BANK OF ENGLAND

External MPC Unit

Discussion Paper No. 30

Macroeconomic stability and the real interest rate: a cross-country analysis

Charlotta Groth and Fabrizio Zampolli

September 2010

This document is written by the External MPC Unit for the Bank of England

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Macroeconomic stability and the real interest rate: a cross-country analysis

Charlotta Groth[(1)] and Fabrizio Zampolli[(2)]

Abstract

We construct a measure of the short-term world interest rate using principal component analysis. Drawing on real interest rate data for 18 OECD countries for the period 1985–2008, persistent deviations from the world interest rate that cannot be explained by movements in the real exchange rate are documented. A theoretical model predicts that these unexplained deviations capture foreign exchange rate risk premia. Using panel data techniques, we test the theoretical prediction that a rise in conditional consumption growth volatility relative to the rest of the world reduces the foreign exchange rate risk premia and, therefore, the real interest rate. Our main result is that we find a robust and signicant negative relation between the volatility in consumption growth and the level of real interest rates relative to the world interest rate, supporting this hypothesis. We also look at the relation between real interest rates and the net foreign asset position. We test the hypothesis that the empirical negative relation between the two variables captures the relation between real interest rates and macroeconomic volatility, on the one hand, and macroeconomic volatility and the net foreign asset position, on the other hand. We are not able to reject this hypothesis.

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The views expressed in this paper are those of the authors, and not necessarily those of the Bank of England. The authors wish to thank Kate Barker for valuable comments and suggestions.

Information on the External MPC Unit Discussion Papers can be found at
www.bankofengland.co.uk/publications/externalmpcpapers/index.htm

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ISSN 1748-6203 (on-line)

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Summary

A simple inspection of the data suggests that there are persistent differences in real interest rates across countries even at short maturities. Over the past 20 years, for example, a measure of the short-term real interest rate has been almost 300 basis points higher in the United Kingdom than in the United States despite similar economic developments. Moreover, in a sample of advanced OECD countries, real interest rates for the highest quartile of countries have, on average, been over 350 basis points higher than those for the lowest quartile of countries. These differences are difficult to explain; the dismantling of barriers to international movements of capital and goods which has taken place over the past means that we should expect an equalisation of real interest rate across countries, at least in the long run. Since persistent interest rate differentials are likely to affect firms' and households' borrowing conditions across countries, they could also impact on economic activity. Understanding what causes these differences in interest rates is therefore of interest, and in this paper we set out to explore this by estimating an empirical model using a cross-country panel data set.

We construct short-term real interest rates for a sample of 18 OECD countries over the period 1985-2008. Based on this data, a measure of the world interest rate is constructed using principal component analysis. We establish the fact that in many countries, real interest rates have deviated from the world interest rate for long periods of time. Persistent deviations should be expected if the price of nontraded goods relative to traded goods move differently across countries, and the share of nontraded goods in consumption expenditure is large. We try to control for this by constructing a measure of the real interest rate mainly based on tradeables, finding that the inclusion of non-tradeables account for some of the differences in real interest rates across countries, but a large part remain unexplained.

We next argue that these unexplained interest rate differentials are likely to reflect risk premia. A standard asset-pricing framework is used to derive a structural equation for the exchange rate risk premium. One prediction from this model that has not previously been tested empirically is that real interest rate differentials across countries should be negatively related to a measure of relative volatility in consumption growth. Taking this hypothesis to the data is the main contribution of the paper. Our focus on the shorter end of the yield curve for a sample of

developed countries with good credit ratings over the sample period means that both inflation risks and the risk of default are likely to be small. This, we argue, allows us to interpret unexplained movements in interest rate differentials in terms of foreign exchange rate risk premia.

Using panel data techniques, we estimate an empirical UIP condition relating the interest rate differential relative to the world interest rate to the expected change in the real exchange rate, and a measure of relative consumption volatility. Our main result is that, consistent with theory, there is a significant negative relation between real interest rate differentials and the volatility of both output and private consumption growth, once we control for expected exchange rate changes. In other words, countries in which economic volatility is high tend to have lower real interest rates, once we control for expected exchange rate changes. This result is robust to different methods for proxying for expectations of exchange rate changes. The estimates imply that a percentage rise in the volatility of output growth reduces the real short-term interest rate relative to the world interest rate by 0.004 percent. We also show that large movements in economic volatility over the past means that the impact of volatility on real interest rates could have been significant in some countries.

We finally tentatively explore the empirical relationship between real interest rates and the net foreign asset position. To analyse this, we draw on theoretical work that show that the net foreign asset position of a country is positively affected by the strength of its precautionary savings motive relative to that of its trading partners, where the precautionary savings motive is positively related to economic volatility. Taken together, this means that theory suggests a *positive* relation between the net foreign asset position and economic volatility. As discussed above, asset-pricing theories predict a *negative* relation between economic volatility and the interest rate differential. Altogether, this suggests a negative relation between real interest rates and the net foreign asset position.

In this paper, we postulate the hypothesis that the documented negative empirical relation between real interest rates and the net foreign asset position is a reduced-form relation, capturing the links described above. We are not able to reject this hypothesis.

1 Introduction

A simple inspection of the data suggests that there are persistent differences in real interest rates across countries even at short maturities. Over the last 20 years, for example, a measure of the short-term real interest rate has been almost 300 basis points higher in the United Kingdom than in the United States despite similar economic developments. Moreover, in a sample of advanced OECD countries, real interest rates for the highest quartile of countries have, on average, been over 350 basis points higher than those for the lowest quartile of countries. These differences are difficult to explain; the dismantling of barriers to international movements of capital and goods which has taken place over the past means that we should expect an equalisation of real interest rate across countries, at least in the long run. Since persistent interest rate differentials are likely to affect firms' and households' borrowing conditions across countries, they could also impact on real activity. Understanding what causes these differences in interest rates is therefore of interest, and in this paper we set out to explore this by estimating an empirical model using a cross-country panel data set.

We start by documenting differences in the levels of real interest rates in a sample of 18 OECD countries over a period of 23 years (1985-2008), and set out to analyse what factors may have caused them to deviate persistently from each other. A main characteristic of our analysis is that we use a measure of the world interest rate as a benchmark rather than the interest rate of a base country. This has two advantages: since we are interested in how the level of interest rates varies across countries, it is more informative to compare cross-country real interest rates to a measure of the world interest rate than to that of an arbitrary base country; estimating bilateral relations will also make the results more sensitive to anomalies in the exchange rate market for the currency of the base country. By focusing on the world interest rate, we mitigate this problem.

In the first part of the paper, we estimate a measure of the world interest rate, using a principal component analysis. This is a flexible approach which allows for different weights be assigned to different countries in the construction of a world interest rate. We argue that this is a necessary requirement, given large cross-country differences in size, wealth and contribution to world trade.

We next conduct an empirical analysis of the determinants of cross-country interest rate differentials, where we base our analysis on the uncovered interest rate parity (UIP) condition.

This framework highlights that persistent interest rate differentials could reflect persistent movements in real exchange rates and risk premia. These risk premia reflect the extra compensation that investors require for taking on different types of risks, such as inflation risk, foreign exchange rate risk and default risk. As our focus is on the foreign exchange rate risk premia, we limit our study to the very short end of the yield curve, which is less likely to be affected by inflation risk premia. In addition, we include only developed countries with relatively high credit ratings in our sample, which allows us to mitigate the size of default risk premia.

We use a standard asset-pricing framework based on Backus, Foresi and Telmer (2001) to derive a structural equation for the foreign exchange rate risk premium. Theory predicts that the risk premium is negatively related to the volatility of consumption growth in the home country relative to abroad. A rise in domestic consumption growth volatility should therefore reduce real interest rates relative to abroad. The intuition is that a rise in economic volatility strengthens the motives to engage in precautionary savings, which puts downward pressure on equilibrium interest rates.

Our paper is related to earlier empirical work that explains cross-country differences in expected returns in terms of the risk premium. Harvey, Solnik and Zhou (1994) use a factor model to analyse interest rates across countries, and find that the first factor resembles a measure of the return on a world market portfolio, and that there is evidence that the second factor is related to foreign exchange rate risk. Sarkissian (2003) finds that consumption dispersion across countries provides some explanatory power for the differences in expected excess returns. Lustig and Verdelhan (2007) document that aggregate consumption growth risk goes some way in explaining changes in the exchange rate, conditional on the foreign interest rate, and that high interest rate currencies relative to the United States tend to depreciate when US consumption growth is low.

One theoretical prediction that has yet to be tested empirically, however, is that cross-country differences in real interest rates are related to a measure of relative volatility in consumption growth. Taking this hypothesis to the data is the main contribution of this paper.

Turning to the results, our estimate of the world interest rate explains almost 70 percent of the volatility in the short-term (3-month) real interest rate in our sample of 18 OECD countries for the period 1985-2007. Moreover, it is positively related to all countries' interest rates, thus suggesting that there is an important role for international factors in shaping real interest rates in open economies. There are nevertheless persistent deviations of interest rates from the world interest rate which cannot be explained by movements in the real exchange rate. Using panel data techniques, we estimate an empirical UIP condition relating the interest rate differential (relative to the world interest rate) to the expected change in the real exchange rate, and a measure of relative consumption volatility. Our main result is that, consistent with theory, there is a significant negative relation between real interest rate differentials and the volatility of both output and private consumption growth. This result is robust to alternative methods for proxying expected changes in future exchange rates.

The empirical estimates imply that a percentage rise in the volatility of output growth reduces the real short-term interest rate relative to the world interest rate by 0.004 percent. To put this number in context, we look at a subset of countries where output growth volatility diminished (the United Kingdom, New Zealand, Sweden and Spain) markedly over the sample period. In these countries, output volatility fell on average by 60 percent between the first and the second sub-sample, while the interest rate differential rose by 1 percent. Given our estimate of the elasticity parameter, the model predicts that around one quarter of that rise could be accounted for by falling volatility. Hence, although the elasticity parameter is small, large movements in economic volatility over the past means that the impact on real interest rates could have been significant.

Another way to state this is that, over the past, interest rates have tended to be a little higher in countries such as the United Kingdom, Spain, New Zealand and Sweden, compared to those prevailing in other countries. One plausible explanation for this difference is the relative economic stability that these countries have enjoyed over the past decade or so. Everything else equal, this stability has tended to depress precautionary savings relative to the rest of the world, which has put upward pressure on real interest rates.

We also tentatively explore the empirical relationship between real interest rates and net foreign asset (NFA) positions. Previous empirical work (Lane and Milesi-Ferretti (2001), Selaive and Tuesta (2003)) have documented a significant negative relationship between these two variables. It is not clear, however, what is driving this result. To analyse this, we draw on the literature that studies how uninsurable aggregate risk, through its impact on precautionary savings, affect the external balance of a country. Durdu et al (2008) and Fogli and Perri (2006) find that increased business cycle volatility strengthens the motive for precautionary savings, and therefore an improvement of the net foreign asset position.[1]

We postulate the hypothesis that the documented negative empirical relation between real interest rates and the NFA position is a reduced-form relation, capturing the *negative* relation between real interest rates and economic volatility, on the one hand, and the *positive* relation between economic volatility and the net foreign asset position, on the other hand. We are not able to reject this hypothesis.

The rest of the paper is organised as follows. Section 2 describes real interest rates, constructs the world interest rate using a principal component analysis, and discusses deviations of real interest rates from the world rate. Section 3 models the risk premia, implements the model empirically, and describes the data and the estimation methods. Section 4 discusses the main results. Section 5 explores the link between interest rate differentials and the net foreign asset position. Section 6 concludes.

2 Real interest rate behaviour

2.1 Real rates

The real interest rate in country i is defined from the Fisher equation as

$$rr_{it} = i_{it} - \mathrm{E}_t\pi_{t+1}$$

where rr_{it} is the one-period *ex ante* real interest rate, i_{it} is the nominal interest rate earned on a one-period bond that matures in period $t+1$, $\mathrm{E}_t\pi_{t+1}$ is the period-t expectation of consumer price inflation from period t to $t+1$. The *ex ante* real interest rate is not directly observable. We instead use a measure of the *ex post* real rate, r_{it}, which is defined according to

$$r_{it} = i_{it} - \pi_{t+1} = rr_{it} + \varepsilon_{it}^{\pi} \qquad \textbf{(1)}$$

where ε_{it}^{π} is the forecast error of inflation, defined as $\varepsilon_{it}^{\pi} = \mathrm{E}_t\pi_{t+1} - \pi_{t+1}$. Below, we assume that expectations are rational, which implies that ε_{it}^{π} is not predictable.

