EUROPEAN CENTRAL BANK

EUROSYSTEM

WORKING PAPER SERIES

NO 1405 / DECEMBER 2011

BOND MARKET CO-MOVEMENTS, EXPECTED INFLATION AND THE EQUILIBRIUM REAL EXCHANGE RATE

by Corrado Macchiarelli

EUROPEAN CENTRAL BANK

EUROSYSTEM

WORKING PAPER SERIES

NO 1405 / DECEMBER 2011

BOND MARKET CO-MOVEMENTS, EXPECTED INFLATION AND THE EQUILIBRIUM REAL EXCHANGE RATE[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.

This paper can be downloaded without charge from http://www.ecb.europa.eu or from the Social Science Research Network electronic library at http://ssrn.com/abstract_id=1965413.

1 The author is grateful to Fabio Bagliano, Claudio Morana for constructive comments and discussions. He is also thankful to Katerina Juselius for useful guidance on an early version of the paper, to Andrew Harvey for comments on the "Modeling Inflation" results, and to Carlo Favero and Andrea Beltratti for further discussions. The paper also benefited from comments provided by Giuseppe Bertola, Guido Ascari and other participants at the Collegio Carlo Alberto 3rd Doctoral Workshop, together with participants at an internal seminar organized by the Monetary Policy Analysis Unit of the Swiss National 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

© European Central Bank, 2011

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

All rights reserved.

Any reproduction, publication and reprint in the form of a different publication, whether printed or produced electronically, in whole or in part, is permitted only with the explicit written authorisation of the ECB or the author.

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

Since the end of the fixed rates in 1973 and after the EMS sterling dismissal in 1992, the value of the pound has undergone large cyclical fluctuations on average. Of particular interest to policy makers is the understanding of whether such movements are consistent with the lack or not of a correction mechanism to some long-run equilibrium. The purpose of the present study is to understand those dynamics, how the external value of the British sterling relative to the USD evolved during the recent floating experiences, and what have been the driving forces. In this paper we assume the real exchange rate to be determined by forces relating to the goods and capital market in a *general equilibrium* framework. This entails testing the purchasing power parity and the uncovered interest parity together. Our findings have two important implications, both for monetary policy. First, we show that some of the observed changes in the real exchange rates can not be solely attributed to changes in inflation rates, but, possibly, also to investors' behavior. Secondly, we show that the special US dollar status of World reserve currency results into a weaker behavior of the US bond rate on international markets.

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

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

Non-Technical Summary

Exchange rates between currencies have been "highly unstable" since the end of Bretton Woods (Krugman, 1993). One of the main arguments in favour of the free float is that, under a flexible regime, nations can pursue independent monetary policy and adjust exchange rates to ease payment imbalances or restore losses in competitiveness (*ibid.*). Since the end of the fixed rates in 1973 and after the EMS sterling dismissal in 1992, the value of the pound has undergone large cyclical fluctuations on average. Of particular interest from a central banking perspective is the understanding of whether such movements are consistent with the existence or not of a correction mechanism to some long-run equilibrium.

The purpose of the present study is primarily to understand those dynamics, how the external value of the British sterling relative to the USD evolved during the recent floating experiences, and what have been the driving forces. In this paper we assume the real exchange rate (RER) to be determined by forces relating to the goods and capital market in a *general equilibrium* framework. Moving from the definition of two well known zero arbitrage conditions, the purchasing power parity (PPP) and the uncovered interest parity (UIP), it is particularly assumed the RER observed deviations not to be exclusively explained by unidirectional inefficiencies on the goods markets. Alternatively, we grant the explanation of these deviations to involve real factors acting through the current account (Juselius, 1991; 1992; 1995; Johansen and Juselius, 1992; Pesaran et al., 2000; Cheng, 1999; Throop, 1993; Zhou and Mahadavi, 1996; Hunter, 1991; Macchiarelli, 2011). In fact, as the PPP does not to hold in the short run, the current account equality states that an increase in the domestic demand for goods can be satisfied by boosting imports and hence with a growing deficit on the balance of payment. The latter can be financed by increasing the interest rate so to raise a relative supply of cash balances creating a wedge from one country to the other. Using this approach, we are able to assess whether interest rates, prices and the real exchange rate are consistent with a UIP-PPP combined equilibrium. There are two main reasons for which one would expect PPP and UIP not to hold separately. First, commodity price movements or speculative activities may lead to significant deviations from PPP/UIP. Secondly, price stickiness out of the equilibrium, together with the presence of limits to arbitrage and related market imperfections, may undermine short-term adjustments. As prices in markets for goods tend to adjust to shocks more slowly than prices in markets for assets, PPP deviations are more likely to persist if compared to UIP deviations (Pesaran *et al.*, 2000). Combining them both hence allows for a gradual convergence to PPP as it seems that the "PPP relation needs a lift [...] to become more stable, and this is what the interest rates do" (Johansen, 1992). Since in the literature the aforementioned relation is not found to hold strictly under the standard theoretical assumption of rational expectations, we assess the robustness of our results by introducing inflationary expectations. During the course of the analysis, a role for "credibility" of the US dollar (USD) is further explored in line with the literature (e.g. MacDonald, 1998; Juselius and MacDonald, 2004) spotting the existence of weak form of PPP (UIP) when using as *numeraire* a currency having the special status of World reserve currency.

Overall the results point to a role for the dollar's special status which is shown into a weaker behavior of the US bond rate on international markets. A combined UIP-PPP relation is moreover found to hold as a long run condition, albeit not strictly because of UK misalignments. Looking to the adjustment to the long run equilibrium, goods market adjustment is very slow, whilst a major adjustment occurs on the capital and on the exchange rate markets. Nonetheless, the US bond rate is found to "push" in the opposite direction, possibly reconciling with a failure of the UIP itself (e.g. Bekaert *et al.*, 2007; Macchiarelli, 2011). Finally, looking at how accumulated empirical shocks to each variable affect the others, we find that empirical shocks to the UK and the US long term yields respectively increase expected inflation in the US but not in the UK; consistent, in the former case, with a Fisherian (1907) view of nominal rates. Moreover, the accumulated shocks to the RER significantly affect the expected inflation in the UK but not in the US. Those latter findings signal imperfect price/capital adjustments on the international markets, and may account for the widespread finding of rejection of the UIP/PPP conditions when tested separately.

1 Introduction

Since the end of the fixed rates in 1973 and after the EMS sterling dismissal in 1992, the value of the pound has undergone large cyclical fluctuations on average. Of particular interest to policy makers is the understanding of whether such movements are consistent with the presence or not of a correction mechanism to some long-run equilibrium.

The purpose of the present study is primarily to understand those dynamics, how the external value of the British sterling relative to the USD evolved during the recent floating experiences, and what have been the driving forces. In this paper we assume the real exchange rate (RER) to be determined by forces relating to the goods and capital market in a *general equilibrium* framework. Moving from the definition of two well known zero arbitrage conditions, the purchasing power parity (PPP) and the uncovered interest parity (UIP), it is assumed the RER observed deviations not to be exclusively explained by unidirectional inefficiencies on the goods markets (i.e. price stickiness, role of tradables vs. non-tradables goods, non linearities).[1] On the contrary these deviations are expected to involve

[1]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

real factors acting through the current account, as foreign net asset position or fiscal imbalances.[2] As the PPP does not to hold in the short run, the current account equality states that an increase in the domestic demand for goods can be satisfied by boosting imports and hence with a growing deficit on the balance of payment.[3] The latter can be financed by increasing the interest rate so to raise a relative supply of cash balances creating a wedge from one country to the other.[4] Using this approach, we are able to assess whether interest rates, prices and the real exchange rate are consistent with a UIP-PPP combined equilibrium.

Based on previous empirical results, we deepen the evidence in favour of a PPP-UIP joint relation on two main grounds. First, we explore the "credibility" implicit in the special USD status as World reserve currency. Secondly, we assess the validity of previous empirical results by including inflationary expectations, modeled here as long-run inflation forecast.

The reason for focusing on the US vs. UK data owes to the fact that (i) the US is Britain's largest single export market, and (ii) the UK and the US are each other's single largest investor.[5] As discussed before, those are key features if one shares the idea that goods and capital markets may interact in keeping the exchange rate in line.[6]

To preview the results in the paper, we find that the special USD status plays some role, as the US bond displays a weaker behavior on international markets, compared to the UK bond rate. Inference based on a standard cointegration analysis (Johansen 1991; 1994) shows that a combined UIP-PPP relation is found to hold as a long run condition, albeit not strictly because of UK misalignments. Looking to the adjustment to the long run equilibrium, goods market adjustment is found to be very slow, whilst a major adjustment occurs on the capital and on the exchange rate markets. Nonetheless, the US bond rate is found to "push" in the opposite direction, possibly reconciling with a failure of

---

sticky prices in local currency lead to PPP deviations (Engle and Rogers, 1996).

[2] Edison (1987) argues that the failure of all prices to adjust *in unison* may be due to capital movements, changes in the international demand and other structural changes. See also Juselius (1995), Johansen and Juselius (1992), Pesaran *et al.* (2000), Cheng (1999), Macchiarelli (2011).

[3] From an empirical view point valid statistical results were achieved when the PPP was firstly tested as a long run condition. Milestone contributions (Edison, 1987; Lothian and Taylor, 1992; Sarno and Taylor, 2002; Taylor, 2002) found PPP to empirically hold in the long run (for one century data or more) with a half-life of about 4 years for the major industrialized countries. For a detailed overview see Frenkel (1980), Fisher and Park (1991), Froot and Rogoff (1994), Rogoff (1996), Sarno and Taylor (2002).

[4] Zhou and Mahdavi (1996) among the others provide evidence for the real exchange rate in the UK to appreciate against the US dollar as the gap between the UK cumulated current account to output ratio and the same ratio for the US grows. Throop (1993) analogously emphasizes the role of government budget deficits in determining *disequilibria* on the real exchange rate.

[5] Johansen and Juselius (1992) have first applied the testing of international parities condition for the case of UK versus a panel of trade weighted foreign countries. They analogously conclude that the "determination of prices, interest rates and exchange rates should be investigated in a balance of payment framework with interrelated movements in the current account and capital account" (*ibid.*, p.66). For an empirical extension see also Hunter (1992).

