Measuring the Risk Premium in Uncovered Interest Parity using the Component GARCH-M Model

1. Introduction

With the development of international financial markets, financial instruments have contributed to international capital market integration by increasing capital mobility between developed and emerging countries. Therefore asset parity has become a vital consideration for all international investors. Uncovered interest parity (UIP) is one of the most important theoretical relations used in analytical work in both international finance and macroeconomics. It is also a key assumption in many of the models of exchange rate determination.

UIP implies that the interest rate differential should be equal to the exchange rate change. However, in reality, low interest rate currencies tend to depreciate relative to high interest rate currencies. This is inconsistent with UIP and has been confirmed by an extensive literature for different countries and periods. Overall there has been no consensus on how to explain the failure of UIP. A number of explanations for the deviations from UIP include the failure of rational expectations, the time-varying risk premium and nonlinear behavior. The time-varying risk premium is one of the most frequently cited reasons leading to the failure of UIP (see Froot and Thaler, 1990; McCallum, 1994; Meredith and Chinn, 2004). Therefore, it is necessary to continue investigating whether the time-varying risk premium could affect the validation of UIP especially over the periods of the Asian financial crisis and the recent credit crisis.

The two contributions of this paper are as follows. First of all, following the financial and credit crises over the sample period considered in this study, there has been a rapid change in risk across the world. To account for this, the CGARCH-in-mean model incorporating asymmetric adjustment is used to reflect the substantial rapid change in risk and separates out the permanent and transitory risk in the UIP condition. It is a superior volatility model for exchange rates, as it can distinguish the permanent and transitory volatility components to describe volatility dynamics better than other GARCH models (Black and McMillan, 2004; Guimarães and Karacadag, 2004; Byrne and Davis, 2005; Pramor and Tamirisa, 2006; Christoffersen *et al*, 2006; Guo and Neely, 2008 and Wei, 2009). Separating permanent and transitory risk is important to assessing whether this uncertainty is driven by macroeconomic fundamentals or by market sentiments, which will affect the investment strategies. This is the first time that the CGARCH model has been used to measure the risk premium in UIP, which could partly explain the UIP puzzle. Secondly, we select both developed and emerging countries for comparison. The majority of the literature on UIP concentrates on low inflation and floating exchange rate regime countries. However Flood and Rose (2002), Huisman *et al* (2007) and Ichiue and Koyama (2008) demonstrate that countries which have high exchange rate and interest rate volatility work better regarding UIP than others. Comparing the different UIP results between developed and emerging countries could help us to understand the volatility effect for both sets of countries.

The main result of this study is that including the risk premium in UIP improves the precision of the estimation, but it is still hard to explain the failure of UIP even using a sophisticated measure of risk. This study finds a significant risk premium in most countries, suggesting that risk is an important part of modeling exchange rates and needs to be considered in both empirical and theoretical models. The transitory shifts in financial market sentiment tend to be less important determinants of risk than shocks to the underlying macroeconomic fundamentals. In general, emerging countries work better in terms of UIP than developed countries.

The remainder of this paper is organized as follows. Section 2 presents the theory of UIP and the previous literature. In section 3, the method is described. Section 4 presents the data and the main empirical analysis in order to see whether UIP holds and Section 5 concludes and suggests further areas of study.

2. Uncovered Interest Parity and the Risk Premium

2.1 Uncovered Interest Parity

UIP suggests that the domestic currency is expected to depreciate when the domestic interest rate exceeds the foreign interest rate[1]. It is a non-arbitrage condition between investing in domestic currency denominated assets and foreign currency denominated assets, so it can be expressed as:

$$(1+i_{t,k}) = (1+i^{*}_{t,k})\frac{E_t S_{t+k}}{S_t} \tag{1}$$

where $i_{t,k}$ represent the domestic interest rate at time t of maturity k, $i_{t,k}^{*}$ is the foreign interest rate, S denotes the exchange rate which is the domestic currency price of a unit of foreign currency and E_t is the expectations at time *t*. Taking natural logarithms of the above equation (1) and imposing the rational expectation and risk neutrality assumptions to get the following empirical equation for UIP:

$$\Delta s_{t+k} = s_{t+k} - s_t = \alpha + \beta(i_{t,k} - i^{*}_{t,k}) + \varepsilon_{t+k} \tag{2}$$

where Δs_{t+k} is the change in the log of the spot exchange rate over k periods and $(i_t - i_t^{*})$ is the current k period home interest rate less the k period foreign interest rate. The

[1] As Holmes *et al.* (2012) note, the UIP literature and the literature on the international links between term structures are closely related, so the literature on these two topics share many common features.

null hypothesis for UIP is that $\alpha = 0, \beta = 1$. We also expect that the error term is Gaussian and stationary. However, most empirical evidence on developed economies suggests that exchange rate changes and interest rate differentials are negatively correlated, with high domestic interest rates predicting an appreciation. Froot and Thaler (1990) summarize the coefficient results from 75 published studies, most giving a negative β coefficient and those with positive coefficients having less than the hypothesized value of one, the average value of the coefficient is -0.88.