[1]Related recent work also include Bems and Carvalho Filho (2009) and Carroll and Jeanne (2009).

We construct a measure of the short-term ex post interest rate for a sample of 18 OECD countries over the period 1985 Q3 – 2008 Q2 using (1), where nominal interest rates are taken from 3-month Treasury bills, and the measure of inflation is the annual consumer price inflation rate over the duration of the bill.[2] Table 1 provides the basic sample statistics for the short-term real interest rates across the countries in our sample.[3]

Countries with higher than average interest rates include the United Kingdom, New Zealand, Australia, Norway and Denmark (column 2). Those with low interest rates include the United States, Japan and Switzerland. However, mean rates may be affected by episodes in which real interest rates have been unusually high or low, which translate into skewed distributions. The median may therefore provide a better handle on the level around which interest rates have normally tended to fluctuate. Based on median values (column 3), real interest rates have been high in United Kingdom, New Zealand, Norway, Ireland and Italy, where the median interest rates where, on average, over 4 percent over the sample period. This compares to low interest-rate countries (Germany, Japan, Switzerland and the United States) where the median interest rate was close to 1.5 percent over the same period. This confirms that there are large differences in real interest rates across countries, both in terms of mean and median.

Another interesting feature of the data is their dispersion as measured by the standard deviation and the inter-quartile range (column 4, 7 and 10). Countries that have had more stable rates include Germany, Switzerland, the UK and the United States. With the exception of the United Kingdom, these are countries that also have had low interest rates.

2.2 The world interest rate

We next construct a measure of the world interest rate using principal component analysis. The idea behind this analysis is to represent the variability of the data with a smaller number of new variables built in such a way that they retain as much as possible of the variation in the original data. These new variables, uncorrelated with each other, are the principal components (PCs). The first PC is constructed as a weighted sum of the real interest rates in our sample countries, where the weights are chosen to explain the largest degree of variation in the data; the second PC

[2] The sample period has been chosen to include the largest possible number of countries.

[3] The data is taken from Global Financial Data.

explains the second largest amount of variation, and so fourth in descending order.[4] One advantage of this analysis is that it focuses on the variability of interest rates, rather than their co-movements.[5] We argue that this is appropriate for the analysis of the level of interest rates. And since the focus is on the level of real interest rates, we choose to conduct the PC analysis on the original data series. However, as there are large variations in the volatility of real interest rates across countries, countries with very volatile interest rates are likely to dominate the first few PCs. We therefore also conduct the PC analysis for standardised data.

The 1st PC is a common factor in that all coefficients have the same sign, capturing 68% of the total variance (column 2-4 in Table 2). However, not all countries contribute in the same way to the common factor. In particular, the coefficient associated with the United States is relatively small suggesting that the US interest rate does not vary much with other countries.[6]

We also note that those countries where interest rates are most volatile (Denmark, Ireland, Italy and Spain) tend to dominate the first PC, whereas countries with relatively low volatility, such as Germany, Switzerland and the United Kingdom, as well as the United States, tend to have less weight in the PC. We therefore also estimate the PCs using standardised data, as shown in Column 5-7 in Table 2. The results from the standardised data support the interpretation of the first PC as a world interest rate, with all coefficients having the same sign, and of similar magnitude, ranging from 0.16 to 0.28. The standardised PC also confirms the result that the US interest rate is relatively less correlated with interest rates in other countries, with a coefficient of 0.17.[7] This is lower than the coefficients of the other countries with low interest rate volatility, which display a weight of 0.20 or above. The 2nd PC appears to show a common pattern of behaviour across Anglo-Saxon countries, which all have large positive values. The remaining principal components individually contribute only to a small share of the total variance without having a straightforward or intuitive interpretation.

Given the ability of the first principal component in explaining a large share of the variance in real interest rates across countries, we take it as an estimate of the world real interest rate and

[4]For a precise and technical definition of principal component analysis see e.g. I.T. Jolliffe (1986).

[5]This would be the case in a factor analysis approach, as in Brzoza-Brzezina and Cuaresma (2008).

[6]This is in line with results by earlier results by Gagnon and Unferth (1995), Chinn and Frankel (2004) and Brzoza-Brzezina and Cuaresma (2008), who find that US rates tend to deviate long periods of time from a measure of world interest rate.

[7]The only country with a lower weight than the United States is Portugal.

hence as a benchmark to assess how different countries' real interest rates differ from each other. The world rate is constructed as a weighted average of the sample countries' real interest rates, where the relative weights are the ones reported in column 2 (for non-standardised data) and 5 (for standardised data) in Table 2.[8] As shown in Figure 1, the two estimates are highly correlated. We therefore focus on the non-standardised measure in the remaining of the paper.

The last row in Table 1 gives the summary statistics for the world interest rate, and Table 3 gives the corresponding correlation coefficients, for the whole sample period, and for the two sub-periods prior to and after 1997 Q1. Figure 2 plots each country's real interest rate together with the estimate of the world interest rate.

Real interest rates appear to be characterized by a break around 1997. Before this date rates were more volatile and on average higher than post 1997. This is reflected both in the world interest rate and across countries. We also find that, in the latter period, a greater number of countries appear to have tracked the movements in the world rate more closely (Table 3). The closer correlation is particularly true for the United States, where the correlation with the estimated world interest rate was close to zero prior to 1997, but strongly positively thereafter. For Japan, the opposite holds, possibly reflecting country-specific developments since the mid-1990s.

As shown in Figure 2, the rates of United States and Japan have been below the world rate for most of the sample period. The fact that these two countries are the two largest and also the relatively most closed economies could explain why their respective real interest rates are less synchronised with the world rate. However, this does not explain why rates should be lower. One explanation is related to the role of the dollar as the reserve currency of the world, which means that international investors are prepared to hold it at a lower return. Other countries with low interest rates include Switzerland and, until the inception of the Euro Area, Germany (the notable exception being the time of the German reunification).

The behaviour of real interest rates in some of the other countries that joined the Euro area seems intuitive. As an example, Ireland's real rate was consistently above the world rate in the pre-Euro period. Subsequently the real rate has always been in line or below the world rate, reflecting the impossibility of a nominal depreciation. The same change has occurred in Portugal and Spain.

[8]Compared to a simple average, the weights do not sum to one.

Finally, the rates of UK, Australia, New Zealand and Norway have been consistently above the world interest rates since 1997.

2.3 Deviations from the world interest rate

The previous section established that there are significant differences in real interest rates across countries, and there are large and persistent deviations from our estimate of the world interest rate. To account for sustained interest rate differentials across countries, we start by postulating a standard augmented interest parity relationship that would hold under the assumption of perfect capital markets:

$$i_{it} = i_{jt} + \mathrm{E}_t \Delta s_{ij,t+1} + \nu_{it} \qquad \textbf{(2)}$$

where i_{it} and i_{jt} are the nominal gross interest rates of country i and j in period t, where $i, j \in 1, ..., N$, $\mathrm{E}_t \Delta s_{ij,t+1}$ is the expected nominal depreciation of the currency in country i relative to the currency in country j (an increase in s_{ij} implies a depreciation of the currency in country i) and ν_{it} is a foreign exchange rate risk premium. All variables (except for the foreign exchange rate risk premium) are denoted in logs. The Fisher parity in country i is given by

$$r_{it} = i_{it} - \mathrm{E}_t \left[\pi_{i,t+1}\right] + \eta_{it} \qquad \textbf{(3)}$$

where η_{it} is the inflation risk premium and $\pi_{i,t+1}$ is a measure of consumer price inflation**.** Combining (2) and (3) gives

$$r_{it} = r_{jt} + \mathrm{E}_t \left[\pi_{j,t+1} - \pi_{i,t+1}\right] + \mathrm{E}_t \Delta s_{ij,t+1} + \left(\eta_{it} - \eta_{jt}\right) + \nu_{it} \qquad \textbf{(4)}$$

We further allow for pricing to market, home bias in consumption, and both traded and non-traded goods. This means that (ex ante) purchasing power parity may not hold. Instead we have

$$q_{ij,t} = s_{ij,t} + p_{jt} - p_{it} \qquad \textbf{(5)}$$

where p_{it} and p_{jt} are the price level in country i and j, with $q_{ij,t}$ being the real exchange rate between country i and country j. Together with (4) we get

$$r_{it} = r_{jt} + \mathrm{E}_t \Delta q_{ij,t+1} + \nu_{it} + \left(\eta_{it} - \eta_{jt}\right) \qquad \textbf{(6)}$$

Thus, interest rate differentials across countries are related to expected changes in the real exchange rate, to the foreign exchange risk premium, and to the difference in the inflation risk premium in the two countries. Given that we focus on the short-term real interest rate in low-inflation countries, it is unlikely that the inflation risk premium has an important role for interest rate differentials. Below, we therefore assume that $\eta_{it} = \eta_{jt} = 0$.

Assume for now that the risk premium υ_{it} in **(6)** is zero. Given N countries, there are $N\,(N-1)$ equations of the type given in **(6)**, but only $N-1$ of them are independent. Previous studies therefore typically estimate a system of $N-1$ bilateral relations expressed relative to the base country k, often chosen to be the US.[9] Here we instead express the interest rate differential relative to the world interest rate for the N countries in the sample, and relate this to the change in the real effective exchange rate:

$$r_{it} = r_t^w + \mathrm{E}_t\,\Delta q_{i,t+1},\ i \in 1, ..., N \tag{7}$$

where r_t^w is a measure of the world interest rate and $q_{i,t}$ is the real effective exchange rate for country i defined in terms of a basket of currencies. It can be shown that a system of N conditions of the type above is approximately equivalent to a system of $N-1$ bilateral relations **(6)**.[10]

There are two main reasons for conducting the analysis in terms of the world interest rate, rather than focusing on the bilateral relations. First, we are interested in how the level of real interest rates varies across countries. For this analysis, it is more informative to compare cross-country real interest rates to a measure of the world interest rate which equates investment and savings at the global level, than to that of an arbitrary base country. Second, estimating bilateral relations will make the results more sensitive to anomalies in the exchange rate market for the currency of the base country.

As a starting point for our analysis of interest rate differentials across countries, we look at the case in which agents are risk neutral and PPP holds for all goods. From **(7)** this implies that real interest rate parity holds: $r_{it} = r_t^w$, for all $i \in 1, ..., N$. To test this relation, we estimate the following equation jointly for all countries:

$$\tilde{r}_{it} = \alpha_i + \varepsilon_{it},\ i = 1, ..., N \tag{8}$$

where $\tilde{r}_{it} = r_{it} - r_t^w$, where r_{it} is the real interest rate of country i at time t.[11] As reported in column 2 in Table 4, the estimated parameter α_i is significant for all countries except for Canada

[9] The literature testing the UIP condition is large. Recent studies include Chinn and Meredith (2005), and earlier contributions by Edison and Pauls (1993) and Baxter (1993).

[10] Had the real exchange rate been constructed using PC weights instead of trade-weights, this would hold exactly.

[11] The equation is estimated jointly for all countries, using generalised least squares (GLS).

and Ireland. This implies that we reject the hypothesis of real interest rate parity across countries. In particular, the rates of the UK, New Zealand, Australia, Denmark, and Norway have on average been above the world interest rate, while the rates of US, Japan, Germany and Switzerland have been markedly below it for most of the sample period.

We would expect deviations from real interest rate parity to be important if the price of nontraded goods relative to traded goods move differently across countries, and the share of nontraded goods in consumption expenditure is large.[12] To account for this, we estimate **(8)** using an alternative measure of the real interest rate. This is constructed using a measure of inflation which is less affected by the inclusion of non-traded goods, based on the wholesale price index (WPI).[13] As shown in column 2 in Table 4, the estimated parameter α_i is in this case significant in only around two thirds of the countries. This suggests that part of the difference in real interest rates across countries reflects the presence of non-traded goods in the consumer price index. However, for a majority of countries, there are still significant deviations from PPP.