[6] Additionally, there is an obvious constraint in considering the same set of parities for, e.g., the euro area vs. the US, given the information on Treasury bond rates for the euro area is clearly missing. Also admitting an analysis of this kind could be replicated, the economic and policy interpretation of the results would be difficult, not only because of the aforementioned asymmetry on the EMU capital markets, but also because many exchange rate movements are related to structural changes, especially at the beginning of our sample. Even if such an analysis could clearly bear on German bonds, this would possibly bias the effect on the existing spot nominal exchange rate. For a similar analysis applied to Germany and the US data see Juselius and MacDonald (2000).

the UIP itself (e.g. Bekaert *et al.*, 2007; Macchiarelli, 2011). Finally, looking at how the accumulated empirical shocks to each variable affect the others, we find that shocks to the UK and the US long term yields respectively increase expected inflation in the US but not in the UK; consistent, in the former case, with a Fisherian (1907) view of nominal rates. Moreover, the accumulated shocks to the RER significantly affect the expected inflation in the UK but not in the US. Those latter findings signal imperfect price/capital adjustments on the international markets, and may account for the widespread finding of rejection of the UIP/PPP conditions when tested separately.

Overall, our analysis have two important implications, both for monetary policy. First, we show that some of the observed changes in the exchange rates can not be solely attributed to changes in inflation rates, but, possibly, also to investors' behavior, overall adjusting to a long run PPP-UIP equilibrium (see also Juselius and MacDonald, 2004). Secondly, we show that the special US dollar status of World reserve currency dampens the appropriate US bond rate's reaction on international markets.

The reminder of the paper is organized as follows. Section 2 presents the theoretical model. Section 3 goes through the econometric strategy and the empirical results. Section 4 is devoted to summary and conclusions.

2 International Parity Conditions

With the purpose of proving co-movements between goods and capital markets, we aim at combining the PPP and the UIP relations. Capital markets are described by the UIP in its standard formulation, stating that for a financial instrument with l periods to maturity to be comparable in a home and a foreign economy it must be:

$$\frac{1}{l}\left(E_t s_{t+l} - s_t\right) = E_t \Delta s_{t+l} = i_t^l - i_t^{l*} - \upsilon_t, \tag{1}$$

where s is the (log) home vs. foreign nominal exchange rate and Δ is the difference operator, i_t^l represents the yield of a bond with maturity l at time t for the home country and υ_t is a time-varying risk-premium.

Equation (1) suggests that the excess of home interest rate over the foreign one (i^{l*}), compounded over l periods, is equal to the expected depreciation of the home currency over the same period (and allowing for a risk premium). Alternatively, one would observe international market arbitrage opportunities.

The UIP describes a general relation. With the aim of considering the special US dollar status, additional comment is required. The dollar is "exceptional" in the sense that it is the main reserve currency Worldwide. The implication is twofold. First, this means agents are "prepared to hold dollars for long periods with little or no change in its relative price" (Juselius and MacDonald, 2000). This is reflected into a feebler interest rate pressure in the US. Secondly, since interest rates in the US do not have to raise as much as in the UK in order to finance a given current account deficit ("USD credibility"), this will result into a weaker pressure over US current account imbalances. The results together imply that the argument according to which the PPP and the UIP can be combined is undermined in this setting.[7]

In order to combine the PPP and the UIP, equation (1) can be re-formulated in terms of the real interest rate $r_{t+l} = i_t^l - E_t \Delta p_{t+l}$ (according to the Fisher's (1907) parity condition, where $E_t \Delta p_{t+l}$ is the expected rate of inflation over l periods at time t in the home country) and the (log) real exchange rate $rer_t = p_t - p_t^* - s_t$ (with p and p^* being the log of the home and the foreign countries overall price level respectively):

$$-\frac{1}{l}\left(E_t rer_{t+l} - rer_t\right) = (i_t^l - i_t^{l*}) - E_t \Delta (p_{t+l} - p_{t+l}^*) - \xi_t, \tag{2}$$

with the error term depending upon the UIP premium (if any), i.e. $\xi_t = \xi(\upsilon_t)$. Following the discussion in Zhou and Madavi (1996) and Throop (1993), we will need assuming the expected value of the real exchange rate over l periods to represent the flexible-price equilibrium value of the real exchange rate. Under this hypothesis, the following long run relation is identified:

$$i_t^l - i_t^{l*} = E_t \Delta \left(p_{t+l} - p_{t+l}^*\right) + rer_t + \xi_t, \tag{3}$$

suggesting that there exists an equilibrium condition towards which the expected inflation, the interest rates and the real exchange rate tend to move in the log run. In other words, for the long-run PPP to hold we would require movements in the RER to be neutralized by movements in the long term interest rates differential, in the (long term) expected inflation differential or a combination of the two, in line with the evidence in Juselius (1995); Juselius and MacDonald (2004) and Sjoo (1992).[8]

As an alternative specification, we test the relation in (3) against a standard real interest parity (RIP) condition, obtained by setting $rer_t = 0$:

$$r_{t+l} - r_{t+l}^* = \xi_t.$$

[7] Under those assumptions the UIP is not expected indeed to hold strictly.

[8] Johansen *et al.* (2007) and Cheng (1999) report for instance the test for the PPP alone - when the relative change in domestic vs. foreign price enters in level - to fail in detecting a valid cointegrating vector, as both the relative price and the nominal exchange rate are "pushing". As a preliminary check for the adequacy of our analysis we tested the null that $(p_t - p_t^*)$ and s_t (with both an unrestricted or restricted constant) cointegrate. The resulting $rank(\Pi)$ is found to be zero with high p-value, i.e. [0.390] using a restricted constant.

Adding and subtracting the term $E_t \Delta s_{t+l}$ in the expression above shows how the RIP can be expressed as a relation conditional on the joint validity of the UIP and the expectational (or *ex ante*) PPP (see Marston, 1997; Macchiarelli, 2011), i.e.

$$r_{t+l} - r^*_{t+l} = \tag{4}$$

$$= \left(i^l_t - i^{l*}_t - E_t \Delta s_{t+l}\right) - \left(E_t \Delta p_{t+l} - E_t \Delta p^*_{t+l} - E_t \Delta s_{t+l}\right) = \xi_t,$$

hence resulting into a formulation capturing the effect of financial markets - to be precise in the nominal interest rates fluctuations - on the expected RER changes.[9]

At this stage it should be noted that the PPP and the UIP are fundamentally different. The PPP is a long run relation whose adjustment is expected to be backward looking, whereas the UIP is forward looking (Mishkin, 1982). The rationale for their combination stands on correctly modeling market expectations (whenever forecast errors are assumed to be systematic). In this respect, Campbell *et al.* (2007) discuss how PPP (and UIP) deviations are due because people generally do not accurately and promptly adjust to changes in the behavior of prices. Hence, if we admit some role for expectations stickiness out of the equilibrium, then the relation in (3) - or its nested version in (4) - would hold in the long run as soon as inflationary expectations, rather than prices, are able to adjust.

3 Methodology and Results

Given the non-stationarity of the variables investigated, current econometric practice suggests to see the relation in (3) as a long run equation with inference carried out with cointegration techniques. In this regard, the existing cointegrating vector is expected to pin down the endogenous variables against speculative activities or commodity price movements leading to short run deviations from the equilibrium.

As a statistical framework for our analysis we refer to the cointegrated VAR (CVAR) model proposed by Johansen (1991; 1994).[10]

For our empirical analysis we use seasonally-adjusted (when needed) monthly data for the US and the UK. We consider three types of data: macroeconomic data (inflation rates), financial data (bond yields), and exchange rate data (bilateral nominal and real exchange rate). In extending our model we input additional macroeconomic variables as each economy's output gap and a World net oil price increase (Hamilton, 1996).

Inflations are measured by annual (log) CPI-based rates. The bond yields are 10-years constant maturity Treasury bond rates, and the rer_t is the (log) RER expressed as $rer_t = p_t - (p^*_t + s_t)$,

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

[10] See also Johansen and Juselius (1990), Pesaran *et al.* (2000), Dennis (2006), Juselius (2006), Lütkepohl (2006).

with s being the (log) sterling/dollar nominal exchange rate. The motivation for focusing on long term interest rates is that those are found to more closely match long run movements in the real exchange rate (Juselius and MacDonald, 2004). GDPs are proxied by Industrial Production Indexes (in logs), whereas output gaps are constructed by applying an HP filter (Hodrick and Prescott, 1997) to GDP, with the smoothing parameter being set to 129.600 following Ravn and Uhlig (2002). Finally, a net oil price increase (NOPI) *à la* Hamilton (1996) is constructed considering the price of oil in the current month oil_t, relative to the maximum value for the level achieved during the previous year, i.e. $NOPI = max\{0, oil_t - max[oil_{t-1} - oil_{t-13}]\}$. In order to get rid of major fluctuations in the global economy the effective sample covers the period 1985-1 to 2008-6. All data sources come from the OECD.stat database.

To assess the degree of persistence of our data, we run a standard ADF unit root test for all the variables entering our benchmark model (Table 1). The test is run for the annualized current home and foreign inflation, the realization of inflation over 10 years, the bilateral real exchange rate and the 10-years constant maturity Treasury bond rate for both the UK and the US. In this paper the US is regarded as foreign economy with all the variables expressed with a star superscript. Additional unit root tests allowing for a structural break (Lumsdaine and Papell, 1997; Perron, 1989; 1997; Zivot and Andrews, 1992) are also run for the UK and the US inflation and their respective bond yields.[11]

The test results suggest both the UK and the US annualized inflations to be non stationary on the overall sample period. The results are robust also to the inclusion of structural breaks (Table 1). Analogously, the real exchange rate and the bond yields are mostly non stationary and sharing the same order of integration. As the realization of inflation over 10 years is simply a moving average of inflation, unit root tests fail in rejecting stationarity at the 5% significance level for both the UK and the US (Figure 1).

3.1 Inflation Measures

On the analytical ground the adoption of a price measure for inflation is itself not without objections. The literature has particularly pointed out several measurement problems.

First, the use of a consumer-price-based index may affect the "true" price measure with a contextual measurement error (MacDonald, 1993; Sarno and Taylor, 2002). In fact, even if we can observe individual goods prices to be the same across countries, the CPI will differ not only in the way prices are combined in the different consumption bundles, but also in the way aggregated prices change over time (Froot and Rogoff, 1994). The same applies for the use of the PPI (produced price index) or the

[11] A simple visual inspection of the data shows an obvious hump down for most of the variables in 1992. Those test are run as we might suspect that for them the non-rejection of the unit root may be due to some misspecification of the data generating process. See Perron (1989).