2.2 Risk Premium

Most empirical tests of UIP are based on assumptions of rational expectations and risk neutrality, one obvious explanation for the UIP failure is the existence of a time-varying risk premium. The time-varying risk premium is a part of the OLS residuals and its correlation with the exchange rate change causes the estimated beta coefficient to be biased. If market participants are risk averse, then the forward rate will equal the expected spot exchange rate plus a risk premium (Meredith and Chinn, 2004; Chinn, 2006). The risk premium δ is written as:

$$f_t = E_t s_{t+1} + \delta_{t+1} \tag{3}$$

If we assume that CIP holds ($f_t - s_t = i_t - i_t^*$), the equation could change to:

$$i_t - i_t^* = E_t s_{t+1} - s_t + \delta_{t+1} \tag{4}$$

We rearrange the equation (4) and get the following equation (5)

$$E_t s_{t+1} - s_t = i_t - i_t^* - \delta_{t+1} \tag{5}$$

In this situation, the interest rate differential could not be interpreted as the expected change in the exchange rate. The interest differential is equal to the expected change in the exchange rate plus a risk premium. In emerging markets, interest rates and price levels are more volatile than those in developed countries. Therefore, investors required more compensation for holding emerging market assets and the interest rates in emerging market have to be higher to maintain capital inflow. Under rational expectations, the UIP model considers the risk premium being expressed as:

$$s_{t+1} - s_t = i_t - i_t^* - \delta_{t+1} + \varepsilon_{t+1} \tag{6}$$

The empirical formulation of the currency risk premium, following previous research by Domowitz and Hakkio (1985) and Tai (1999) is defined as:

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + \gamma\sigma_{t+1} + \varepsilon_{t+1} \tag{7}$$

where σ_{t+1} is the conditional component of the standard deviation of the error term. The risk premium has a constant component (α) and a time-varying component, which is the conditional standard deviation. If both α and γ are insignificantly different from zero, there is no risk premium. If $\alpha \neq 0$ but $\gamma = 0$, there is a constant risk premium. Only when $\gamma \neq 0$ does the time-varying risk premium exist. The previous finding of $\beta < 0$ means that the increase in the interest rate differential is combined with the decline in the expected change in the exchange rate and a larger rise in the risk premium. Investors are demanding a large risk premium for holding risky currencies and expect the currency to appreciate rather than depreciate. Investors of the risky currencies are compensated by both higher interest rates and by currency appreciation[2].

Froot and Frankel (1990) use survey data on exchange rate expectations to decompose the deviations from UIP into deviations caused by expectation error and a time-

[2] This is consistent with the carry trade where investors borrow from low interest rate currencies and invest in countries with high interest rates, where they could benefit from the interest rate differential. In this case the UIP condition does not hold, as the low interest rate currency tends to appreciate. Therefore, investors could earn a profit from both the interest rate differential and the exchange rate change.

varying risk premium. They find that the largest part of deviations from UIP is caused by expectation error, while the time-varying risk premium plays a minor role. However, the results of MacDonald and Torrance (1989), Taylor (1989) and Cavaglia *et al*. (1993) indicate an important role for the time-varying risk premium rather than the expectation error. Froot and Thaler (1990) demonstrate in their paper that the risk premium significantly affects UIP. Fama (1984) and Anker (1999) provide evidence that when the correlation between the risk premium and the change in the exchange rate is negative, the estimated coefficient on the interest differential is less than zero. McCallum (1994) and Meredith and Chinn (2004) point out that the risk premium is the main reason leading to UIP failure over the short-term horizon. However, in the long term, exchange rates are determined by fundamentals. Berk and Knot (2001) following the seminal work of Engle et al. (1987), allow for a time-varying risk premium by estimating the UIP relationship as the conditional mean in an ARCH model. Poghosyan *et al.* (2008) test UIP in Armenia and find that UIP holds better than other studies and there exists a positive time-varying risk premium based on the GARCH-M model. Melander (2009) tests the risk premium in the UIP condition, which is measured by the GARCH-M model in Bolivia and provides evidence that UIP does not hold, but the deviation from UIP is smaller than before.