In its weak form, real interest rate parity allows for constant risk premia which means that although rates are not equalised across countries they move similarly across time. To analyse this, we look at the properties of the residuals ε_{it} in **(8)**. Rejecting the hypothesis of white noise residuals would imply that there are persistent deviations of real interest rates from the world interest rate, even when allowing for constant term premia. Based on the Ljung-Box Portmanteau test for serial correlation in the residual, we find that for all countries we reject the null hypothesis of no serial correlation at the 5% level, for both the CPI- and the WPI based measure of the real interest rate.[14] To abstract from relatively short-term deviations from the world interest rate, we reestimate **(8)** using annual data:

$$\tilde{r}_{it}^{a} = \alpha_i^{a} + \varepsilon_{it}^{a}, \; i = 1, ..., N \tag{9}$$

where $\tilde{r}_{it}^{a} = r_{it}^{a} - r_t^{wa}$ is an annualised measure of the deviation of real interest rates in country i in year t from the world interest rate, where the annualised interest rate data is calculated as the 4-quarter average of the quarterly interest rate data. Column 2-4 in Table 5 reports the Q statistics

[12] For an analysis with barriers to trade and tradeables and non-tradeable goods, see eg Obstfeld and Rogoff (2000), Dutton (1993) and Dutton and Strauss (1997).

[13] WPI data for Portugal is not available. Portugal is therefore excluded from the analysis based on WPI inflation.

[14] This is in contrast to findings by Gagnon (1995) who create a similar measure of the residual for a sample of 9 OECD countries the period 1978-1993, and cannot reject the null hypothesis of no serial correlation for a majority of the countries. However, Gagnon uses yearly data, whereas we focus on quarterly data.

at lag lengths 1 2 and 3 for the residual ε^a_{it}. Based on annualised data, in roughly two thirds of the countries we reject the null hypothesis that the deviation of interest rates from the world rate is white noise, once we account for a constant term premium. Column 5-7 reports the same test statistics for the WPI-based real interest rate. As expected, the null is rejected for fewer (in around one half of the) countries. This means that there remain significant interest rate differentials that cannot be explained only by the existence of constant term premia, and that persist also when we use a measure of inflation less affected by non-traded goods.

3 Empirical model

Having established that real interest rates across countries differ from the world interest rate for long periods of time, we next turn to an empirical analysis of these deviations.

3.1 Modelling risk premia

Equation (7) reflects the fact that, even under the assumption that the risk premia are zero, we would expect deviations from real interest rate parity to be important when there are persistent deviations from PPP. As was clear from the analysis above, also when we try to control for this using a measure of the real interest rate based on the WPI, there remain persistent interest rate differentials across countries. Analysing these therefore requires some modelling of the risk premium in (6).

We take as our starting point the standard consumption-based asset pricing model, adopted to include an intertemporal price for currencies. Under the assumptions of no arbitrage opportunities and complete financial markets, the investor's Euler condition for any asset can be represented as:

$$1 = \mathrm{E}_t \left[M_{t+1} R_t \right] \tag{10}$$

where R_t is the gross real one-period return on the asset and M_{t+1} is a stochastic discount factor:

$$M_{t+1} = \frac{U'(C_{t+1})}{U'(C_t)}$$

where $U'(C_t)$ denotes the marginal utility of consumption. A similar relation holds in the foreign country (denoted by *):

$$1 = \mathrm{E}_t \left[M^*_{t+1} R^*_t \right] \tag{11}$$

We could use arbitrage conditions to express this as

$$1 = \mathrm{E}_t \left[M_{t+1} \frac{Q_{t+1}}{Q_t} R_t^* \right] \tag{12}$$

where Q_t denotes the real exchange rate. Combining the last two equations give

$$\mathrm{E}_t \left[M_{t+1}^* R_{t+1}^* \right] = \mathrm{E}_t \left[M_{t+1} \frac{Q_{t+1}}{Q_t} R_t^* \right] \tag{13}$$

Under the assumption of complete markets, the solution to this equation fulfils:[15]

$$\frac{Q_{t+1}}{Q_t} = \frac{M_{t+1}^*}{M_{t+1}} \tag{14}$$

This equation links the returns R_{t+1} and R_{t+1}^* by specifying a relation for the exchange rate depreciation.[16] As is shown in Appendix A, under the assumption of log-normal distributions of M_{t+1} and M_{t+1}^*, a second-order approximation of **(10)** and **(11)** give rise to the following equations for the real interest rate

$$r_t = -\mathrm{E}_t m_{t+1} - \frac{1}{2} var_t \left(m_{t+1} \right), \tag{15}$$

$$r_t^* = -\mathrm{E}_t m_{t+1}^* - \frac{1}{2} var_t \left(m_{t+1}^* \right) \tag{16}$$

where r_t and m_{t+1} are the log of R_t and M_{t+1}, and $var_t \left(m_{t+1} \right)$ is the conditional variance of m_{t+1}. Equation **(15)** implies that when individuals expect the marginal utility of consumption to be high in the future relative to the current period, real interest rates are low. The reason for this is that a lower rate of interest is required to equate savings and investment. The conditional variance of m_{t+1} also has a negative impact on real interest rates. Uncertainty about the future increases individuals' incentives to engage in precautionary savings, which puts downward pressure on real interest rates.

Combining **(15)-(16)** with **(14)** gives

$$r_t = r_t^* + \mathrm{E}_t \Delta q_{t+1} + \frac{1}{2} \left[var_t \left(m_{t+1}^* \right) - var_t \left(m_{t+1} \right) \right] \tag{17}$$

Hence the risk premium is affected by the volatility of the stochastic discount factor abroad relative to home, through its impact on the incentives to engage in precautionary savings. To better understand the intuition behind this result, we show in the Appendix that the risk premium

[15]This solution is unique when markets are complete such that there is a complete set of currencies and state-contingent claims. When markets are not complete, this solution is not unique, but still constitutes one possible solution (see Backus, Foresi and Telmer (2001) and Chari, Kehoe and McGrattan (2002)).

[16]As stressed by Backus, Foresi and Telmer (2001), this relation states that out of three variables - M_{t+1}^*, M_{t+1} and Q_{t+1}/Q_t, one is effectively redundant and can be constructed from the other two. Most of the existing literature focuses on the domestic stochastic discount factor and the depreciation rate. We instead follow Backus, Foresi and Telmer (2001) and recent applications by Verdelhan (2008) and de Paoli and Sondergaard (2008), in focusing on the two SDFs.

can alternatively be expressed as:[17]

$$\upsilon_t = \frac{1}{2}var_t\left(q_{t+1}\right) + cov_t\left(m_{t+1}, q_{t+1}\right) \qquad \textbf{(18)}$$

The risk premium is related to the variance of the exchange rate and the covariance between the real exchange rate and the domestic stochastic discount factor. When this covariance is positive, the exchange rate is expected to depreciate when the marginal utility of consumption is high. This implies that the domestic-currency return to foreign investment is high when the utility of an extra unit of consumption is high. For a given level of exchange rate volatility ($var_t\left(q_{t+1}\right)$), a positive covariance between m_{t+1} and q_{t+1} therefore puts downwards pressure on the foreign interest rate r_t^*, and causes the interest rate differential to rise. We show in the appendix that, by solving for the volatility and covariance of the exchange rate in terms of the domestic and foreign stochastic discount factors, we obtain **(17)**. Hence, the two ways of modelling the risk premium are linked.

Under some assumptions about the utility function, we have:[18]

$$var_t\left(m_{t+1}\right) = \gamma^2 var_t\left[c_{t+1} - c_t\right] \qquad \textbf{(19)}$$

where c_t is the log of consumption and γ is the inverse of the intertemporal elasticity of substitution. Under the assumption that γ is equal across countries the risk premium will be negatively affected by the conditional volatility of consumption growth in the home country relative to abroad.[19] As discussed above, the intuition is that when the economy is more volatile, individuals have an incentive to build up precautionary savings, which puts downward pressure on real interest rates, through the impact on the risk premia.

Equation **(17)** implies that, in a sample of N countries, there exists $N - 1$ independent relations of the following form

$$r_{it} = r_{jt} + \mathrm{E}_t\Delta q_{i,t+1} - \frac{1}{2}var_t\left(m_{i,t+1}\right) + \frac{1}{2}var_t\left(m_{j,t+1}\right), \qquad i = 1, ..., N-1 \qquad \textbf{(20)}$$

[17]This follows the approach by eg Lustig and Verdelhan (2007) and Lewis (1999).

[18]We need constant relative risk aversion and no habits in consumption.
Previous work have introduced habits in consumption to motivate time-varying risk premia to explain the apparent puzzle in foreign exchange markets (Verdelhan (2008) and de Paoli and Sondergaard (2008)). Since we are not interested in addressing that puzzle, per se, we assume a simpler utility function.

[19]Under the assumption that shocks to consumption growth are iid distributed, we have $c_{t+1} - c_t = g + u_{t+1},\ \ u_{t+1} \sim N\left(0, \sigma^2\right)$ where g is trend growth, in which case the conditional volatility of consumption growth will be proportional to the underlying shock, σ^2.

for any country $i \neq j$. This is approximately equal to a system of N relations expressed in terms of deviations from the world interest rate:

$$\tilde{r}_{it} = \mathrm{E}_t \Delta q_{i,t+1} - \frac{1}{2} var_t \left(m_{i,t+1} \right) + \frac{1}{2} var_t \left(m^{w}_{t+1} \right), \qquad i = 1, ..., N \tag{21}$$

where $var_t \left(m^{w}_{t+1} \right)$ is a measure of the volatility of the stochastic discount factor in the rest of the world.

3.2 Empirical implementation

To evaluate empirically the model as captured by equation **(21)**, we specify the following regression equation:

$$\tilde{r}_{it} = c + \gamma_i + \beta_1 \mathrm{E}_t \Delta q_{i,t+1} + \beta_2 var_t \left(\Delta C_{i,t+1} \right) + time_t \tag{22}$$

where $\tilde{r}_{i,t} = \log((1 + R_{it}) / (1 + R^{world}_t))$ and $\Delta q_{i,t+1} = \log(Q_{t+1}/Q_t)$. Variable $var_t \left(\Delta C_{i,t+1} \right)$ is the conditional variance of private consumption growth. We further assume that each country i is small so that $var_t \left(m^{w}_{i,t+1} \right) \approx var_t \left(m^{w}_{j,t+1} \right)$ for each country $i \neq j$. In this case, the term $var_t \left(m^{w}_{i,t+1} \right)$ in **(21)** will be captured by the common time-effect $time_t$. We also include a fixed effect, to allow for factors other than $var_t \left(\Delta C_{i,t+1} \right)$ to affect the risk premia. Theoretical predictions for the parameters are $\beta_1 > 0$ and $\beta_2 < 0$.

3.3 Data

We employ the real interest rates and the estimate of the world interest rate as described in section 2 and 3 to construct a measure of the interest rate differential ($\tilde{r}_{i,t}$). Here we focus on a sample of developed countries to guard against the possibility that the results are due to sovereign risk instead of foreign exchange rate risks. Also, the countries in our sample all score relatively high on an index on financial openness.[20] The data on the real effective exchange rate (q_{it}) is taken from OECD, and is a consumer-price based measure of the real effective exchange rate. A rise in q_{it} means that the exchange rate is depreciating. We use four measures to proxy for the volatility term $var_t \left(\Delta C_{i,t+1} \right)$, calculated as the midpoint of the 3-year and the 5-year rolling estimate of the standard deviation of quarterly real GDP and consumption growth, respectively. We include a number of controls: gdp and cpi are the mid-point of a three-year rolling average

[20] On the Index of Financial Openness (Ito and Chinn (2008)), all countries in our sample score over 1.4, with an average value of 2.0. On the index as a whole, the average value is 0, with scores varying between -1.8 and 2.5.

of quarterly GDP growth and cpi inflation, respectively. $volcpi$ is the mid-point of a three-year rolling average of the standard deviation of cpi inflation.