GDP deflator.

Additional concern is related to the adoption of price measures in levels or differences. Most of the previous studies use price differentials, rather than levels, primarily because transformed prices show better statistical properties.[12] As those assumption are statistically, rather than theoretically, driven they disregard that the expected inflation has to be considered over the same horizon of the bond maturities (over l periods).

Finally, many contributions believe that price series for tradable goods (proxied by the wholesale price index - WPI) may be preferable.[13]

In this paper we use a CPI series for inflation as the consumer indices are similarly defined in countries having common *habits* in consumption, as it may be for the UK-US case.[14] The use of a proxy for tradable goods prices seems less than appropriate in this context. The latter would not only bias the calculation towards the existing spot market foreign exchange rate, but it would also account for the non-tradable components embedded in the tradable goods prices, as transportation costs, tariffs barriers, as well as different patterns of consumption (Jore *et al.*, 1993; Sjoo, 1992).

3.2 Benchmark Analysis

In order to understand the dynamics endogenous to our system, we proceed as follows. We initially model an unrestricted VAR with $k = 2$ lags and an unrestricted constant term μ_0.[15] In light of the unit-root test results - and following Juselius (1995), Juselius and MacDonald (2004), Sjoo (1992) - rather than the 10 years inflation rate we use the current inflation rate, albeit we prefer its year-on-year version in line with the idea that a central bank quotes headline inflation on an annualized basis.[16] The vector of dependent variables takes the form (see Figure 1-2):

$$Y_t = \left[\Delta_{12}p_t, \Delta_{12}p^*_t, i^{120}_t, i^{120*}_t, rer_t\right].$$

where $\Delta_{12}p$ and $\Delta_{12}p^*$ denote the year-on-year *cpi*-inflation in the UK and in the US respectively, i^{120} and i^{120*} are constant maturity Treasury bonds at 10 years and rer is the home vs. foreign real exchange rate, as explained in Section (3).

[12] In Juselius (1995) and Sjoo (1992) price levels are - for instance - transformed in their respective growth rates because of the presence of I(2) trends in the model (see also Johansen, 1991). With our data, prices in levels analogously confirm the existence of I(2)-ness in the model.

[13] See *inter alias* Johansen and Juselius (1992) and Throop (1993).

[14] Moreover, the official measurement of price inflation for most industrialized countries is computed starting from that index.

[15] With this specification, the variables are allowed to contain linear trends, whereas there are no trends in the cointegrating relations. A constant restricted to the cointegrating space has been found not to be significant in the final model. The trace test results and the magnitude of the beta coefficients are nonetheless robust whenever a restricted, unrestricted (or both) constant is selected.

[16] In the purpose of analyzing the relation in (3) as a long run relation (Johansen, 1991; 1994; Johansen and Juselius, 1990; Juselius, 2006) we conduct our analysis requiring all the endogenous variables to be I(1), hence ruling out *a priori* the 10-year inflation rate.

Following the VAR representation:

$$Y_t = \Pi_1 Y_{t-1} + \Pi_2 Y_{t-2} + \mu_0 + \epsilon_t,$$

the VECM can be written in the standard form, as:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \mu_0 + \epsilon_t.$$

where α and β are $(p \times r)$ matrices with $\Pi = \alpha\beta'$. If Π has reduced rank, $\beta' Y_t$ represents the $p \leq r$ valid cointegrating vectors, while α contains the loadings measuring the adjustment of ΔY_t to the long run equilibrium. For estimation purposes we will need assuming ϵ_t to be *i.id.* $N(0, \Omega)$. Preliminary tests point to the rejection of normality for some of the variables. To secure valid statistical inference we need to control for intervention effects falling outside the normality confidence band. If a residual is observed to be larger than $[3.74\sigma^\epsilon]$, this corresponds to a known intervention which needs being controlled with a dummy variable.[17]

Following this criterion, the model includes four dummies. Each dummy *Dmm.yy* measures a blip interventional shock in month *mm* of year *19.yy* (Table 2). A shift dummy (structural break) is also included in order to amount for a shift in the mean for the UK inflation due to the ERM sterling admission in 1990.[18] The break is broadly consistent with the one identified from the unit root tests reported in Table 1.[19] The lag length determination test finally reports the existence of $k = 4$ lags in the model.[20] The baseline VAR is finally represented as:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + \Gamma_3 \Delta Y_{t-3} + \mu + \Phi D_t,$$

where D_t accounts for interventional dummy variables.

As a preliminary check of the adequacy of the model we run a set of misspecification tests. In Table 3 we report several multivariate and univariate tests (see Dennis, 2006). In the first section of the Table, we report a Ljung-Box test based on the auto and cross-correlation of the first $[T/4]$ lags, an ARCH test and a normality Doornik-Hansen test. The second section reports some univariate statistics based on the estimated residuals for each equation. The results show that the multivariate LM test for residuals autocorrelation is not significant. The multivariate LM tests for no (first order)

[17] The critical value is derived as the inverse normal distribution of $(1 - 0.025) * (1/es)$, with es being our effective sample. For further reference see Dennis (2006) and Juselius (2006).

[18] Following the frenetic speculation over the British currency, in September 1992 UK will be driven out of the ERM (Black Wednesday). As the UK left the agreement (September, 16th), interest rates were able to fall.

[19] In the model we control for a blip in 09:2006 and 04:1991 in the US and the UK inflation respectively; in 03:1990 in the UK bond rate and in 10:1992 in the RER. Finally, the model includes a break in the UK inflation level in 04:1992.

[20] Lag-length determination test is based on log-likelihood and information criteria.

conditional heteroskedasticity is rejected whilst joint normality is violated, essentially for the two inflation rates. According to the univariate tests results, the rejection of the normality assumption is however mainly imputable to excess kurtosis and hence no serious for our estimation results (Juselius, 2006).[21]

Under the assumption that the model is well specified, we determine the number of cointegrating vectors. In line with the idea that a CVAR can approximate well a monetary/fiscal authority long-run policy target (e.g. Christensen and Nielsen, 2009;), the Johansen's trace test statistic suggests the existence of one valid cointegrating relation (Table 4) (i.e. with the relation in (3) approximating the behaviour of the equilibrium real exchange rate, see Juselius, 1995; Juselius and MacDonald; 2004).[22] Given rank equal to one, we report the test for long run exclusion and weak exogeneity. The former checks whether any of the variables can be excluded from the cointegrating space by testing for a zero element in the β vector.[23] Alternatively, the test of long run weak exogeneity investigates the absence of long-run feedback effects, being formulated with a zero element in α. If the hypothesis of weak exogeneity is accepted, the variable can be considered as a driving force in the system, and hence not-adjusting to the long-run relation depicted by the cointegrating space.

None of the variables can be excluded under the assumption of $r = 1$.[24] The weak exogeneity assumption for the UK bond rate is not rejected at the 5% critical level, whereas the US bond rate is found not to be driving. This result supports the idea that the special USD status results into a weaker behavior of the US bond rate in this system.

The unrestricted estimates for α and β are reported in Table 6 under H_0. The UK bond yield is normalized to unity. The β coefficients are far from accommodating the *symmetry* (between the domestic and the foreign economy) and *proportionality* (between the relative inflation and the exchange rate) conditions implicit in the relation in (2). Moreover, we find evidence in favour of a "forward premium" puzzle, due to the positive sign in the RER term differently from what (2) would predict.[25] Imposing the joint restriction $[-1, 1, 1, -1, 1]$ is moreover rejected at a standard significance level (H_2), whereas the rer_t can not be excluded from the cointegrating space (H_1) (or the RIP does not hold here; see Macchiarelli, 2011). These results, while confirming the special USD status, confirm the evidence in favour of a weak version of (3) found by Juselius (1995), Juselius and MacDonald (2004), Jore *et al.* (1993) and Sjoo (1992), for currency pairs other than the UK sterling vs. the US dollar

[21] Simulation studies have shown that valid statistical inference is sensitive to the validity of some of the assumptions, like parameter non-constancy, autocorrelated residuals and skewed residuals, while quite robust to others, like excess kurtosis and residual heteroscedastisity (Juselius, 2006).

[22] As an additional check of the validity of our analysis, we input the US and the UK realization of inflation over 10 years, instead of the current changes in prices. The number of long run relations sensitively increases from $r = 1$ to $r = 3$, confirming the unit-root test results.

[23] Recall that in this setting both α and β are (5×1) vectors.

[24] The exclusion restriction for the US inflation can not be barely rejected at the 5% but it is rejected at the 10% level. As this assumption is not motivated theoretically, we proceed with the original vector of variables.

[25] The "forward premium" puzzle is well documented in the UIP literature, predicting high interest rate currencies to appreciate rather then depreciate as theory would suggest.

during the post Bretton-Woods.

3.3 Modeling Inflation

The theoretical framework grounding the analysis assumes individuals to trade off between a 10 years constant maturity bond and their inflationary expectations over the next 10 years. Ideally, we would extract expected inflation on the bond markets by comparing nominal and inflation-indexed yields at the same maturity. Unfortunately, for the US the inflation-indexed market yield on U.S. Treasury securities at 10 years (constant maturity) is available only since January 2003, yielding a too short sample for estimation. Instead the inflation-indexed yield from British Government Securities at 10 years (zero coupon) dates back to 1985-1.[26] Due to missing data, we then employ a statistical measure of expected inflation derived from a structural unobserved component model. In practice, we consider two models all assuming expected inflation to be described by the long-run inflation forecast or permanent inflation component. For each country j (with $j = UK, US$), Model 1 assumes a standard permanent-transitory decomposition for inflation:

[Model 1]

$$\pi_{j,t} = \pi^e_{j,t} + v_{j,t},$$

with

$$\pi^e_{j,t} = \pi^e_{j,t-1} + u_{j,t},$$

where, differently from our previous notation, π is the current inflation while π^e is its long-run component, following a driftless random walk as in Cogley (2002) and Cogley and Sargent (2007). The error terms, $v_{j,t}$ and $u_{j,t}$, are serially and mutually uncorrelated with zero mean and variance σ^2, i.e. $v_t \sim NID(0, \sigma^2_v)$, where the notation NID stands for normally and independently distributed (see Koopman *et al.*, 2009). Based on this model, inflation depends exclusively on its expectation plus an irregular component.

