hypothesis of no change in the parameters cannot be rejected for any of the four cases, even when the liberal 10% significance level is employed. Therefore, it can be concluded that the adoption of a European common monetary policy has not altered the long-horizon UIP relationships.

We should finally note that we have also carried out the same exercise for the observations trimmed at the beginning of the sample. P–values are those at the top of Figure 2, from which similar conclusions with respect to the stability in the cointegration relationship are obtained.

5 Conclusions

Existing studies testing for the uncovered interest parity over wide periods of time have not taken into account the (possibly) non-stationary behavior of the variables included in the model specification. If we follow Valkanov (2003) it seems clear that the variables commonly employed in long-run regressions may exhibit a great amount of persistence and, therefore, they can be better characterized as being I(1) than I(0).

Adopting this result as the keypoint of our paper, we have studied long-horizon UIP for the economy of the G6 countries versus the US economy. The sample covers the post Bretton-Woods era. We have first tested the integration order of the variables included in the UIP relationship and we cannot offer strong evidence against the presence of a unit root, which confirms that the long-run analysis requires the use of those econometric techniques that reflect the fact that the variables are not stationary, as is the case of the cointegration methods.

The application of the standard statistics for testing the no-cointegration null hypothesis only offers evidence against it for the case of the United Kingdom. However, this lack of evidence may be simply the consequence of a misspecification problem, as the no consideration of shifts in the intercept of the model could be. This parameter reflects the risk premium and we have offered evidence in this paper that this variable exhibits a time-varying pattern of behavior across the sample size used. Then, if we extend the statistics proposed in Gregory and Hansen (1996) to those cases where the cointegration relationship may show two changes in
the intercept, the no-cointegration null hypothesis can be rejected for the cases of France and Germany. The estimation of the periods where these changes appears is coincident with those when variations of the risk premium are of greater magnitude. Therefore, we can conclude that the evolution of the risk premium can help us to explain the UIP. Moreover, these shifts are clearly related to changes in monetary policies, a result that can also help us explain the PPP relationship between the US currency and the German and French currencies. Once we have determined the correct specification of the model, we can then test for the existence of a total transmission of the interest differentials to exchange rates. We cannot reject the $\beta = 1$ null hypothesis and, consequently, we conclude in favour of the existence of a total UIP for the United Kingdom, France and Germany.

Finally, we have analyzed whether the adoption of the European common monetary policy has altered the long-horizon UIP relationship. Our analyses lead us to admit that this particular event do not seem to have involved any modification.

References


Table 1. Unit root tests for long-horizon (10-years) exchange rate depreciation and interest rate differentials. 1973:Q1-2004:Q3

<table>
<thead>
<tr>
<th>Country</th>
<th>Canada</th>
<th>France</th>
<th>Germany</th>
<th>Italy</th>
<th>Japan</th>
<th>United Kingdom</th>
</tr>
</thead>
<tbody>
<tr>
<td>Variable</td>
<td>$\Delta^{40}s_t$</td>
<td>$i_{t,40}^<em>-i_{t,40}^</em>$</td>
<td>$\Delta^{40}s_t$</td>
<td>$i_{t,40}^<em>-i_{t,40}^</em>$</td>
<td>$\Delta^{40}s_t$</td>
<td>$i_{t,40}^<em>-i_{t,40}^</em>$</td>
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<tr>
<td>c.v. (5%) Test</td>
<td>-17.30</td>
<td>-16.69</td>
<td>-3.23</td>
<td>-16.22</td>
<td>-4.53</td>
<td>-2.35</td>
</tr>
<tr>
<td>MZ$_t$</td>
<td>-2.58</td>
<td>-16.69</td>
<td>-3.23</td>
<td>-16.22</td>
<td>-4.53</td>
<td>-2.35</td>
</tr>
<tr>
<td>ADF$^{GLS}$</td>
<td>-1.08</td>
<td>-3.24</td>
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<td>-3.19</td>
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<td>-1.10</td>
</tr>
<tr>
<td></td>
<td>Canada</td>
<td>France</td>
<td>Germany</td>
<td>Italy</td>
<td>Japan</td>
<td>United Kingdom</td>
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<tr>
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<td>---------</td>
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<tr>
<td><strong>Engle and Granger (1987)</strong></td>
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<td></td>
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<tr>
<td>c.v. (5%):−3.17 ADF</td>
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<td></td>
<td>-2.45</td>
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<tr>
<td><strong>Gregory and Hansen (1996)</strong></td>
<td></td>
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</tr>
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<td>1 level shift</td>
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</tr>
<tr>
<td>c.v. (5%):−4.83 ADF(^1)</td>
<td>−3.52</td>
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<td>-3.24</td>
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<td>-3.55</td>
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<td><strong>Extended Gregory and Hansen (1996)</strong></td>
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<td>2 level shifts</td>
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<tr>
<td>c.v. (5%):−5.44 ADF(^2)</td>
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<td>-5.79</td>
<td>-5.50</td>
<td>-4.37</td>
<td>-4.82</td>
<td>-3.81</td>
</tr>
</tbody>
</table>
Table 3. Cointegration relationships estimation. Long-horizon UIP.

Fully modified OLS. 1973:Q1-2004:3

<table>
<thead>
<tr>
<th>Country</th>
<th>t₁</th>
<th>t₂</th>
<th>α</th>
<th>α₁</th>
<th>α₂</th>
<th>β</th>
<th>H₀: β = 1</th>
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</thead>
<tbody>
<tr>
<td>United Kingdom</td>
<td>-0.01</td>
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<td>0.69†</td>
<td></td>
<td></td>
<td>-1.15</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td></td>
<td>(0.27)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>France</td>
<td>1981:1</td>
<td>1987:1</td>
<td>0.02</td>
<td>-0.09†</td>
<td>0.07†</td>
<td>1.34†</td>
<td>0.71</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(0.48)</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Germany</td>
<td>1983:2</td>
<td>1986:3</td>
<td>-0.01</td>
<td>-0.07†</td>
<td>0.09†</td>
<td>0.64†</td>
<td>-1.29</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(0.02)</td>
<td>(0.28)</td>
<td></td>
<td></td>
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</tr>
</tbody>
</table>

Note: This Table reports the estimation of model (1). Standard errors in parentheses. † and ‡ means statistical significance at the 1 and 5 % level.
Figure 1. Estimated average long-term risk premia
Figure 2. p-values Andrews and Kim (2003) cointegration breakdown tests