3.3.1 Unit root tests and the real exchange rate

To evaluate whether the real exchange rate data used here contain a unit root, we use a variety of unit root tests, including panel data tests.[21] We start by conducting four unit root tests on the individual time series, which all aim at overcoming the problems of size distortion and low power associated with unit root tests: The Dickey-Fuller (DF-GLS) test based on GLS detrending, the Phillips-Perron (PP) test, the Kwiatkowski, Philips, Schmidt and Shin (KPSS) test, and the Elliot, Rothenberg and Stock (ERS) test.[22] The results are reported in Table 6. The results for the 18 series are inconclusive; in none of the cases do all four tests point to the existence of a unit root, and in 14 of the 18 cases do at least 2 tests point to stationarity in the real exchange rate.[23] Since panel data unit root tests tend to have higher power than those based on pure time series models, we also conduct four panel data tests:[24] the Levin Lin and Chu (LLC) test, the Breitung (B) test, the Im, Pesaran and Shin (IPS) test, and Fisher-type tests (ADF and PP tests).[25] The results are reported in Table 7. The two tests that implement a common unit root across all countries do not reject the null hypothesis of a unit root. By contrast, the tests that allow for individual unit root processes all reject a unit root.

Based on these results, we argue that there is no strong evidence in favour of a unit root process in the real exchange rate data. Nevertheless, real exchange rates are characterised by very persistent movements, consistent with Rogoff's (1996) empirical result according to which PPP

[21] Following Baxter (1994), many studies have argued that because real exchange rates appear to contain nonstationary components, the data need to be filtered to remove the nonstationarities. However, as discussed by eg Taylor and Taylor (2004) and Rogoff (1996), the power of unit root tests is generally low, and it is therefore often unclear whether the real exchange rate contains a unit root. For this reason, panel data methods and long time samples provide more evidence in favour of a trend-reverting real exchange rate than do pure time series methods and short samples (Chinn, 2006).

[22] All series include a constant as a regressor.

[23] The DF, PP and KPSS test the null hypothesis that the series contain a unit root. The ERS test the unit root that the data is stationary.

[24] Although panel data unit root tests have higher power than those based on individual time series, they may also have problems. O'Connell (1998) shows that panel data unit root tests are problematic when there is cross-sectional correlation in the data. Also, as discussed by Taylor and Taylor (2004), the panel data tests applied test the null hypothesis that none of the real exchange rates contain a unit root. When the null is rejected, the most that can be inferred is that at least one of the rates is stationary. For this reason, we look at panel data tests in conjunction with tests applied on time series data.

[25] The LLC and B test assume that there is a common unit root process across countries, while the remaining tests allow for individual roots. All tests except for the B test include individual fixed effects. For the B test, we also include individual trends.

deviations tend to damp out, but only at a very slow rate.[26]

There are two main assumptions that one can rely on to explain highly persistent movements in the exchange rate: the Harrod-Balassa-Samuelson (HBS) effect and changes in the pattern of trade and specialisation that influences the terms of trade. The HBS effect predicts that countries experiencing trend real appreciation should have high productivity growth in the tradeable sector relative to the nontradeable sector.[27] The second explanation focuses on the composition of the tradeable basket. As countries become more developed, the range of goods in their exports shift towards higher-quality goods or goods requiring more sophisticated technology, with a subsequent rise in the price of their tradeables. Thus, the real exchange rate (and the terms of trade) of less developed countries should exhibit a trend appreciation. One reason for our lack of finding of a unit root in the real exchange rate may therefore be that our sample only contains advanced OECD countries, for which a trend appreciation of the real exchange rate is less likely. Below, we proceed with the estimation under the assumption that the data contain no unit root.

3.3.2 Cross-sectional dependence

Since the dependent variable contains a common component, one potential issue when estimating (22) is that the panel data may exhibit cross-sectional dependence.[28] We therefore implement the Pesaran (2004) test for cross-sectional dependence. The test is conducted after **(22)** has been estimated using the fixed effects model, where we replace $\mathrm{E}_t \Delta q_{it+1}$ with the realised change in the exchange rate:

$$\tilde{r}_{it} = c + \gamma_i + \beta_1 \Delta q_{i,t+1} + \beta_2 var_t \left(\Delta C_{i,t+1}\right) + time_t + u_{it} \tag{23}$$

where u_{it} is a residual. The null hypothesis tested is given by $H_0 : \rho_{ij} = \rho_{ji} = corr(u_{it}, u_{jt}) = 0$ for $i \neq j$. We obtain a test statistics equal to -5.474, implying that we cannot reject the null hypothesis of no cross-sectional dependence. The test

[26] Rogoff (1996) finds that empirical studies often suggests that shocks to the real exchange rate die out at the rate of around 15 percent per annum.

[27] The HBS effect assumes that PPP holds in the tradeable sector. In a world where labour is mobile between sectors but not between countries, a rise in productivity in the tradeable sector relative to that in the non-tradeable sector causes wages to rise in the non-tradeable sector as well. This leads to higher prices for non-tradeable goods, and a rise in the real exchange rate.

[28] As discussed by Hoyos and Sarafides (2006), the impact of cross-sectional dependence depends on the nature of the dependence. If the dependence is caused by a common factor which is uncorrelated with the included regressors, this can cause standard fixed-effect and random estimators to be consistent, but not efficient, and standard errors to be biased. When the common factor is correlated with the regressors, both the FE and the RE estimators will be biased and inconsistent.

results therefore suggest that standard panel data techniques would be appropriate for this dataset, despite the inclusion of a common variable in the dependent variable.

3.4 Estimation methods

3.4.1 GMM estimation

We estimate (22) using Generalized Methods of Moments (GMM). We replace the conditional expectations of the exchange rate change in (22) with actual data and introduce an expectation error, $\varepsilon_{i,t+1} = \tilde{r}_{it} - c - \gamma_i - \beta_1 \Delta q_{i,t+1} - \beta_2 var_t \left(\Delta C_{i,t+1}\right) - time_t$. Under rational expectations, $\varepsilon_{i,t+1}$ is uncorrelated with any information known at the decision date (period t).[29] Given this identifying assumption, any period t variable that is included in the decision-maker's information set can be used as an instrument to form the moment condition to estimate the model parameters. Under a more general representation, which allows for potential misspecification, identification requires some additional assumptions about the error term. To allow for higher-order processes for the error term ε_{it}, we here use an instrument set consisting of the third lags (compared to the model variables) of the interest rate differential and the change in the exchange rate, together with time and country dummy variables (denoted by D). The instrument set, Z_t, is given by: $Z_t = \left[\tilde{r}_{i,t-3}, \Delta q_{i,t-2}, D\right]$.[30]

Denoting the instrument set containing variables dated period t and earlier as Z_t and the parameter vector as $\theta = [\theta_1, \theta_2, ..., \theta_k]$, we can define the unconditional moment condition as

$$\mathrm{E}\left[\varepsilon_{t+1}(\theta)\, Z_t\right] = 0 \qquad \textbf{(24)}$$

To estimate the model, we use the iterative-GMM estimator and compute the Newey and West (1987) heteroscedasticity and autocorrelation consistent (HAC) estimator of the optimal weight matrix using four lags.

The instruments also need to be adequately correlated with the endogenous model variables.

[29] We use an *ex post* measure of the real interest rate to proxy for the *ex ante* rate. The two measures will differ when expected and realised inflation differs. In that case, the expectation error ε_{it} in **(22)** also contains the expectation error made on inflation. Under the assumption that financial market expectations are rational, the forecast error of inflation is not predictable, meaning that it will be uncorrelated with time t information.

[30] When choosing instrument set, we find that an instrument set consisting of the first and the second lag of the model variables do not perform well, to the extent that the overidentifying restrictions are rejected. For this reason, we include as instruments the third lag of these variables. The country and time dummy variables are used in the estimation of the fixed effects.

Otherwise they will provide limited ability to discriminate among different parameter values, and weak identification would arise, Ideally, the instrument set should be ’strong’ for the expected variable $\Delta q_{i,t+1}$. To assess instrument weakness, we compute the partial R^2 for the first-stage regression, which equals 0.07. The low R^2 suggests that the model can only be weakly identified. A potential weak-instrument problem could lead to imprecise estimates of the structural parameters, and the standard J statistics to draw inference may also be unreliable. To address this issue, we also compute an identification robust test statistics considered in recent literature - the Anderson and Rubin (AR) statistics. The main advantage of this statistic is that its limiting distribution is robust to weak and excluded instruments.

3.4.2 Switching model for the exchange rate

One issue with modelling expectations about exchange rate movements is the evidence of systematic bias in currency forecast, which appears at odds with investors’ rationality. Indeed, the covered interest parity condition, which relates the forward rate to the expected future value of the currency, is often rejected.[31] One explanation for this puzzle is the so called ’peso problem’. The idea is that even if agents are fully rational and learn instantly, they may be uncertain about a future shift in the exchange rate regime, driven by for example shifts in monetary policy. The peso problem occurs when individuals attach a small probability to a large change in economic fundamentals which has not occurred in the available sample (small sample problem).[32]

To see how the peso problem would affect the estimation results, suppose that there are two regimes in the economy: M_1 and M_2. If the economy is in regime 1 in period t, and agents assign a positive probability p_t to there being a shift in the regime in the following period, the expected exchange rate will follow

$$\mathrm{E}_t q_{t+1} = p_t \mathrm{E}_t \left(q_{t+1} | M_2\right) + \left(1 - p_t\right) \mathrm{E}_t \left(q_{t+1} | M_1\right) \qquad \textbf{(25)}$$

Under the assumption that the shift does not occur, the forecast error satisfies

$$q_{t+1}\left(M_1\right) - \mathrm{E}_t q_{t+1} = \left[q_{t+1}\left(M_1\right) - \mathrm{E}_t\left(q_{t+1}|M_1\right)\right] + p_t\left[\mathrm{E}_t\left(q_{t+1}|M_2\right) - \mathrm{E}_t\left(q_{t+1}|M_1\right)\right] \qquad \textbf{(26)}$$

[31] For an overview of this literature, see Sarno and Taylor (2002).

[32] Empirical work on the peso problem includes Engel and Hamilton (1990) and Kaminsky (1993). They try to explain the persistent movements in the dollar exchange rate during the 1980s using a switching regime model, where agents are uncertain about the future state of the world. Danthine and Donaldson (1999) show in a theoretical model that the expectation of a rare event affects the properties of an otherwise standard model so as to better fit the data.

We can write this as

$$q_{t+1}(M_1) - \mathrm{E}_t q_{t+1} = \eta_{t+1} + p_t \left[\mathrm{E}_t (q_{t+1}|M_2) - \mathrm{E}_t (q_{t+1}|M_1)\right] \quad \textbf{(27)}$$

where η_{t+1} is a disturbance term which, under the assumption of rational expectations is uncorrelated with any information available in period t. The second term represents the difference in the expected value of the future exchange rate in the two different regimes. Resulting from this term, the forecast error will be serially correlated with a non-zero mean, regardless of whether the regime switch occurs or not.

To allow for regime switches in the exchange rate, we estimate a Markov-Switching model for each country i :

$$q_t = c_{s_t} + \sum_{k=1}^{p} \rho_{k,s_t} q_{t-k} + \epsilon_t \quad \textbf{(28)}$$

where $\epsilon_t \sim iid\,(0)$ and s_t is a Markov-chain taking the value of 1 or 2 with transition matrix P. This model thus assumes that there are two regimes for the real exchange rate in country i, and the probabilities of switching between the two regimes are given by the probabilities summarised in the matrix P. We can estimate the model via maximum likelihood, using the filter suggested by Hamilton (1989). The forecast from this model is then given by

$$q_{t+1|t} = \left[\sum_{j=1}^{2} \left(c_j + \sum_{k=1}^{p} \rho_{k,j} q_{t+1-k}\right) \Pr(s_{t+1} = j|t)\right] \quad \textbf{(29)}$$

where $\Pr(s_{t+1} = j|t)$ refers to the probability that the state in period $t+1$ is given by j conditional on information available in period t. We can construct the expected change in the exchange rate, given **(29)**, to be used as a proxy for $\mathrm{E}_t \Delta q_{t+1}$ in the structural equation **(22)**. For comparison, we also estimate an AR(k) for the exchange rate for each country i, given by

$$q_t = c + \sum_{k=1}^{p} \rho_k q_{t-k} + \epsilon_t$$

The expected change in the exchange rate is constructed in the same way as for the Markov estimate.