Model 2 is instead a Phillips-type equation where deviations of inflation from its long-run equilibrium depend on the economy output gap and a net oil price increase (NOPI) following Hamilton (1996) (see Figure 3).

[26] Real rates are end-of-month and are extracted respectively from the Federal Reverve Bank of St. Luis and the National Bank of England.

[Model 2]

$$\pi_{j,t} = \pi^e_{j,t} + \delta' z_{j,t} + v_{j,t},$$

with

$$\pi^e_{j,t} = \pi^e_{j,t-1} + u_{j,t},$$

and

$$z_t = \begin{bmatrix} y_t - \bar{y}_t \\ NOPI_t \end{bmatrix}.$$

For Model 2 as well, standard assumptions on the irregular components apply, i.e. $w_t \sim NID(0, \sigma^2_w)$. Overall, the two models tell a similar story (Figure 5), although inflation gaps ($\omega_t = \pi_t - \pi^e_t$) show an overall higher degree of correlation for the UK rather than for the US (Table 8).

Under the characterizations in Model 1 and 2, we clearly depart from the *rational expectation* hypothesis, as we assume market participants having no knowledge of the causal mechanism driving macroeconomic fundamentals, forming expectations about what inflation would be in the future on the basis of its past realization. Hence, macroeconomic determinants (output gap, net oil price shocks...) do not enter the information set used by agents in forming long run beliefs.[27] In Figure 6 we plot Model 2 estimates together with the yield spread - between nominal government bonds and those inflation-indexed - which proxy the return investors require for being compensated for the expected inflation and the inflation risk over the lifetime of the bond (*inflation compensation*, see Gürkaynak *et al.*, 2010). In both the UK and the US case, the model behaves quite well being roughly consistent with financial markets expectations. Albeit this is not our focus here, this framework can be used to assess the validity of - more or less explicit - inflation targeting regimes (e.g. Figure 6 suggests how an explicit numerical inflation objective seems to have improved the anchoring of long-term inflation expectations in the UK).

3.4 Inflation-Consistent Analysis

Given our new expected-inflation estimates, we assess the robustness of the above results by repeating the analysis proposed in Section 3.2. The new vector of dependent variables is

$$Y_t = \left[\pi^e_t, \pi^{e*}_t, i^{120}_t, i^{120*}_t, rer_t\right],$$

[27] This complements the evidence reported by Frydman *et al.* (2008, 2010). According to the authors, once imperfect knowledge is recognized in the model, a monetary framework is able to account for the PPP puzzle, as well as for the observed long-swings in the exchange rates data.

where the expected inflation from Model 2 appears instead of the current annualized inflation. For sake of simplicity, from now on we focus on the results from Model 2 (yielding more economically grounded estimates), and refer to the latter as the VAR model where Model 2 expected inflation appears.[28]

As before, we control for intervention effects falling outside the normality confidence bands. Again, the model includes some impulse dummies plus a level shift amounting for the change in the UK inflation mean after 1990.[29] The lag length of the model has been set to $k = 4$ consistent with the information criteria considered. Finally, the Johansen's trace test statistic confirms the existence of one cointegrating vector (Table 9).

Given $r = 1$, the unrestricted estimates for α and β are reported in Table 10 under H_0. As before, normalization occurs on the UK bond rate. We first check whether the unrestricted results are consistent with the hypothetical long-run relation (equation (4)): the coefficients have the correct signs, as they are consistent with the benchmark analysis proposed in Section 3.2. The *symmetry* and *proportionality* conditions are nonetheless still rejected at a standard significant level (H_2), albeit this is now mainly imputable to UK misalignments. As evidenced by the unrestricted beta coefficients, imposing a *proportional* adjustment in between the US bond yield, the US inflation and the RER term is in facts not rejected *per sé* (for the joint restriction the p-value of the LR test is [0.983]).[30] Moreover, while it would be theoretically plausible setting $\beta_{rer} = 0$, based on the t-ratio values reported in Table 10 this restriction is rejected at the 1% level (under H_1) (see also Macchiarelli, 2011). To this extent, the empirical evidence in favour of a long-run PPP-UIP equilibrium persists also when accounting for expectations, thus supporting the evidence in favour of the RER acting as *error correction mechanism* in such a combined equilibrium (see also Juselius, 1995; Juselius, 2006; Johansen, 1992; Sjoo, 1992).

Given the first column values of α under H_0, we finally proceed by testing for weak exogeneity. We separately test for weak exogeneity of (i) the UK bond rate, consistently with our benchmark analysis (under H_3), (ii) the US inflation, consistently with the idea that US inflation drives inflation in the UK (under H_4), (iii) and the US bond rate, meaning no USD "credibility" in the long run (H_5). Albeit the α restrictions are tested separately in Table 10, the former two hold also jointly with p-value [0.850]. Instead, when the restriction $\alpha_{b_{US}} = 0$ is imposed, the p-value for the LR drops considerably below the 5% critical level. As accommodating the weak exogeneity of the US bond would imply no role for USD "credibility", we do not find evidence that in the long run - as soon as expectations,

[28]The current expected inflation appears in the VAR instead of its 10 years expectation, as by construction the inflation expected over 10 years equals the expectation today plus a sum of white noise errors.

[29]In the model we control for a blip in 10:2005 in the US inflation; in 03:1990 in the UK bond rate and in 10:1992 in the RER. Finally, the model includes a break in the UK inflation level in 05:1991.

[30]Proportionality among US inflation and the bond yield is not rejected alone with p-value [0.985]. The whole results are available upon request from the author.

rather than prices, adjust - the dollar special status vanishes. This result deserves further discussion. First, it shows that our findings are somewhat robust to the inclusion of inflationary expectations. Nonetheless expectations help the fitting of our model, yielding a coefficient for the RER term which is much closer to that observed for the US real interest rate.

Additionally, as stressed in Juselius and MacDonald (2004), the fact that the US bond is not found to be weakly exogenous is, in any case, surprising, given that "US financial shocks are generally believed to lead international capital markets". Nonetheless, we believe such a result to reconcile with the traditional observed gap between US and UK government bonds, with a high deficit but proportionally low(er) rates in the US (see discussion in Section 3.2). Investors' high willingness to hold dollar-denominated assets possibly dampens the necessary US capital reaction on international markets. In fact, concerns regarding the sustainability of the US deficit do not lead to an increase in the US risk premium, resulting into a lower than expected upward pressure on long term interest rates.

3.4.1 Characterizing the International Parity Condition

The long run relation given by Model 2 can be written as:

$$\left[(i_t^{120} - 4.3\pi_t^e) - (2.9i_t^{120*} - 2.8\pi_t^{e*})\right] + 2.7rer_t = 0, \tag{5}$$

As mentioned, the RER term has the opposite sign with respect to what the relation in (2) would predict. This is in line with the presence of a "forward premium" puzzle in the UIP test (a negative regression coefficient), predicting high interest rate currencies to appreciate rather then depreciate as the standard interest parity would suggest. Overall, our results are consistent with the idea that the US real interest rate would increase over the UK one when the RER term is negative (i.e. the real interest spread is positive for a negative RER). In other words, when US prices are above prices in the UK measured in the same currency (RER term negative), the US long term interest rate has always be to be average higher with respect to the UK bond rate.[31] Capital market adjustments are hence consistent with a *general equilibrium* framework where equation (5) is able to keep exchange markets in line against speculative activities or commodity prices deviations.

Based on the unrestricted α coefficients in Table 10 (under H_0), deviations from (5) are indeed (slowly) corrected *via* the UK inflation, eliminating about 0.2% of misalignment each month.[32] Most of the misalignments are indeed corrected *via* the RER term (-0.9%). Surprisingly, the US bond rate is found

[31] In addition, this somewhat reconciles with theories stating that when the real interest rate is high the currency of the same country tends to be strong in real terms (e.g. Dornbush, 1976). See also Engel (2011).

[32] As discussed, adjustments on the side of the US inflation are not significant, in line with the weak exogeneity assumption under H_4.

to "push" in the opposite direction: +1.2%, which is quite high compared to the other αs. While, statistically, expectations do not help impose the restriction of both rates to be driving, when looking at bond yields corrections the US is still puzzling, possibly reconciling with a failure of the UIP alone (e.g. Bekaert *et al.*, 2007; Macchiarelli, 2011).

The overall results are nevertheless consistent with our prior expectations, supporting the idea that when the RER term is out of its PPP equilibrium level (long run), the equilibrium is restored with a very slow adjustment on the goods markets and with a major adjustment on the capital and hence on the exchange rate markets.[33]

As an important issue for the interpretation of our results is the structural stability of our estimates, we finally subject our chosen model to a set of Hansen-Johansen (1999) stability tests. For sake of easiness, we report only a test for *beta constancy* (Figure 7). The test gives important information about the constancy of our individual cointegrating relation, revealing the model to be acceptable and showing a fair degree of stability.[34] The technical discussion of the test is left to the Figure's footnote.

3.4.2 The Long Run Impact of Shocks

Based on a stable cointegrating vector, we can obtain the common trends representation of our restricted model by simply inverting the vector process trough the vector moving-average representation (VMA), i.e.

$$Y_t = \tau_0 + C^*(1)\mu_0 + C\sum_{i=1}^{t}\epsilon_i + C\sum_{i=1}^{t}D_i + C^*(L)(\epsilon_t + \Phi D_t),$$

where $\tau_0 = \tau(Y_0)$ and $C^*(L) = \sum_{i=0}^{\infty} C_i^* L_i$ is a convergent matrix polynomial in the lag operator. The long run impact matrix, $C = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp = \tilde{\beta}_\perp \alpha'_\perp$, with its decomposition being similar to that of Π, with the difference that now $\tilde{\beta}_\perp$ determines the *loadings* whilst $\alpha'_\perp$ identifies the common stochastic trends. Under this assumption, the non-stationarity of Y_t is originated indeed by the cumulative sum of the $(p-r)$ combinations of $\alpha'_\perp \sum_{i=1}^{t} \epsilon_i$.

We rely on Model 2, when the UK bond and the US inflation are set to be weakly exogenous (see Section 3.4).[35] Given the results reported at the bottom of Table 11, the estimates of the common stochastic trends imply the non stationarity of the VAR to be driven by business cycle shocks to both goods and capital markets. Those are shown in the expected inflations and RER term interactions with the US bond rate, in the former two stochastic trends.

Based on an inspection of the long run impact matrix C, we can additionally check how the cumulated residuals from each VAR equation impact the others, given the corresponding weights (row wise

[33] This approach belongs to the sticky prices monetary models. For further details see Dornbusch (1976).

[34] For further details see Juselius (2006).

[35] Looking at the top of Table 11, the results for the lung run impact matrix C - when no weak exogeneity assumption is made (H_0) - are not significantly different.

analysis). Inspecting the third and fourth rows of C reveals how shocks to the UK and the US long term yields respectively increase expected inflation in the US but not in the UK; consistent, in the former case, with a Fisherian (1907) view of nominal rates. This finding is in line with the idea that nominal long term rates have historically had an important role in influencing expected inflation in the US. In fact, episodes where longer-term rates have crucially driven inflation expectations, thus prompting a monetary policy reaction, are well documented in the literature (see Goodfriend, 1998). From an analytical perspective, this suggests that occasions when long bond rates jumped prompting an inflation scare, or, on the contrary, fell due to a gain of credibility for future lower inflation by the Fed, are well captured by the data.

Finally, the results suggest that the cumulated empirical shocks to the RER have a positive, though hardly significant, effect on the UK bond rate, and a negative, but not significant, effect on the US bond rate. This latter finding, a part from suggesting imperfect price adjustments on the international markets, may explain the failure of the PPP alone.[36]

4 Conclusions

The purpose of this analysis was to illustrate the extent to which the real exchange rate (RER) observed deviations could explain bilateral capital movements. This led us to combine the PPP and the UIP hypothesis in a CVAR *general equilibrium* framework, admitting goods and capital markets to interact in keeping exchange markets in line. There are two main reasons for which one would expect PPP and UIP not to hold separately. First, commodity price movements or speculative activities may lead to significant deviations from PPP/UIP. Secondly, price stickiness out of the equilibrium, together with the presence of limits to arbitrage and related market imperfections, may undermine short-term adjustments. As prices in markets for goods tend to adjust to shocks more slowly than prices in markets for assets, PPP deviations are more likely to persist if compared to UIP deviations (Pesaran *et al.*, 2000). Combining them both hence allows for a gradual convergence to PPP as it seems that the "PPP relation needs a lift [...] to become more stable, and this is what the interest rates do" (Johansen, 1992).

The econometric strategy adopted in our "benchmark" analysis (Section 2.3) gave initial suggestions of the expected adjustments on the side of the exchange markets (RER term) to be very strong. This supports the evidence in favour of the RER acting as *error correction mechanism* in a PPP-UIP combined relation (see also Juselius, 1995; Juselius, 2006; Johansen, 1992; Sjoo, 1992).

Relatively to the existing literature, the current paper innovates on two fundamental grounds. First,

[36]The two bond yields are equally, though not significantly, affected by cumulated shocks to the RER.

we explore a role for the "credibility" of the USD, implicit in its special status as World reserve currency. The literature reports indeed weak forms of PPP to hold for USD bilateral parings (see MacDonald, 1998; Juselius and MacDonald, 2004) as the result of adopting as *numeraire* a currency which is "exceptional". Secondly, we extend the analysis by allowing for markets expectations. As discussed in Campbell *et al.* (2007), PPP (and UIP) deviations may be due to people not generally and promptly adjusting to changes in the behavior of prices, so that in our set up expectations are accounted for by means of a dynamic latent factor model.

The overall results point to a role for the special status of the USD both when the observed inflation and its long-run expectation enter the CVAR. Using our new estimates for inflation, we find a combined UIP-PPP relation not to hold strictly as a long run condition essentially because of UK misalignments: the cointegrating vector is not found to accommodate *symmetry*, but only *proportionality* between the US inflation, the US bond yield and the RER term. Inference based on a typical cointegration analysis moreover shows that the goods market adjustment is very slow, whilst a major adjustment occurs on the capital and on the exchange rate markets. Nonetheless, the US bond rate is found to "push" in the opposite direction, possibly reconciling with a failure of the UIP itself (e.g. Bekaert *et al.*, 2007; Macchiarelli, 2011). Finally, looking at how accumulated empirical shocks to each variable affect the others, we find that empirical shocks to the US and the UK long term yields respectively increase expected inflation in the US but not in the UK; consistent, in the former case, with a Fisherian (1907) view of nominal rates. Moreover, the accumulated empirical shocks to the RER significantly affect the expected inflation in the UK but not in the US. Those latter findings signal imperfect price/capital adjustments on the international markets, and may account for the widespread finding of rejection of the UIP/PPP conditions when tested separately.

References

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

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

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

[4] 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.

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

[6] Christensen A.M., Nielsen H.B. (2009), Monetary Policy in the Greenspan Era: A Time Series Analysis of Rules vs. Discretion, *Oxford Bulletin of Economics and Statistics*, **71**(1), 69-89.

[7] Cogley T. (2002), A simple adaptive measure of core inflation, *Journal of Money, Credit and Banking*, 34, 94-113.

[8] Cogley T., Sargent T. (2007), Inflation-Gap persistence in the US, Mimeo in Harvey (2008).

[9] Dennis J.G. (2006), CATS in RATS, Cointegration Analysis of Time Series - version 2, Estima, Evanston, Illinois (USA).

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

[11] 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.

[12] Engel C. (2011), The Real Exchange Rate, Real Interest Rates, and the Risk Premium, Manuscript.

[13] Fisher E.O'N., Park J.Y. (1991), Testing Purchasing Power Parity under the Null Hypothesis of Co-Integration, *The Economic Journal*, **409**(101), 1476-1484.

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

[15] Frenkel J.A. (1980), The collapse of Purchasing Power Parities during the 1970s, NBER Working Paper Series, No. 569.

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

[17] Frydman *et al.* (2008), A Resolution of the Purchasing Power Parity Puzzle: Imperfect Knowledge and Long Swings, University of Copenhagen.

[18] Frydman *et al.* (2010), Testing Hypotheses in an I(2) Model With Piecewise Linear Trends. An Analysis of the Persistent Long Swings in the Dmk/$ Rate, *Journal of Econometrics*, **158**(1), 117-129.

[19] Goodfriend M. (1998), Using the Term Structure of Interest Rates for Monetary Policy, Federal Reserve Bank of Richmond *Economic Quarterly*, **84**(3), 13-30.

[20] Gürkaynak R.S. *et al.* (2005), The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models, *American Economic Review*, **95**(1), 425-436.

[21] Gürkaynak R.S. *et al.* (2006), Does Inflation Targeting Anchor Long-Run Inflation Expectations? Evidence from Long-Term Bond Yields in the U.S., U.K., and Sweden, FRBSF Working Paper 2006-09.

[22] Gürkaynak R.S. *et al.* (2010), The TIPS Yield Curve and Inflation Compensation, *American Economic Journal: Macroeconomics*, American Economic Association, **2**(1), 70-92.

[23] Hamilton, J.D. (1996), This is what happened to the oil price-macroeconomy relationship, *Journal of Monetary Economics*, Elsevier, **38**(2), 215-220.

[24] Harvey A., 2008. Modeling the Phillips curve with unobserved components, Cambridge Working Papers in Economics No. 0805, Faculty of Economics, University of Cambridge.

[25] Hodrick R.J., Prescott E.C. (1997), Postwar US Business Cycles: An Empirical Investigation, *Journal of Money, Credit and Banking*, 29, 1-16.

[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. (1991), Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models, *Econometrica*, **59**, 1551-1580.

[28] Johansen S. (1994), Likelihood Based Inference for Cointegration of Non-Stationary Time Series, in Katarina Juselius (edited by), Econometric Modeling of Long Run Relations and Common Trends: Theory and Applications - Volume I, University of Copenhagen.

[29] Johansen S. (2002), A small sample correction for tests of hypotheses on the cointegrating vectors, *Journal of Econometrics*, **111**, 195-221.

[30] Johansen S. *et al.* (2008), Allowing the Data to Speak Freely. The Macroeconometrics of the Cointegrated Vector Autoregression, *American Economic Review*, **98**(2).

[31] Johansen S., Juselius K. (1990), Maximum Likelihood Estimation and Inference on Cointegrating - With Applications to the Demand for Money, *Oxford Bulletin of Statistics and Economics*, **52**, 169-211.

[32] 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.

[33] 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.

[34] 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.

[35] Juselius K. (1992a), Domestic and Foreign Effect on Prices in an Open Economy. The case of Denmark, *Journal of Economic Policy Modeling*, **14**, 401-428.

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

[37] 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.

[38] Juselius K. (2006), The Cointegrated VAR Model, Methodology and Applications, Oxford University Press, New York.

[39] Juselius, K., MacDonald, R. (2000), International Parity Relationships between Germany and the United States: A Joint Modelling Approach, University of Copenhagen Discussion Paper, no. 00-10.

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

[41] Koopman S.J. *et al.* (2009), STAMP 8.2: Structural Time Series Analyzer, Modeller and Predictor, London: Timberlake Consultants Press.

[42] Krugman K. (1993), Exchange Rates, *The Concise Encyclopedia of Economics* (Henderson D.R., ed.). Originally published as *The Fortune Encyclopedia of Economics*, Warner Books. Library of Economics and Liberty [Online] available from http://www.econlib.org/library/Enc1/ExchangeRates.html

[43] 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.

[44] Lumsdaine L., Papell D.H. (1997), Multiple Trend Breaks and the Unit-Root Hypothesis, *The Review of Economics and Statistics*, **79**(2), 212-218.

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

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

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

[48] Macchiarelli C. (2011), A VAR Analysis for the Uncovered Interest Parity and the Ex-Ante Purchasing Power Parity. The Role of Macroeconomic and Financial Information, ECB Working Paper 1404.

[49] 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.

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

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

[52] Perron P. (1997), Further evidence on breaking trend functions in macroeconomic variables, *Journal of Econometrics*, **80**, 355-385.

[53] Ravn M.O., Uhlig H. (2002), Notes on Adjusting the HP Filter for the Frequency of Observations, *The Review of Economics and Statistics*, MIT Press, **84**(2), 371-375.

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

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

[56] Sjoo B. (1992), Foreign transmission effects in Sweden: do PPP and UIP hold in the long run?, *Advances in International Banking and Finance*, **1**, pp. 129-149.

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

[58] 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.

[59] 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.

[60] 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.

[61] Zivot E., Andrews D.W.K. (1992), Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis, *Journal of Business and Economic Statistics*, American Statistical Association, vol. **10**(3), 251-70.