4 Main results

To investigate the relationship between real interest rate differential, exchange rate depreciation and volatility of GDP growth, we start by splitting our sample into two periods: 1985-1996 and 1997-2007. For each period, and for each country, we calculate the average real interest rate differential, the average quarterly real exchange rate depreciation and the average standard deviation of GDP growth over these two periods. We then look at the relationship between the changes in these variables across the two time periods. As shown in Figure 3, there is a weak positive relation between the change in the interest rate differential and the average exchange rate depreciation (correlation coefficient equal to 0.19), although this relation is not significant. By contrast, countries that experienced an increase in the volatility of GDP growth across the two periods displayed declining real interest rate differentials (correlation coefficient equal to -0.53, significant at the 5 percent level (Figure 4)).

Although most countries experienced a fall in the volatility of GDP growth between the two sample periods, there is considerable variation in the extent of this reduction. Some countries have experienced a reduction in the volatility by over 50 percent, including the United Kingdom, New Zealand, Sweden and Spain. In these countries, the average interest rate differential (towards the world interest rate) rose by over 100 basis points between the first and the second period, which was substantially higher than the average of around 50 basis points. By contrast, volatility rose by over 50 percent between the two periods in two countries (Germany and Ireland), where the interest rate differential fell by over 200 basis points over the two subperiods. This suggests that long-run differences in changes in the economic volatility across countries have been associated with movements in the interest rate differential. Below, we explore the co-movements between the variables both over time and over countries while controlling for potential sources of spurious correlation, using the panel dataset.

4.1 GMM estimation

Column 1 in Table 8 show the results from estimating (22) using GMM on annual data, under the assumption that $\beta_2 = 0$, controlling for time and country fixed effects. In line with theory, the estimated coefficient on the exchange rate term is positive and significant. This is in contrast to earlier work that often obtains a negative estimate of the slope coefficient (the UIP puzzle).[33] Nevertheless, the estimated coefficient is significantly smaller than one, which is its predicted theoretical value. The p-value associated with the J statistics is 0.02, suggesting that we marginally reject the overidentifying restrictions. Since there is evidence that the instruments are

[33] Early contributions include Fama (1984) and Hansen and Hodrick (1980). More recent discussions include Lewis (1995) and Lustig and Verdelhan (2007).

weak, the J statistics may be unreliable. For this reason, we also look at the AR statistics, which is robust to weak instruments. The AR statistics reject the null that the estimated parameters lie in the instrument-robust confidence set.

Column (2) also estimates β_2, which is negative and significant. That is, countries where output is more volatile tend to have lower real interest rates, as predicted by theory. The regression equation is able to explain around 40% in the variation of interest rates across countries and time. The J-statistics now indicates that we cannot reject the overidentifying restrictions. However, the AR statistics still signals that the model may be misspecified. Column (3) includes some controls: GDP growth (gdp), inflation (cpi) and the volatility of inflation ($volcpi$). The rationale for adding these variables is to control for policy changes that could have an impact on both the level of interest rates, and macroeconomic volatility. We try to do this indirectly by including variables that would be affected by such shift in policy. The results are robust to the inclusion of these controls. Column 4-6 report results from the same regression, using the 5 year rolling average measure of GDP growth volatility, and the 3- and 5-year rolling average measure of consumption growth volatility. The results remain similar.

One drawback with our dataset is that it contains the Euro countries for which exchange rate risks have changed markedly since the introduction of the Euro. The model is, however, specified in terms of real exchange rates, which are not equalised across Euro countries. So although risks around the real exchange rate are likely to have fallen with the introduction of the Euro, they still remain. Instead of excluding the Euro area countries from our sample, we control for this issue by including a dummy variable that takes the value of one for countries that belong to the Euro area for the period 2000 and onwards. In all other cases, it takes the value of zero. Table 9 reports the results when we add the Euro area dummy ($euro$) in the regressions reported above. The results are very similar to those reported in Table 8. The dummy variable $euro$ enters negatively and significant in all regressions, while remaining coefficients are similar to those reported in Table 8. In particular, the estimated coefficient β_2 is negative and significant, in most of the cases. Hence, our results appear to be robust even when we try to control for the introduction of a common currency in the Euro area.

4.2 Regime switch model

We next estimate the exchange rate model for each country using a regime switching model, as described by equation **(28)**. For these estimations, we use quarterly data. Column 2-5 in Table 10 gives the sum of the estimated AR coefficients for the two stages and the estimated constant.[34] The model gives reasonable estimates of the AR coefficients for most countries and, in line with the unit root tests, the estimated sum of AR coefficients is smaller than one in both states in all countries except for Sweden, the United States and Germany, where it is greater than one in one of the states. The mean of the exchange rate process in state s is given by $\mu_s = c_s / \left(1 - \sum_{k=1}^{p} \rho_{k,s}\right)$, for $s = 1, 2$, where c_s and $\rho_{k,s}$ are the estimated constant and the AR coefficient of order k. Column 6-7 in Table 10 show the estimated mean associated with the two estimated states. Although the estimated means are similar, or identical, for most countries, there are large differences in a few cases: Canada, Sweden, the United Kingdom, the United States, Denmark and France. Hence, for these countries, allowing for a regime switch could have a substantial impact on the estimates of the change in the exchange rate.

Regression (1)-(3) in Table 11 show the estimates of equation **(22)** where we use the one-period-ahead forecasts from the Markov model to get estimates of $\mathrm{E}_t \Delta q_{i,t+1}$. We estimate the model using least squares, controlling for country and time fixed effects. The estimated coefficient on the exchange rate term is positive and significant, but significantly smaller than one. The estimated coefficient on the volatility term is negative and significant, and similar in magnitude to those reported in Table 8. The results are robust also when we include the control variables: gdp, cpi and $varcpi$. Regression (4)-(6) in the same table show the results based on the AR(k) model. Interestingly, the estimated coefficient on the exchange rate term is now negative, although insignificant. The negative sign is inconsistent with theory but, as discussed above, in line with previous empirical work on the UIP condition which often document that high interest-rate countries tend to find their exchange rate appreciating, rather than depreciating, and vice versa.

[34] Lag length is selected by the Akaike information criterion.

4.3 *Economic interpretation of the results*

To interpret the economic content of the volatility parameter, we compute the elasticity of the interest rate differential with respect to economic volatility. Using **(22)**, we get the following expression for the elasticity

$$\xi = \frac{\partial \log((1+R_{it})/(1+R_t^{world}))}{\partial \log var_t\left(\Delta C_{i,t+1}\right)} = \frac{\partial \tilde{r}_{it}}{\partial var_t\left(\Delta C_{i,t+1}\right)} var_t\left(\Delta C_i\right) = \beta_2 var_t\left(\Delta C_i\right)$$

Given the estimated value of β_2 reported in Table 8 (-0.005) and an average value of $var_t\left(\Delta C_i\right)$ of 0.72 (across countries and time), we get an estimate of the average elasticity parameter $\xi = -0.004$. This means that a percentage rise (fall) in volatility is expected to reduce (increase) the interest rate differential by 0.004 percent. To put this in context, Table 12 shows the average percentage changes in the interest rate differential $((1+R_{it})/(1+R_t^{world}))$ and economic volatility between the two periods 1985-1996 and 1997-2008. The countries that faced the largest falls in volatility over this period (United Kingdom, New Zealand, Sweden and Spain, where output volatility fell by 61 percent on average)) saw their interest rate differential increase by 1 percent on average. Given our estimate of the elasticity parameter, the model predicts that around one quarter of that rise was due to falling volatility. Hence, although the elasticity estimate is small, large variations in economic volatility over the past means that the impact on real interest rate differentials could have been significant.

We can also interpret the results in terms of the level of interest rates, returning to the question of what affects differences in the level of interest rates across countries. We do so by testing the null hypothesis that the intercept in (22) is zero: $c + \gamma_i = 0$. That is, we test if there are persistent deviations in the level of real interest rates once we control for the expected exchange rate change and the variability of output. Table 13 shows the χ^2 test statistics and the significance levels associated with the test, based on the estimation results reported in column 2 in Table 8.[35] We reject the null hypothesis for 6 countries: Japan, United States and Switzerland (with positive values of $c + \gamma_i$), and New Zealand, Norway and Australia (with negative values of $c + \gamma_i$).

We finally note that, over the sample period, three currencies have dominated the reserve currency holdings across the world: The US dollar, the euro, and the German mark (prior to the

[35] We base this evaluation on the estimation results obtained on annual data. The reason for doing so is to abstract from relative short-term fluctuations in real interest rates and exchange rates.

formation of the Euro Area).[36] It is often argued that, because of the role of these currencies as reserve currencies, some countries are able to issue bonds at a lower rate than other countries due to lower liquidity premia. To analyse this, we include a dummy variable ($reserve$) that takes the value one if the currency is a reserve currency, and zero otherwise. As countries with reserve currencies, we here include the US, Germany and, for the period after 1999, countries that belong to the Euro Area (Belgium, France, Italy, Netherlands, Ireland, Portugal, and Spain). The regression results are reported in the last column in Table 8. The estimated coefficient on the reserve dummy is negative and significant, in line with theory. The estimates of the remaining coefficients are similar to those in the regression without the reserve dummy (column 1). Once we control for whether a currency is one of the major reserve currencies or not, the estimated country-specific constant $c + \gamma_i$ is significant for only three countries: Japan, New Zealand and Switzerland (column 3 in Table 13).

5 Exploring the link between interest rate differentials and the net foreign asset position

One of the weaknesses of the above framework is that it does not model the portfolio choices made by individuals. It is therefore unable to establish whether the net foreign asset (NFA) position of a country is affected by, or affects, the real interest rate differential.

Previous empirical work by Lane and Milesi-Ferretti (2001) and Selaive and Tuesta (2003) find evidence of a significant negative relation between net foreign asset position and the interest rate differential relative the rest of the world. It is not clear, however, what drives this negative relation.[37]

Recent work has analysed the relation between the build-up of net foreign assets and business cycle volatility through the motive for precautionary savings. In an analysis of Asian countries, Durdu et al (2008) show that precautionary acquisitions of foreign assets are partly driven by higher business cycle volatility. That is, as volatility increases, risk averse individuals want to

[36] On average over the period 1995-2007, 66% of foreign exchange reserves consisted of US dollars. Over 1995-1999, 15% of reserves were made up of the Mark and, since 1999, 23% of reserves have been made up of Euro.

[37] The existence of a 'portfolio' balance effect can be interpreted as reflecting a home bias in asset markets, and/or upward-sloping supply of international funds. Theoretical models also imply a negative relation between the interest rate differential and the NFA position (Benigno (2001)). These, however, rely on an ad hoc formulation which assumes rather than explaining that it is costly to undertake positions in the international asset market for households in the home country, and that this cost depends on the NFA position of the domestic economy.

increase precautionary savings, to guard against bad shocks. Similar results are obtained by Fogli and Perri (2006).[38] These models thus imply that if the volatility of shocks varies between countries and over time, the precautionary savings motive also changes. In an open-economy setting, this generates external imbalances, with the more volatile countries building up a positive foreign position compared to less volatile ones. In an empirical study, Fogli and Perri (2008) also find a significant positive relation between the net foreign asset position and the volatility of GDP growth, consistent with theory. These studies, however, assume that there is a single internationally traded bond, and therefore do not model interest rate differentials across countries.

In this section, we explore the empirical relation between the NFA position and the real interest rate differential. Again, we start by splitting the sample into two periods: 1985-1996 and 1997-2007. For each period, and for each country i, we calculate the average NFA position as a ratio to GDP.[39] We look at the relationship between the change in this variable and the change in the interest rate differential between the two time periods. As shown in Figure 5, there is a weak negative relation between the two variables (correlation coefficient equal to -0.32), although the relation is insignificant.

Some countries have seen a substantial worsening of their position towards the rest of the world. For example, the external position of the United States fell from -3.5% of GDP in the first period to over -22% in the second. Similar deteriorations occurred in the United Kingdom, the US, Spain, Portugal and the Netherlands. These countries also saw an increase in the interest rate differential towards the world interest rate by on average over 100 basis points, compared to around 50 basis points on average across all sample countries. Interestingly, though, countries which saw a large build-up of net foreign assets (Belgium, Norway, Sweden, Switzerland and Ireland) did not face a substantial fall in real interest rates relative to the rest of the world. Instead, the average differential in these countries rose by around 30 basis points.

5.1 An empirical model

To evaluate the relationship between the net foreign asset position and the real interest rate differential, we postulate two relations: one that links the relative volatility of consumption

[38]Similar results are obtained in a recent paper by Bems and Carvalho Filho (2009), and are consistent with the theoretical predictions in Carroll and Jeanne (2009).

[39]The construction of the data is discussed in section 5.2.

growth to the foreign exchange rate premium, and one that links economic volatility to the acquisition of net foreign assets:

$$\nu_t = \chi \left[var_t \left(\Delta C_{t+1}\right) - var_t \left(\Delta C^*_{t+1}\right)\right] \tag{30}$$

$$nfa_t = \phi \left[var_t \left(\Delta C_{t+1}\right) - var_t \left(\Delta C^*_{t+1}\right)\right] + \xi Z_t \tag{31}$$

where ν_t is the foreign exchange risk premium of the home country, nfa_t is the NFA position of the home country, and Z_t is a vector of variables, other than economic volatility, that affect the net foreign asset position. The results in Durdu et al (2008) and Fogli and Perri (2006) are consistent with **(31)** under the assumption that $\phi > 0$ and, as discussed in section (3.1), asset-pricing theories imply that $\chi < 0$ (under the assumption of no habits in consumption).

We can write the UIP condition as

$$r_t = r^*_t + \beta_1 \mathrm{E}_t \Delta q_{i,t+1} + \beta_2 \left[var_t \left(\Delta C_{t+1}\right) - var_t \left(\Delta C^*_{t+1}\right)\right] \tag{32}$$

Combining **(31)-(32)** gives the following reduced-form relation between the interest rate differential and the NFA position:

$$\tilde{r}_{it} = \delta_1 \mathrm{E}_t \Delta q_{i,t+1} + \delta_2 nfa_{it} + \Gamma Z_{it} + \varepsilon_{it} \tag{33}$$

where parameters δ_1 and δ_2 fulfil:

$$\delta_1 = \beta_1, \qquad \delta_2 = \frac{\beta_2}{\phi} \tag{34}$$

We next want to test if the reduced-form relationship between the interest rate differential and the NFA position in **(33)** indeed seem to capture the structural relationships between the NFA position and economic volatility, on the one hand, and between economic volatility and the real interest rate differential, on the other hand, or whether the NFA position has a separate impact on the interest rate differential. To do so, we specify the following system of equations:

$$nfa_{it} = \varphi_i + \varphi_1 var_t \left(\Delta C_{i,t+1}\right) + \Phi_2 Z_{it} + \upsilon^1_{it} \tag{35}$$

$$\tilde{r}_{it} = \delta_i + \delta_1 \mathrm{E}_t \Delta q_{i,t+1} + \delta_2 nfa_{it} + \Gamma Z_{it} + \upsilon^2_{it} \tag{36}$$

$$\tilde{r}_{it} = \beta_i + \beta_1 \mathrm{E}_t \Delta q_{i,t+1} + \beta_2 var_t \left(\Delta C_{i,t+1}\right) + time_t + \upsilon^3_{it} \tag{37}$$

where the constants capture country fixed effects. The three equations constitute a system of simultaneous equations, where variables Z_{it}, $var_t \left(\Delta C_{i,t+1}\right)$ and $\mathrm{E}_t \Delta q_{i,t+1}$ are taken to be exogenous, and nfa_{it} and $\tilde{r}_{it}$ are endogenous. To obtain values for $\mathrm{E}_t \Delta q_{i,t+1}$, we construct estimates based on the Markov switching model discussed in section (4.2).[40] We estimate equation **(35)-(37)** simultaneous using least squares where, prior to estimation, we impose the constraint that $\delta_1 = \beta_1$.

When estimating **(35)-(37)**, we need to control for variables other than volatility that are likely to affect the net foreign asset position of a country. As stressed by Lane (2002), significant movements in the NFA position are likely to reflect differences across countries in savings and investment behaviour rather than business cycle shocks. Lane and Milesi-Ferretti (2001) identify a number of fundamental variables that are likely to affect these differences: A relative rise in output per capita is likely to affect the NFA position positively through its impact on the marginal product of capital, and on the saving rate. When Ricardian equivalence does not hold, a high level of public debt may give rise to a decline in the external position, if the increase in public debt is not fully offset by a rise in private savings. Demographic factors are likely to affect the NFA position. A country with an ageing population is likely to save more as they foresee a rising share of retirees, leading to an improvement in the NFA position. By contrast, when the share of young is high relative to the share of working population, the savings rate is likely to be lower.

5.2 Data

We take data on the net foreign asset position from Lane and Milesi-Ferretti (2006) for the period 1985-2004, and update it using current account and capital transfers data (taken from IMF Balance of Payment Yearbook) for 2005-2007. The net foreign asset position is expressed as a ratio to GDP (nfa). We control for the following variables: The level of real GDP per capita ($GDPcap$), public debt as a percentage of GDP ($debt$), the share of population of age 15-64 ($working$) and the share of population above 54 (old). The demographics data is available as five-year averages. Data on public debt is taken from OECD (central government debt), and population data is taken from the United Nations, World Population Prospects.

[40]Since the net foreign asset position is only available at the annual frequency, we here estimate the Markov regime switching model on annual, instead of quarterly, data.

5.3 Empirical results

Column 2 in Table 14 reports the results from estimating equation **(35)**. In line with previous results, there is a significant positive relationship between the measure of economic volatility and the net foreign asset position. One interpretation of this result is that, when economic volatility is high, the incentive to build up precautionary savings is higher, resulting in an increase in the NFA position. We also find that a rise in GDP per capita improves the net foreign asset position, while a rise in the share of old in the population tends to decrease it. These results are both in line with theory. By contrast, a rise in debt increases the NFA position. A rise in the proportion of old people (relative to the share of young) has a negative impact on the NFA position, while a rise in the share of working age population has no significant impact on it.[41]

Column 2 report the results when we estimate equation **(36)**, capturing the reduced-form relation between the net foreign asset position and the interest rate differential. The estimated coefficient on nfa is negative and significant, implying that an improvement in the net foreign asset position reduces the interest rate differential.

Column 3, finally, estimates **(35)**-**(37)** jointly. The signs of the main variables of interest are in line with theory, and the estimated coefficients are significant. We next test whether the estimated coefficient δ_2 in **(36)** is equal to β_2/φ_1, as predicted by the model. The estimates of these coefficients are significant, with $\hat{\delta}_2 = -0.0008$, $\hat{\beta}_2 = -0.006$, and $\hat{\varphi}_1 = 7.621$. The ratio $\hat{\beta}_2/\hat{\varphi}_1$ is thus equal to -0.0008. We next test the null hypothesis: $H_0 : \delta_2 = \beta_2/\varphi_1$ using a χ^2 test. The $\chi^2(1)$ test statistics for testing the null hypothesis is close to zero (0.00), implying that we cannot reject the null hypothesis that the relation between the net foreign asset position and the interest rate differential in **(36)** is a reduced-form relation, capturing the link between economic volatility and the NFA position, on the one hand, and the link between economic volatility and the interest rate differential, on the other hand.

[41] The contemporaneous correlations between the NFA position and the explanatory variables are: (i) negative but insignificant for the level of debt, (ii) positive and significant for gdp per capital, (iii) negative and significant for the proportion of young, (iv) positive and significant for the proportion of working age, (v) positive and significant for the proportion of old. These are all in line with theory. However, once we control for country fixed effects and the volatility of output growth, some of these correlations become insignificant, or switch sign.

6 Conclusion

We construct a measure of the world interest rate using principal component analysis. The world interest rate is taken as a starting point to analyse differences in the level of short-term real interest rates across a sample of 18 developed countries. We establish the fact that in many countries, real interest rates have deviated from the world interest rate for long periods of time. Deviations from real interest rate parity could be due to the fact that the price of nontraded goods relative to traded goods move differently across countries, or that the share of nontraded goods in consumption expenditure is large. We control for this by constructing a measure of the real interest rate that is less affected by the inclusion of non-tradeables. We find that the inclusion of non-tradeables is likely to account for some of the differences in real interest rates across countries, but a large part remains unexplained.

These unexplained interest rate differentials are likely to reflect risk premia. Our focus on the shorter end of the yield curve for a sample of developed countries with good credit ratings during our sample period allows us to focus on the foreign exchange rate risk premium. We use a standard asset-pricing framework to derive a structural equation for this premium. This relation states that it is negatively related to the volatility of consumption growth in the home country relative to abroad. One prediction from this model, that has not been tested empirically, is therefore that differences in real interest rates across countries should be negatively related to a measure of relative volatility in consumption growth. Taking this hypothesis to the data is the main contribution of the paper.

Using panel data techniques, we estimate an empirical UIP condition relating the interest rate differential relative to the world interest rate to the expected change in the real exchange rate, and a measure of relative consumption volatility. Our main result is that, consistent with theory, there is a significant negative relation between real interest rate differentials and the volatility of both output and private consumption growth. This result is also robust to different methods for proxying expectations of future exchange rate changes. The empirical results imply that a percentage rise in the volatility of output growth reduces the real short-term interest rate relative to the world interest rate by 0.004 percent. Although the elasticity parameter is small, large movements in economic volatility over the past means that the impact on real interest rates could have been significant.

We finally explore the empirical relationship between real interest rates and the net foreign asset position. To analyse this, we draw on recent theoretical work by Durdu et al (2008) and Fogli and Perri (2006) that show that the equilibrium external balance of a country is positively affected by the strength of its precautionary savings motive relative to that of its trading partners. The precautionary savings motive, in turn, is positively related to economic volatility. Together with asset-pricing theories that predict a negative relation between the real interest rate differential and economic volatility, this suggests a negative relation between real interest rates and the net foreign asset position.

In this paper, we test the hypothesis that the negative empirical relation between real interest rates and the net foreign asset position is a reduced-form relation, capturing the links described above. We are not able to reject this hypothesis.