Table 1: Augmented Dickey Fuller Test Statistics

	Standard ADF[a]	Lumsdaine and Papell ADF[b,d]	Zivot and Andrews ADF[b,d]	Perron ADF[c,d]
$\Delta_{12}p_t$	-1.628	-3.441	-3.562	-3.434
	(12)	(12)	(12)	(12)
$\Delta_{12}p_t^*$	-1.954	-3.231	-3.316	-3.375
	(15)	(15)	(15)	(15)
$\Delta_{120}p_t$	-3.128	–	–	–
	(14)			
$\Delta_{120}p_t^*$	-3.763	–	–	–
	(1)			
i_t^{120}	-0.803	-3.263	-3.263	-3.252
	(15)	(10)	(10)	(10)
i_t^{120*}	-1.142	-5.622	-5.621	-5.759
	(14)	(8)	(8)	(11)
rer_t	-2.992	–	–	–
	(7)			
CVs				
1%	-3.46	-7.24	-5.34	-5.70
5%	-2.87	-6.82	-4.80	-5.10
10%	-2.57	-6.65	-4.58	-4.82
Break	none	1	1	1

Notes: [a] For the ADF the max lag length is chosen by mean of information criteria. The number of lags is in parenthesis. [b] The lag length is automatically chosen with a general-to-specific recursive t-test. The test selects the number of lags for which the last lag included has a marginal significance level less than the cutoff value 0.10. In all cases the max lag length to start the recursion with is 15. [c] The reported critical values are for finite samples. The asymptotic values are however not dramatically different (see Perron, 1997). [d] The methods employed are not directly geared at providing a consistent estimate of the dated change. In this respect the dated break is irrelevant and should be viewed as approximate.
Lumsdaine-Papell (1997) Unit Root Test with one break in Intercept (UK inflation 1992:03; US Inflation 1992:07; UK bond 1997:04; US bond 1992:04). Zivot-Andrews (1992) Unit Root Test with one break (UK inflation 1989:04; US Inflation 1989:04; UK bond 1989:02; US bond 1992:05). Perron (1997) Unit Root Test with breaking trend (UK inflation 1992:02; US Inflation 1992:06; UK bond 1997:03; US bond 2005:05).

Table 2: The Unrestricted VAR(4) Model with Dummy Variables

	$\Delta Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + \Gamma_3 \Delta Y_{t-3} + \mu + \Phi D_t$ with $D_t = [Dum06.09, Dum91.04, Dum92.10, Dum90.03]$
$Dum06.09$	During 2005 the US wholesale price indexes rose at their fastest rate in 15 years, as a consequence of growing energy prices. The FED pursued the strategy of gradually increasing the interest rates to prevent energy prices pressure to infect other parts of the economy with inflationary trends. In 2006 inflation remained high with services being the main critical source of such an upward pressure.
Other dummies	With inflation getting out of control (9.5 % in 1990), the UK government decided to join the ERM (October, 8th). In the monetary system the value of the sterling was kept within certain boundaries against the D-Mark (± 2,25). However, under the pressure of increasing inflation, the British currency became weaker on the exchange markets. The government let interest rates to increase in order to prevent the currency to fall. In 1991, this led to a steep fall in consumption. The UK economy moved close to a recession.

Table 3: Tests for Misspecification of the VAR(4)

Multivariate tests					
Resid. autocorr.	Ljung-Box(69)	$\chi^2(1625)$	1765.851	(0.008)	
ARCH	LM(1)	$\chi^2(225)$	242.684	(0.199)	
ARCH	LM(2)	$\chi^2(450)$	521.516	(0.011)	
Normality	LM	$\chi^2(10)$	33.937	(0.000)	
Univariate tests	$\Delta_{12} p_t$	$\Delta_{12} p_t^*$	i_t^{120}	i_t^{120*}	rer_t
ARCH (4)	5.856	16.253	8.616	7.702	3.290
	(0.210)	(0.003)	(0.071)	(0.103)	(0.510)
Normality	10.699	9.984	9.231	2.972	3.410
	(0.005)	(0.007)	(0.010)	(0.226)	(0.182)
Skewness	-0.001	0.427	-0.212	0.233	-0.028
Kurtosis	3.929	3.816	3.857	2.891	3.450

Notes: In the first section of the Table, we report a Ljung-Box test based on the auto and cross-correlation of the first $[T/4]$ lags, an ARCH test and a normality Doornik-Hansen test. The second section reports some univariate statistics based on the estimated residuals for each equation. p-values are in parenthesis.

Table 4: Trace Test Statistics

	r = 0	r = 1	r = 2	r = 3	r = 4
Trace test	101.414	48.056	25.662	12.772	3.878
	(0.016)	(0.680)	(0.835)	(0.776)	(0.671)
Trace test (Bartlett corrected)	94.892	45.177	23.931	11.693	3.385
	(0.053)	(0.793)	(0.895)	(0.825)	(0.735)

Notes: The Table reports the estimates for the Johansen's (1994) trace test statistics. p-values are in parenthesis. The test also reports the results corrected for small-sample bias according to a Bartlett correction (see Johansen, 2002).

Table 5: Tests for Long Run Identification (under Rank of $\Pi = 1$)

	$\Delta_{12}p_t$	$\Delta_{12}p_t^*$	i_t^{120}	i_t^{120*}	rer_t	DGF	5% CVs
Exclusion restriction	28.037	2.899	12.504	9.134	5.683	1	3.84
	(0.000)	(0.089)	(0.000)	(0.002)	(0.015)		
Weak exogeneity	25.924	3.803	0.002	9.646	2.646	1	3.84
	(0.000)	(0.051)	(0.926)	(0.002)	(0.104)		

Notes: The Table reports the test for long run exclusion and weak exogeneity under the assumption of $rank(\Pi) = 1$. The former test checks whether any of the variables can be excluded from the cointegrating space by testing for a zero element in the β vector. Alternatively, the test of long run weak exogeneity investigates the absence of long-run feedback effects, being formulated with a zero element in α. p-values are in parenthesis.

Table 6: The Long Run Identification Structure

	H_0		H_1		H_2	
	α	β	α	β	α	β
$\Delta_{12}p_t$	0.060	-1.103	0.071	-1.012	0.033	-1.000
	(5.679)	(-6.714)	(5.283)	(-6.282)	(3.767)	–
$\Delta_{12}p_t^*$	0.029	0.522	0.019	0.669	0.010	1.000
	(2.276)	(2.142)	(1.167)	(3.356)	(0.958)	–
i_t^{120}	-0.001	1.000	-0.004	1.000	-0.008	1.000
	(-0.101)	–	(-0.062)	–	(-0.917)	–
i_t^{120*}	0.036	-1.173	0.048	-1.358	0.019	-1.000
	(3.431)	(-6.391)	(3.561)	(-9.443)	(2.149)	–
rer_t	-0.022	0.573	-0.006	0.000	-0.024	1.000
	(-1.939)	(3.162)	(-0.383)	–	(-2.584)	–
$C(1992:04)$	–	-2.586	–	-2.091	–	0.140
		(-3.334)		(-3.095)		(0.235)
Log likelihood	2088.263		2085.331		2078.510	
LR test	–		$\chi^2(1) = 5.863$		$\chi^2(4) = 19.505$	
p-value			(0.015)		(0.001)	
Bartlett corrected LR			$\chi^2(1) = 3.441$		$\chi^2(4) = 12.321$	
p-value			(0.064)		(0.015)	

Notes: In the VECM representation, $Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \mu + \Phi D_t + \epsilon_t$, α and β are (5×1) vectors with $\Pi = \alpha\beta'$. If Π has reduced rank, $\beta' Y_t$ represents the $p \leq r$ valid cointegrating vectors, while α contains the loadings measuring the adjustment of ΔY_t to the long run equilibrium. C(1992:04) represents the break in levels spotted by the large residuals detection procedure. For each estimate asymptotic t-ratios are reported in parenthesis. Under the restrictions in column 2 and 3 a Likelihood-Ratio test is reported at the bottom of the Table. A version of the test corrected for small sample bias is also reported in the last line (for Bartlett correction see Johansen, 2002).

Table 7: Modelling Inflation Results

		Model 1	Model 2		Model 1	Model 2
Log-likelihood	US	975.039	929.049	UK	1500.24	1426.05
	σ_v^2	9.369	9.053		37.667	34.080
		(1.000)	(1.000)		(1.000)	(1.000)
	σ_u^2	0.366	0.169		0.524	0.248
		(0.039)	(0.017)		(0.014)	(0.007)
	σ_w^2					
	δ_1	–	0.058		–	0.015
			[0.010]			[0.019]
	δ_2	–	0.501		–	0.512
			[0.161]			[0.201]

Notes: (q-ratios) and [standard errors] in parentheses. q-ratios are signal-to-noise ratios computed as $q = \sigma_u^2/\sigma_v^2$. Normalization occurs on higher variance values. Model 1 is a simple permanent-transitory decomposition for inflation; Model 2 augment the measurement equation in Model 1 with two exogenous regressors (the economy output gap (δ_1) and a net oil price increase *à la* Hamilton (δ_2)). In each model the variances are the one associated to the measurement (σ_v^2) and the transition equations (σ_u^2).

Table 8: Inflation Gaps Cross-Correlation Matrix

		Model 1	Model 2		Model 1	Model 2
US	Model 1	1.000	0.329	UK	1.000	0.770
	Model 2		1.000			1.000

Notes: The Table reports cross-correlation results for the inflation gaps from Model 1 and 2 computed as $\omega_t = \pi_t - \pi_t^e$. Model 1 is a simple permanent-transitory decomposition for inflation; Model 2 augment the measurement equation in Model 1 with two exogenous regressors (the economy output gap and a net oil price increase *à la* Hamilton (1996)).

Table 9: Inflation-Consistent Trace Test Statistics

		r = 0	r = 1	r = 2	r = 3	r = 4
Model 2	Trace test	99.093	52.987	26.754	12.143	0.018
		(0.017)	(0.407)	(0.760)	(0.809)	(0.566)
	Trace test (Bartlett corrected)	88.536	46.269	23.526	9.404	4.149
		(0.103)	(0.529)	(0.925)	(0.809)	(0.566)

Notes: The Table reports the estimates for the Johansen's (1994) trace test statistics. p-values are in parenthesis. The test also reports results corrected for small-sample bias according to a Bartlett correction (see Johansen, 2002). The results are displayed for Model 2 (trend-cycle decomposition model). For sake of simplicity, it is referred to Model 2 as the VAR model where Model 2 expected inflation appears.

Table 10: The Long Run Structure with Inflationary Expectations

Model 2

	H_0		H_1		H_2	
	α	β	α	β	α	β
π_t^e	0.002	-4.428	0.005	-0.740	0.002	-1.000
	(5.201)	(-5.149)	(2.975)	(-3.156)	(1.387)	–
π_t^{e*}	0.000	2.837	0.003	0.035	-0.002	1.000
	(0.537)	(2.741)	(1.310)	(0.110)	(-1.309)	–
i_t^{120}	-0.002	1.000	-0.025	1.000	-0.014	1.000
	(-0.443)	–	(-1.706)	–	(-1.544)	–
i_t^{120*}	0.012	-2.886	0.042	-1.693	0.020	-1.000
	(3.258)	(-4.721)	(2.965)	(-8.883)	(2.233)	–
rer_t	-0.009	2.675	0.014	0.000	-0.016	1.000
	(-2.133)	(4.694)	(0.869)	–	(-1.625)	–
$C(1991:05)$	–	-14.921	–	-3.387	–	-0.050
		(-5.587)		(-4.279)		(-0.065)
Log likelihood	3153.153		3147.232		3140.632	
LR test	–		$\chi^2(1) = 11.844$		$\chi^2(4) = 25.043$	
p-value			(0.001)		(0.000)	
Bartlett corrected LR			$\chi^2(1) = 6.760$		$\chi^2(6) = 15.378$	
p-value			(0.009)		(0.004)	
	H_3		H_4		H_5	
	α	β	α	β	α	β
π_t^e	0.002	-6.257	0.002	-5.110	0.001	-11.309
	(5.276)	(-5.120)	(5.186)	(-5.202)	(5.571)	(-4.866)
π_t^{e*}	0.000	4.360	0.000	3.519	0.000	6.759
	(0.498)	(2.965)	–	(2.977)	(0.810)	(2.417)
i_t^{120}	0.000	1.000	-0.001	1.000	-0.002	1.000
	–	–	(-0.345)	–	(-1.568)	–
i_t^{120*}	0.009	-3.725	0.011	-3.082	0.000	-4.165
	(3.391)	(-4.289)	(3.310)	(-4.414)	–	(-2.521)
rer_t	-0.007	3.845	-0.008	3.114	-0.004	7.449
	(-2.251)	(4.749)	(-2.157)	(4.784)	(-2.352)	(4.837)
$C(1991:05)$	–	-21.315	–	-16.720		-37.042
		(-5.617)		(-5.481)		(-5.133)
Log likelihood	3153.093		3153.028		3148.723	
LR test	$\chi^2(1) = 0.121$		$\chi^2(1) = 0.250$		$\chi^2(1) = 8.860$	
p-value	(0.728)		(0.617)		(0.003)	
Bartlett corrected LR						
p-value						

Notes: In the VECM representation, $Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \mu + \Phi D_t + \epsilon_t$, α and β are (5×1) vectors with $\Pi = \alpha\beta'$. If Π has reduced rank, $\beta' Y_t$ represents the $p \leq r$ valid cointegrating vectors, while α contains the loadings measuring the adjustment of ΔY_t to the long run equilibrium. C(1992:04) represents the break in levels spotted by the large residuals detection procedure. For each estimate asymptotic t-ratios are reported in parenthesis. Under the restrictions in column 2 and 3 a Likelihood-Ratio test is reported at the bottom of the Table. A version of the test corrected for small sample bias is also reported in the last line (for Bartlett correction see Johansen, 2002). For sake of simplicity, it is referred to Model 2 as the VAR models where Model 2 (trend-cycle decomposition model) expected inflation appears.

Table 11: Common Trend Representation

Alpha orthogonal (transposed)					
	π_t^e	π_t^{e*}	i_t^{120}	i_t^{120*}	rer_t
CT(1)	0.000	0.000	0.000	0.722	1.000
				(1.856)	
CT(2)	1.000	0.000	0.000	-0.200	0.000
				(-2.696)	
CT(3)	0.000	1.000	0.000	-0.032	0.000
				(-0.539)	
CT(4)	0.000	0.000	1.000	0.140	0.000
				(0.412)	
Long-run impact matrix C					
	π_t^e	π_t^{e*}	i_t^{120}	i_t^{120*}	rer_t
π_t^e	3.850	0.619	0.263	-0.512	0.337
	(4.158)	(1.074)	(2.803)	(-3.949)	(2.704)
π_t^{e*}	1.173	3.197	0.087	-0.262	0.087
	(1.440)	(6.304)	(1.051)	(-2.299)	(0.796)
i_t^{120}	0.902	3.277	1.089	0.038	0.237
	(0.615)	(3.588)	(7.334)	(0.186)	(1.202)
i_t^{120*}	-2.147	3.167	0.187	0.810	0.630
	(-1.513)	(3.585)	(1.301)	(4.080)	(3.302)
rer_t	2.476	-0.173	0.138	0.291	1.056
	(1.713)	(-0.192)	(0.941)	(1.438)	(5.435)
Weakly exogenous variables (b_t, π_t^*)					
Alpha orthogonal (transposed)					
	π_t^e	π_t^{e*}	i_t^{120}	i_t^{120*}	rer_t
CT(1)	0.000	0.000	0.000	0.728	1.000
				(1.950)	
CT(2)	1.000	0.000	0.000	-0.193	0.000
				(-2.787)	
CT(3)	0.000	1.000	0.000	0.000	0.000
CT(4)	0.000	0.000	1.000	0.000	0.000
Long-run impact matrix C					
	π_t^e	π_t^{e*}	i_t^{120}	i_t^{120*}	rer_t
π_t^e	3.835	0.771	0.225	-0.478	0.362
	(4.158)	(1.331)	(2.522)	(-3.852)	(2.808)
π_t^{e*}	1.406	3.176	0.059	-0.226	0.062
	(1.751)	(6.299)	(0.758)	(-2.095)	(0.556)
i_t^{120}	0.781	3.378	1.080	0.041	0.264
	(0.535)	(3.686)	(7.638)	(0.210)	(1.296)
i_t^{120*}	-2.045	3.319	0.133	0.864	0.644
	(-1.413)	(3.652)	(0.946)	(4.438)	(3.189)
rer_t	2.504	-0.279	0.161	0.270	1.037
	(1.767)	(-0.313)	(1.174)	(1.418)	(5.239)

Notes: The common trends representation of the model is obtained by inverting the vector process (under the hypothesis of cointegration) trough the vector moving-average representation (VMA), i.e. $Y_t = \tau_0 + C^*(1)\mu_0 + C\sum_{i=1}^t \epsilon_i + C\sum_{i=1}^t D_i + C^*(L)(\epsilon_t + \Phi D_t)$, where $\tau_0 = \tau(Y_0)$ and $C = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp = \tilde{\beta}_\perp \alpha'_\perp$. The decomposition of C is similar to that of Π, with the difference that now $\tilde{\beta}_\perp$ determines the *loadings* whilst $\alpha'_\perp$ identifies the common stochastic trends. The non-stationarity of Y_t is originated indeed by the cumulative sum of the $(p-r)$ combinations of $\alpha'_\perp \sum_{i=1}^t \epsilon_i$. Asymptotic t-ratios are in parenthesis.

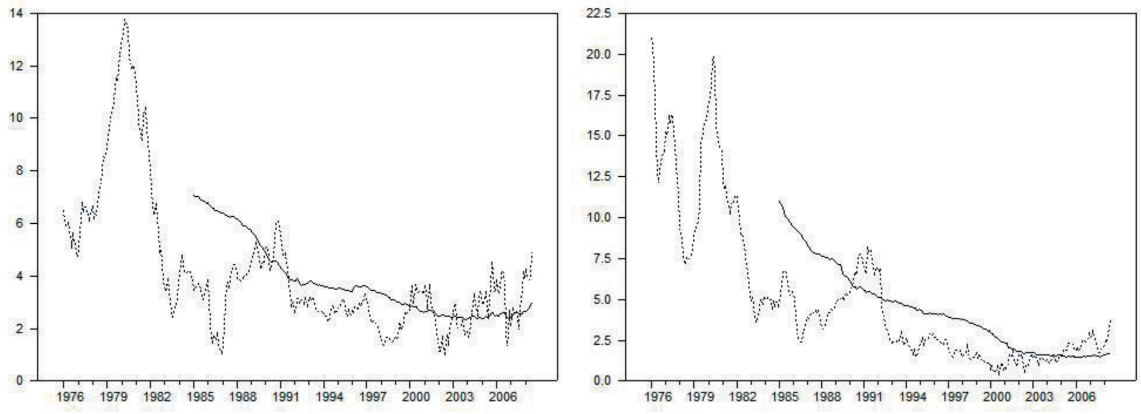

Figure 1: UK vs. US. Annualized Inflation and Realization of Inflation over 10 Years

Notes: US (left panel) and UK (right panel) plots for annualized inflation (dotted line) and realization of inflation over 10 years (black line).

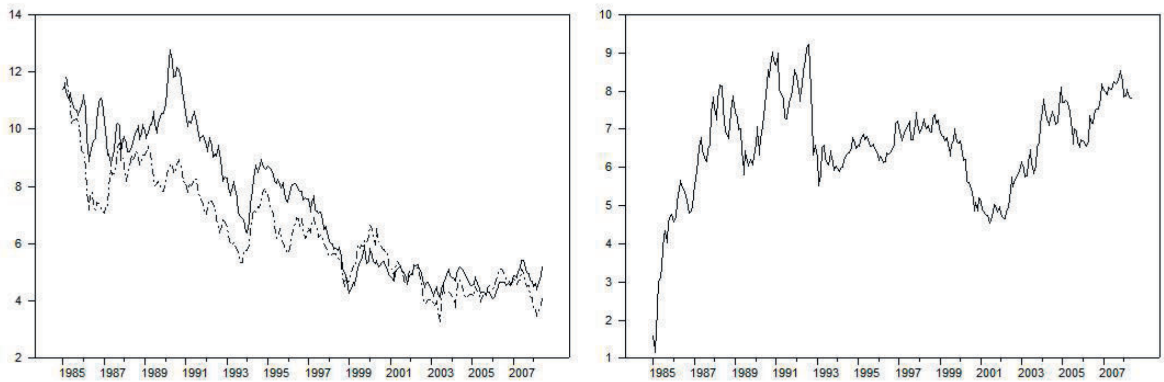

Figure 2: UK vs. US. 10 Years Constant Maturity Treasury Bond Rates and Real Exchange Rate

Notes: UK (black line) and US (broken line) 10 years constant bond-yields (left panel) and sterling/dollar bilateral real exchange rate (right panel).