Appendix A

Assuming that the M_t and Q_t are jointly log-normal, **(10)** and **(12)** imply the following relations

$$r_t = -\mathrm{E}_t m_{t+1} - \frac{1}{2} var_t\left(m_{t+1}\right)$$

$$r_t^* = -\mathrm{E}_t m_{t+1} - E_t \Delta q_{t+1} - \frac{1}{2} var_t\left(m_{t+1}\right) - \frac{1}{2} var_t\left(q_{t+1}\right) - cov_t\left(m_{t+1}, q_{t+1}\right)$$

Together, this implies that

$$r_t - r_t^* = E_t \Delta q_{t+1} + \upsilon_t \qquad \textbf{(A-1)}$$

where

$$\upsilon_t = \frac{1}{2} var_t\left(q_{t+1}\right) + cov_t\left(m_{t+1}, q_{t+1}\right) \qquad \textbf{(A-2)}$$

We can continue by solving for the volatility and covariance of the exchange rate in terms of the domestic and foreign stochastic discount factors. To do so, we use **(14)** to get

$$\Delta q_{t+1} = m_{t+1}^* - m_{t+1} \qquad \textbf{(A-3)}$$

Combining **(A-2)** and **(A-3)** gives

$$\begin{aligned} \upsilon_t &= \frac{1}{2} var_t\left(m_{t+1}^* - m_{t+1}\right) + cov_t\left(m_{t+1}, m_{t+1}^* - m_{t+1}\right) \\ &= \frac{1}{2} var_t\left(m_{t+1}^*\right) + \frac{1}{2} var_t\left(m_{t+1}\right) - cov_t\left(m_{t+1}^*, m_{t+1}\right) - var_t\left(m_{t+1}\right) + cov_t\left(m_{t+1}, m_{t+1}^*\right) \end{aligned}$$

Simplifying gives

$$\upsilon_t = \frac{1}{2} var_t\left(m_{t+1}^*\right) - \frac{1}{2} var_t\left(m_{t+1}\right)$$

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Table 1: Sample statistics for real interest rates (sample period 1985Q2-1997Q4)

Country	Mean	Median	st dev	25 prct	75 prct	iq range	min	max	mm range
Austria	4.2	3.6	2.4	2.8	5.2	2.4	−1.1	11.2	12.4
Belgium	3.1	2.5	2.4	1.3	5.2	4.0	−1.6	8.2	9.8
Canada	3.2	3.1	2.3	1.5	4.9	3.4	−1.2	9.1	10.4
Denmark	4.6	2.9	3.9	1.4	8.0	6.6	−0.4	13.0	13.4
France	3.5	2.9	3.6	1.8	5.2	3.4	−0.7	9.1	9.8
Germany	2.4	2.2	3.7	1.4	3.0	1.5	−0.6	7.8	8.4
Ireland	3.2	3.8	3.7	−0.6	5.8	6.5	−2.4	15.1	17.5
Italy	3.7	4.1	2.9	1.1	5.7	4.5	−0.7	13.3	14.0
Japan	1.4	0.9	1.6	0.3	2.8	2.5	−2.0	5.5	7.5
Netherlands	2.8	2.2	2.3	1.0	5.0	4.0	−0.9	7.2	8.1
NZ	5.3	4.9	2.2	3.7	6.5	2.8	−0.5	13.0	13.5
Norway	4.0	3.7	2.8	2.3	6.3	4.0	−6.3	13.7	20.0
Portugal	2.0	1.5	2.2	0.0	3.6	3.6	−1.3	6.4	7.7
Spain	2.6	2.2	2.9	0.4	4.9	4.5	−1.9	9.3	11.2
Sweden	3.7	3.2	2.5	1.6	5.3	3.7	−0.1	15.6	15.8
Switzerland	1.4	1.3	1.3	0.5	2.1	1.6	−1.1	4.9	6.0
UK	4.3	3.8	1.8	2.9	5.3	2.4	1.0	9.1	8.1
US	1.5	1.9	1.7	0.3	2.7	2.5	−3.1	4.6	7.7
World	3.3	2.9	2.2	1.4	5.4	4.0	−0.2	8.3	8.4

Table 2: Principal components

	Non-standardised data			Standardised data		
Country	1st	2nd	3rd	1st	2nd	3rd
Austria	0.18	0.20	0.45	0.21	0.35	0.01
Belgium	0.27	0.14	−0.07	0.28	−0.08	0.12
Canada	0.22	0.08	0.28	0.26	0.20	−0.05
Denmark	0.41	0.24	−0.26	0.26	−0.14	0.12
France	0.27	0.02	−0.12	0.28	−0.18	0.06
Germany	0.14	0.19	0.02	0.24	0.06	0.28
Ireland	0.41	−0.18	−0.04	0.27	−0.15	−0.07
Italy	0.32	−0.29	−0.02	0.28	−0.18	−0.19
Japan	0.15	0.27	−0.06	0.24	0.01	0.35
Netherlands	0.24	0.23	−0.05	0.27	−0.05	0.22
NZ	0.13	−0.01	0.49	0.18	0.35	−0.23
Norway	0.23	0.17	−0.40	0.21	−0.24	0.24
Portugal	0.14	−0.54	−0.01	0.16	−0.14	-0.56
Spain	0.28	−0.14	0.02	0.25	−0.14	−0.07
Sweden	0.21	−0.47	0.02	0.21	−0.17	−0.39
Switzerland	0.11	0.03	0.03	0.23	0.02	0.04
UK	0.13	0.22	0.31	0.20	0.46	0.07
US	0.10	−0.03	0.36	0.17	0.45	−0.29
Variance	0.68	0.06	0.06	0.61	0.08	0.08
Tot variance	0.68	0.74	0.80	0.61	0.69	0.77

Notes: The table reports the values of the principal components' coefficients. Variance is the variance explained by each PC.
Tot variance is the the cumulative variance explained by the PCs.

Table 3: Correlations with world interest rate

Country	Full sample	pre-1997	post-1997
Austria	0.7	0.5	0.5
Belgium	0.9	0.9	0.8
Canada	0.8	0.5	0.8
Denmark	0.9	0.8	0.5
France	0.9	0.8	0.7
Germany	0.7	0.6	0.7
Ireland	0.9	0.8	0.9
Italy	0.9	0.6	0.9
Japan	0.8	0.7	−0.3
Netherlands	0.9	0.9	0.4
NZ	0.6	0.1	0.8
Norway	0.7	0.7	0.2
Portugal	0.6	−0.5	0.8
Spain	0.8	0.4	0.9
Sweden	0.7	0.2	0.8
Switzerland	0.7	0.5	0.7
UK	0.6	0.6	0.8
US	0.5	0.0	0.9
Average	0.8	0.5	0.7

Notes: The pre-1997 sample covers 1985Q3-1996:Q4. The post-1997 sample covers 1997Q1-2008:Q2.

Table 4: Testing PPP

Country	α_i	α_i
Austria	0.79**	0.86**
Belgium	−0.22*	0.15
Canada	−0.09	−0.03
Denmark	1.20**	1.24**
France	0.17*	−0.53**
Germany	−0.89**	−0.88**
Ireland	−0.16	1.65**
Italy	0.41**	0.19
Japan	−1.90**	−1.56**
Netherlands	−0.49**	−1.03**
NZ	1.91**	1.78**
Norway	0.75**	0.24
Portugal	−1.29**	
Spain	−0.69**	0.21
Sweden	0.34*	0.41
Switzerland	−1.86**	−1.31**
UK	0.99**	0.52*
US	−1.83**	−1.91**

Notes: Column 2 and 3 shows estimate of the constant in (8) using CPI and WPI, respectively.
*(**) denotes significant at the 10(5) percent level.

Table 5: Portmanteau Q statistics for ε_{it}^{a}

	Model 1			Model 2		
Country	Lag 1	Lag 2	Lag 3	Lag 1	Lag 2	Lag 3
Austria	1.5**	1.9**	2.7**	0.4**	0.7**	3.8**
Belgium	0.5**	0.7**	2.1**	0.9**	0.9**	0.7**
Canada	0.9**	4.7*	14.9	7.9	8.3	11.2
Denmark	11.2	19.6	24.6	10.7	17.3	18.7
France	5.7	6.7	8.6	3.0*	3.4**	6.1**
Germany	6.9	8.6**	10.9*	2.5**	2.7**	4.3**
Ireland	11.9	15.4	16.3	1.6**	2.2**	3.0**
Italy	9.6	11.9	14.3	0.2**	1.4**	2.3**
Japan	10.2	15.8	20.3	3.9	4.8*	9.4
Netherlands	4.1	7.6	8.6	2.9*	3.7**	7.2*
NZ	5.2	5.4*	5.7**	5.4	5.4*	8.6
Norway	2.2**	2.6**	2.6**	4.3	4.7*	5.8**
Portugal	11.6	14.6	15.0			
Spain	6.7**	7.5	7.5*	12.4	16.1	17.6
Sweden	1.2**	1.6**	3.4**	0.4**	1.1**	9.7
Switzerland	10.3	20.6	29.0	11.3	14.3	14.9
UK	13.9	18.5	20.4	11.2	13.3	13.6
US	16.4	24.4	27.5	4.5	5.0*	5.0**

Notes: Q statistics for Portmanteau test for white noise of ε_{it}^{a} in (9).
*(**) denote p-value greater than 0.10 and 0.05 respectively.

Table 6: Unit root tests for the real exchange rate

Country	DF-GLS	PP	KPSS	ERS
Austria	*A*	*A*	*R*	*R*
Belgium	*A*	*A*	*R*	*R*
Canada	*A*	*A*	*R*	*R*
Denmark	*A*	*R*	*A*	*R*
France	*A*	*A*	*A*	*R*
Germany	*R*	*A*	*R*	*R*
Ireland	*A*	*A*	*R*	*R*
Italy	*R*	*A*	*R*	*A*
Japan	*A*	*A*	*R*	*R*
Netherlands	*A*	*A*	*R*	*R*
NZ	*A*	*A*	*R*	*R*
Norway	*R*	*A*	*R*	*A*
Portugal	*A*	*A*	*A*	*R*
Spain	*A*	*A*	*R*	*A*
Sweden	*A*	*A*	*A*	*R*
Switzerland	*A*	*R*	*R*	*R*
UK	*A*	*A*	*A*	*R*
US	*A*	*R*	*R*	*R*

Notes: A and R denote evidence of unit root, or no unit root, respectively. For DF-GLS, PP and KPSS tests, R implies that null of unit root is rejected at the 5 percent level.
For ERS test, A implies that null of no unit root is rejected at the 5 percent level.

Table 7: Panel unit root tests for real exchange rate

test	Test statistics	Prob
	Common unit root process	
LLC	−1.04	0.15
B	−0.84	0.20
	Individual unit root processes	
IPS	−2.81	0.00
Fisher - ADF	59.26	0.01
Fisher - PP	58.83	0.00

Table 8: Dependent variable: Interest rate differential

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δs	0.105*	0.120*	0.103	0.121*	0.129*	0.115*	0.092
	[0.073]	[0.070]	[0.880]	[0.071]	[0.072]	[0.100]	[0.075]
var		−0.005***	−0.004**	−0.005***	−0.004***	−0.005***	−0.003**
		[0.002]	[0.002]	[0.002]	[0.001]	[0.002]	[0.001]
gdp			0.001				
			[0.002]				
cpi			−0.001				
			[0.001]				
$varcpi$			−0.002				
			[0.001]				
$reserve$						−0.006**	
						[0.001]	
R^2	0.39	0.41	0.45	0.40	0.39	0.41	0.46
J-test	5.17	2.81	2.37	3.55	4.23	4.76	1.06
	[0.02]	[0.09]	[0.12]	[0.06]	[0.04]	[0.03]	[0.30]
AR-test	23.06	23.16	16.7	23.18	23.0	23.06	19.75
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Obs	342	342	342	342	342	342	342

Notes: Estimation of (20) using GMM with fixed year and time effects. Measure of var in (1)-(3) and (7) based on 3-year rolling average of standard deviation of output growth, (4)-(6) use alternative measures. *(**)(***) significant at 10-, 5- and 1-percent level. Standard errors in square brackets.
For test statistics, p-value in square bracket.

Table 9: Dependent variable: Interest rate differential controlling for Euro Area

	(1)	(2)	(3)	(4)	(5)	(6)
Δs	0.075	0.096	0.079	0.093*	0.099	0.089
	[0.067]	[0.068]	[0.084]	[0.069]	[0.069]	[0.068]
var		−0.004**	−0.003	−0.003*	−0.003**	−0.004**
		[0.002]	[0.002]	[0.002]	[0.001]	[0.002]
gdp			0.001			
			[0.002]			
cpi			−0.001			
			[0.001]			
varcpi			−0.002			
			[0.001]			
euro	−0.005***	−0.004***	−0.004**	−0.005***	−0.005***	−0.005***
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
R^2	0.43	0.44	0.48	0.44	0.44	0.44
J-test	3.06	1.84	1.46	2.39	2.62	2.96
	[0.08]	[0.17]	[0.23]	[0.12]	[0.11]	[0.09]
Obs	342	342	342	342	342	342

Notes: Estimation of (20) using GMM with fixed year and time effects. Measure of *var* in (1)-(3) and (7) based on 3-year rolling average of standard deviation of output growth, (4)-(6) use alternative measures. *(**)(***) significant at 10-, 5- and 1-percent level. Standard errors in square brackets.
For test statistics, p-value in square bracket.

Table 10: Estimated coefficients Markov Switching model

	Sum of AR coefficients		Constant		Mean	
	State 1	State 2	State 1	State 2	State 1	State 2
Australia	0.95	0.95	0.01	0.01	0.23	0.23
Belgium	0.74	0.95	0.01	0.01	0.35	0.59
Canada	0.90	0.93	0.31	−0.10	3.17	−1.44
Denmark	0.81	0.90	−0.05	0.16	−0.24	1.57
France	0.90	0.96	−0.02	0.04	−0.16	0.99
Germany	0.93	0.93	0.01	0.01	0.10	0.10
Ireland	0.98	0.98	0.02	0.02	0.92	0.92
Italy	0.93	0.93	0.01	0.01	0.18	0.18
Japan	0.91	0.91	0.01	0.01	0.07	0.07
Netherlands	0.90	0.90	−0.01	−0.01	−0.13	−0.13
NZ	0.91	0.91	0.02	0.02	0.22	0.22
Norway	0.88	0.88	0.01	0.01	0.12	0.12
Portugal	0.98	0.98	0.02	0.02	1.39	1.39
Spain	0.94	6.44	0.02	0.45	0.34	−0.08
Sweden	1.01	0.83	0.02	−0.08	−2.04	−0.47
Switzerland	0.86	0.86	0.03	0.03	0.22	0.22
UK	0.40	0.91	−0.78	0.07	−1.31	0.75
US	1.08	0.84	0.02	0.02	1.78	0.49

Table 11: Dependent variable: Interest rate differential

	Markov			AR		
	(1)	(2)	(3)	(4)	(5)	(6)
β_1	0.180***	0.196***	0.163***	−0.001	−0.007	−0.009
	[0.039]	[0.039]	[0.038]	[0.008]	[0.008]	[0.008]
β_2		−0.004***	−0.003***		−0.004***	−0.003***
		[0.001]	[0.000]		[0.000]	[0.000]
gdp			−0.001*			−0.0003**
			[0.000]			[0.000]
cpi			−0.001***			−0.001***
			[0.000]			[0.000]
$varcpi$			−0.001***			−0.001*
			[0.000]			[0.000]
R^2	0.33	0.37	0.39	0.32	0.37	0.38
Obs	1656	1656	1656	1656	1656	1656

Notes: LS estimation of (20). Standard errors in parenthesis: *(**)(***) significant at 10-, 5- and 1-percent level. p-values in square brackets.

Table 12: Change in interest rate differential and volatility between 1985-1996 and 1997-2008 (percent)

Country	Interest rate	Volatility
Austria	1.07	−29.03
Belgium	−0.02	24.35
Canada	0.39	−38.63
Denmark	−2.23	65.06
France	0.02	−13.55
Germany	1.89	−46.10
Ireland	−2.20	89.05
Italy	−0.94	−14.69
Japan	1.31	−18.10
Netherlands	0.39	−45.54
NZ	1.72	−63.17
Norway	0.86	−22.10
Portugal	0.88	14.09
Spain	−0.51	73.62
Sweden	0.50	−53.15
Switzerland	2.28	−24.43
UK	2.34	−54.25
US	2.45	−11.34

Table 13: Test statistics, $H_0 : c + \gamma_i = 0$

Country	*var*	*var and reserve*
Austria	2.91*	0.96
Belgium	0.06	0.00
Canada	0.17	0.08
Denmark	2.17	0.44
France	0.70	0.55
Germany	0.57	0.67
Ireland	0.44	0.28
Italy	0.53	1.19
Japan	5.62**	10.38***
Netherlands	0.56	0.36
NZ	15.13***	9.12***
Norway	4.53**	1.29
Portugal	1.08	0.97
Spain	0.02	0.01
Sweden	0.60	0.01
Switzerland	4.45**	8.82***
UK	1.89	0.99
US	2.71*	0.55

Notes: $\chi^2(1)$ statistics. *(**)(***) significant at 10-, 5- and 1-percent level.

Table 14: System of equations

	(1)	(2)	(3)
Dependent variable: Net foreign asset position			
var	6.176* [3.252]		7.621***
gdpcap	0.003***[0.000]		0.003***
debt	13.318*		15.430**
old	−3.229***		−3.282***
working	−1.369[1.430]		−1.526*
Dependent variable: Interest rate differential			
nfa		−0.00004*	−0.0008***
ds		−0.020	0.038**
gdpcap		0.000***	0.000***
debt		0.011***	0.013**
old		−0.0004[0.0004]	−0.003***
working		−0.001**[0.0003]	−0.001*
Dependent variable: Interest rate differential			
var			−0.006***
ds			0.038**

Notes:

Figure 1: PC measures of the world interest rate

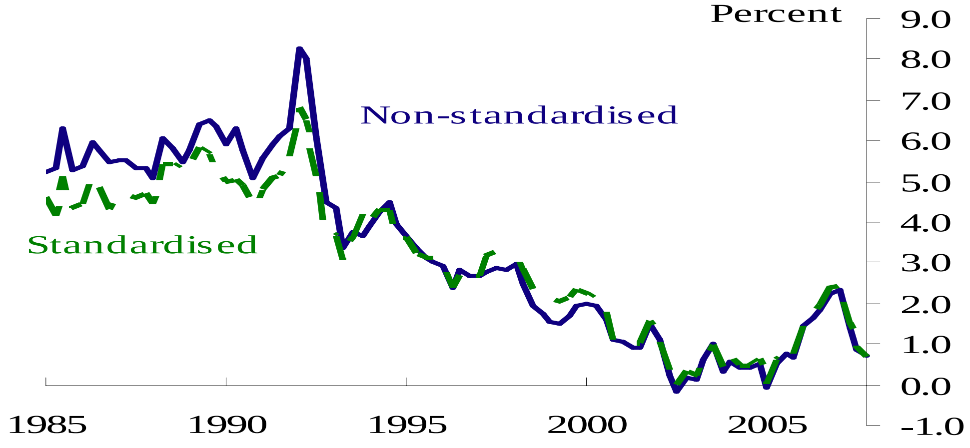
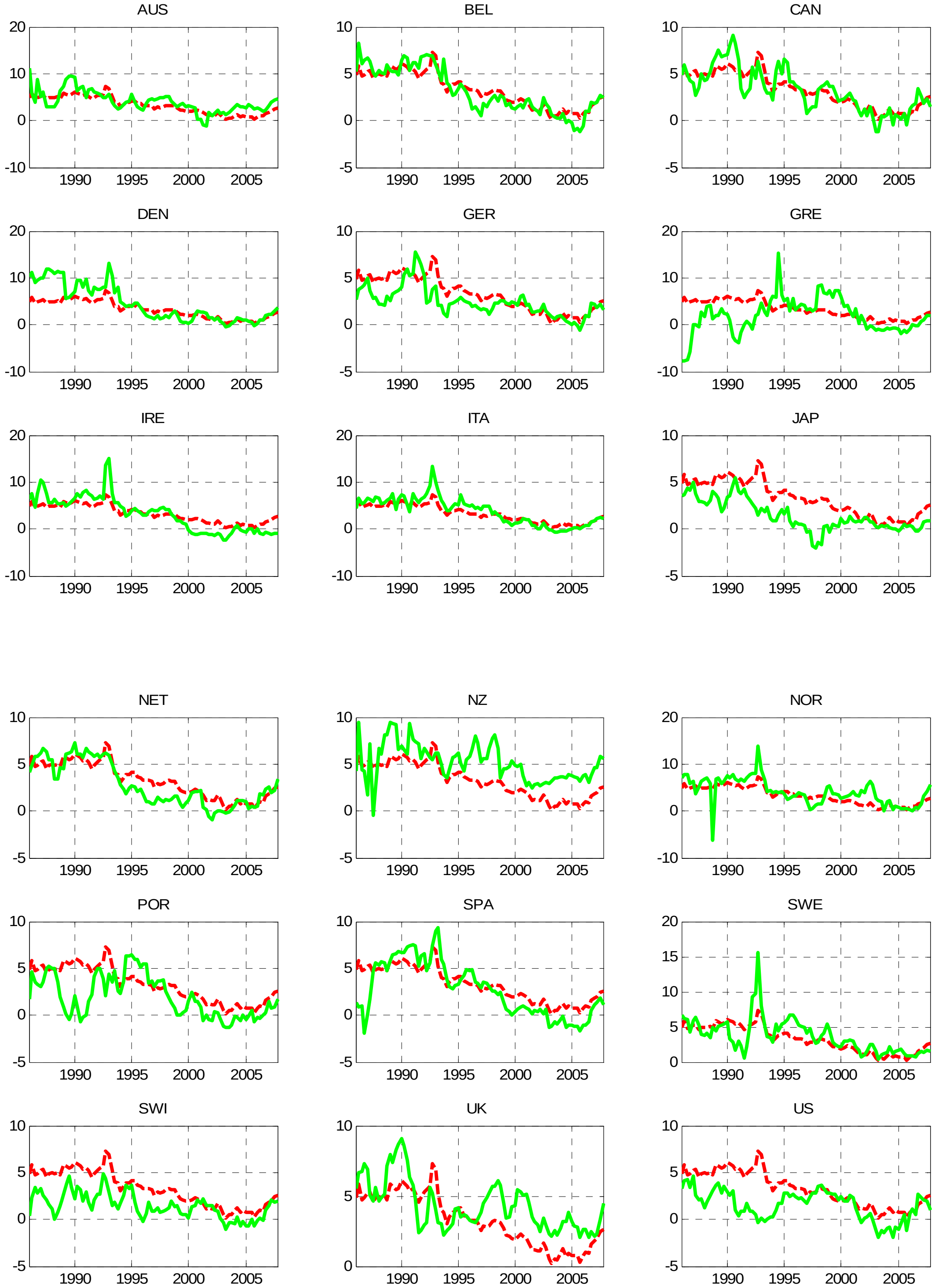


Figure 2: Real interest rates vs. PC estimate of world real interest rate



Note: The green line is the respective country's real interest rate; the dotted red line is the 1st principal component.

Figure 3: Changes in interest rate differential and average exchange rate depreciation

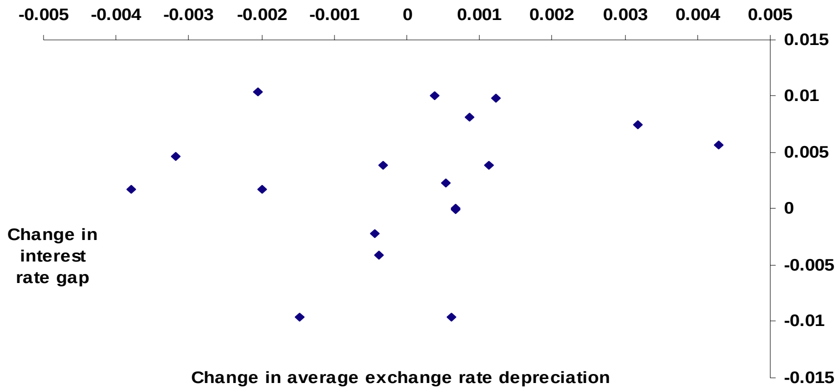


Figure 4: Changes in interest rate differential and economic volatility

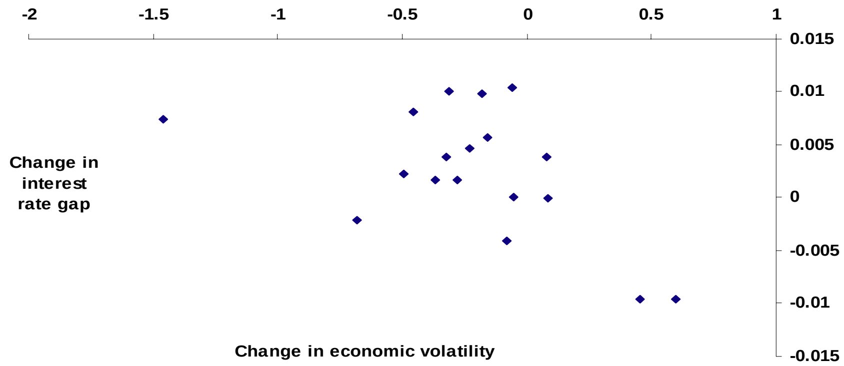


Figure 5: Changes in interest rate differential and net foreign asset position (% of GDP)