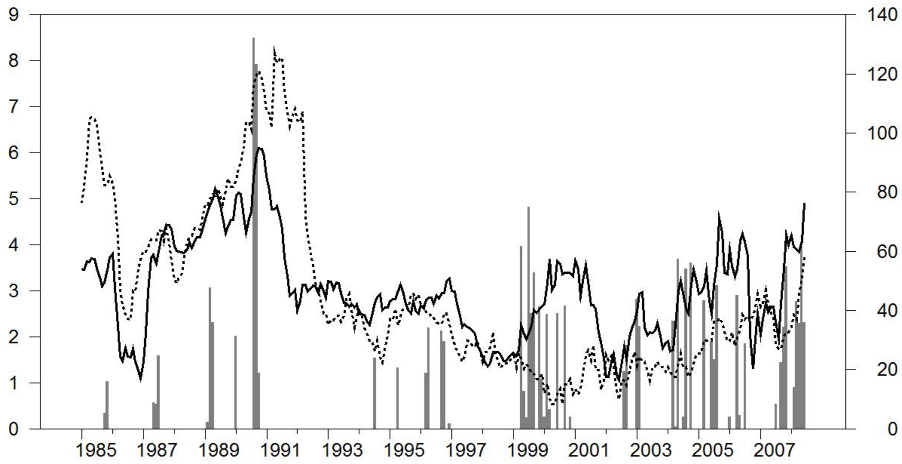

Figure 3: UK vs. US Annualized Inflation and Net Oil Price Increase

Notes: US (black line) and UK (dotted line) actual inflation and net oil price increase (see Hamilton, 1996) (right scale). The latter is constructed considering the price of oil in the current month, relative to the maximum value for the level achieved during the previous year.

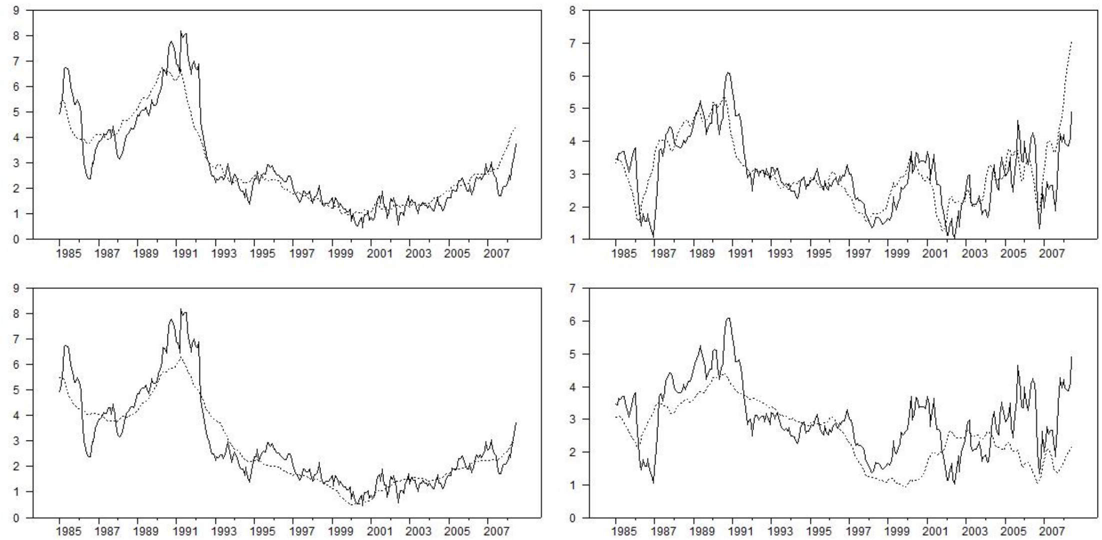

Figure 4: Modelling Inflation Results

Notes: UK (left panel) and US (right panel) results for Model 1 (upper panel) and Model 2 (lower panel) estimates. Actual inflation (black line) is plotted together with the expected inflation (dotted line). Model 1 is a simple permanent-transitory decomposition for inflation; Model 2 augment the measurement equation in Model 1 with two exogenous regressors (the economy output gap and a net oil price increase *à la* Hamilton (1996)).

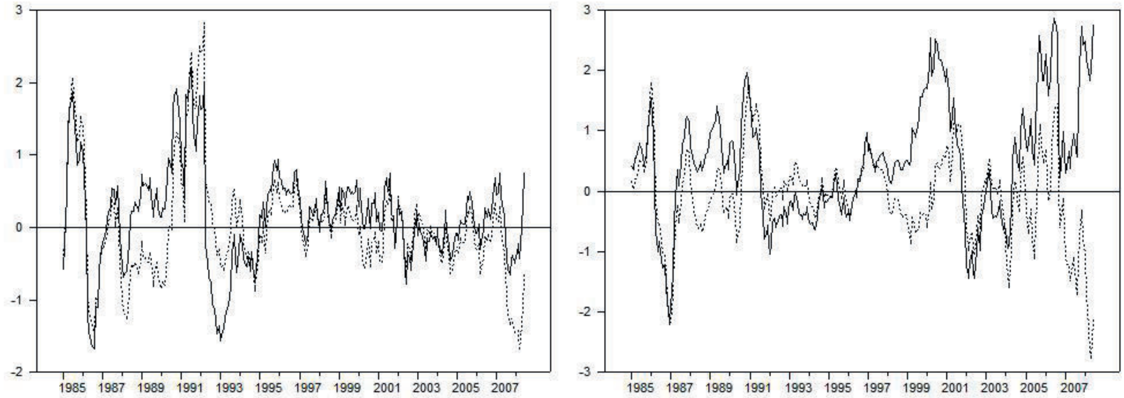

Figure 5: Estimated Inflation Gaps

Notes: UK (left panel) and US (right panel) results for Model 1 (broken line) and Model 2 (black line). Inflation gaps are computed as $\omega_t = \pi_t - \pi_t^e$. Model 1 is a simple permanent-transitory decomposition for inflation; Model 2 augment the measurement equation in Model 1 with two exogenous regressors (the economy output gap and a net oil price increase *à la* Hamilton (1996)).

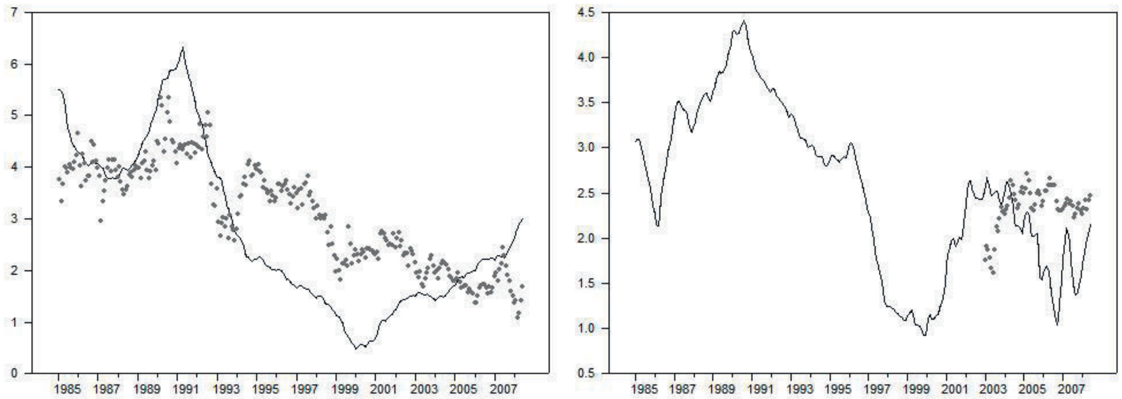

Figure 6: Model 2 Expected Inflation and Financial Markets Expectation on Yield from Government Securities at 10 years Inflation

Notes: UK (left panel) and US (right panel) results for Model 2 (black line). Dots represent inflationary expectations (plus forecast error) extracted from financial markets information. Model 2 is a trend-cycle decomposition model using the economy output gap and a net oil price increase *à la* Hamilton (1996) as exogenous regressors.

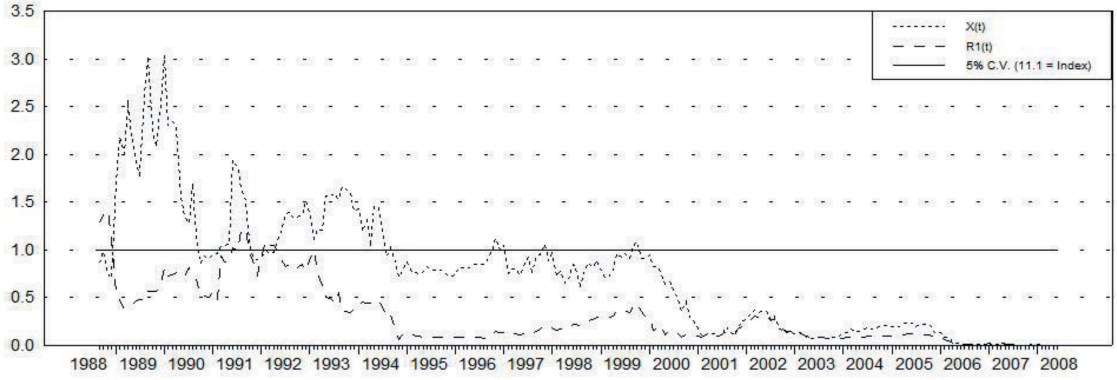

Figure 7: Test of Beta Constancy

Notes: The results illustrate a test based on the constancy of the cointegrating space $\beta' Y_t$. This shows how, under the null of the constancy of the parameters, the R-form is safely below the rejection line of 1.0 for almost all t belonging to the sample period. As indicated in the picture, recursive estimates refer to 5% critical level. The base sample for the recursive estimates is from 1986:03 to 1989:08. In each sub sample the short run parameters are re-estimated. Similar results are obtained if short run parameters are fixed at their full sample estimates. The results refer to the model augmented with inflationary expectations.