Earlier empirical studies on the UIP condition mostly focus on developed economies rather than emerging markets because of a lack of data. However, the deviations from UIP in emerging countries are likely to be larger and more persistent than in the developed economies. The increase in the degree of financial liberalization in emerging markets enabled many researchers to analyze the emerging foreign exchange market. Some recent studies indicated that the UIP condition might work better in emerging countries with fixed exchange rates than in developed countries. Flood and Rose (1996) find that UIP holds better for fixed exchange rates than floating exchange rates. But they said there is no theoretical reason to explain this difference of exchange rate regime change. Flood and Rose (2002) conclude that UIP works better for countries in crisis which have high exchange rate and interest rate volatility[3]. Bansal and Dahlquist (2000) find that the negative correlation between the expected currency depreciation and interest rate differentials is only present in developed countries where its interest rate is lower than the interest rate of the US, but not in the emerging countries. In other word, UIP tends to hold in emerging countries rather than developed countries. Frankel and Poonawala (2006) find small deviations from UIP in emerging countries.

The CGARCH model has been widely used recently in both economics and finance. It decomposes volatility into permanent and transitory components. Separating the permanent and transitory risk premium could help us to understand the source of uncertainty, because investment decisions heavily depend on whether this uncertainty is permanent or transitory (Byrne and Davis, 2005). Black and McMillan (2004) find evidence of short-run and long-run components in exchange rates, which exhibit different rates of volatility persistence and decay from a shock to volatility. They also find that the CGARCH specification provides a more adequate description of exchange rate volatility than a GARCH specification. Pramor and Tamirisa (2006) analyze exchange rate volatility trends in Central and Eastern European (CEE)

[3] There is also extensive literature on the related theme of UIP and speculation. For instance Craighead *et al*. (2010) discuss the 'extreme support' for UIP theory, where using the large outlier observations offers more support for UIP, although their evidence fails to provide complete support for this theory.

currencies and the euro and find the long-run volatility is mainly driven by shocks to economic fundamentals rather than shifts in market sentiment. Simón and Amalia (2011) examine the relationship in the volatility of sovereign yields using a CGARCH model to decompose permanent and transitory volatility. The results suggest that transitory shifts in debt market sentiment tend to be less important determinants of bond yield volatility than shocks to the underlying fundamentals.

There is a large body of literature providing evidence that the CGARCH model works better than the standard GARCH models. Christoffersen et al (2006) show that distinguishing between short-run and long-run components, enables the CGARCH model to describe volatility dynamics better than the standard GARCH model. Ghysels *et al* (2005) and Guo and Whitelaw (2006) state that better measures of conditional volatility might produce more precise estimates of the risk-return relationship. Guo and Neely (2008) use the CGARCH model to distinguish between the effects of the long-run and short-run volatility components in stock prices and the results favor the CGARCH model over the standard GARCH model. Consistent with the US evidence, the long-run volatility is a more important determinant of the conditional equity premium than the short-run for most international markets. Wei (2009) investigates the spillover effect of unexpected exchange rate shocks based on the symmetric and asymmetric CGARCH model and finds that the asymmetric effect is weakly statistically significant for all three exchange rate markets, although the forecasting performance of the symmetric CGARCH model outperforms the asymmetric version.

3. Methodology

The GARCH-in-mean (GARCH-M) model introduced by Engle *et al* (1987) was designed to capture the relationship between return and risk, such as with the CAPM. The applications of GARCH-M models to stock returns, interest rates and exchange rates can be found in Bollerslev *et al* (1992). We follow the paper of Berk and Knot (2001) and Melander (2009) and add the conditional standard deviation as a time-varying risk premium in the mean equation to construct the GARCH-M model. The GARCH-M model used in UIP empirical analysis is written as follows:

$$\begin{aligned} s_{t+1} - s_t &= \alpha + \beta(i_t - i_t^*) + \gamma\sigma_{t+1} + \varepsilon_{t+1} \\ \sigma_{t+1}^2 &= \delta_0 + \varphi_1\varepsilon_t^2 + \varphi_2\sigma_t^2 \end{aligned} \tag{8}$$

where σ_{t+1} is the standard deviation of the error term and denotes the time-varying risk premium that directly affects the exchange rate.

The GARCH class model used in this paper is the component GARCH (CGARCH) model proposed by Engle and Lee (1999) as many researchers find it is a superior volatility model. They extended the GARCH model to ensure that the volatility is not constant in the long-run and decomposed volatility into two components, the long-run trend[4] and short-run deviations from that trend. The two components of volatility are typically interpreted as driven by different factors: the long-run trend in volatility as reflecting shocks to economic fundamentals, and transitory volatility being driven by market sentiment and short-term position-taking. However as yet no one has used it to test the risk premium in UIP. This paper includes the asymmetric term in the CGARCH-M model based on the GJR GARCH model by Glosten, Jagannathan and Runkle (1993) to examine the difference in volatility associated with exchange rate

[4] An alternative to the CGARCH model for long memory in conditional variance has been provided by the FIGARCH model.

depreciation and appreciation. The model is described by the following set of equations:

$$\begin{aligned} s_{t+1} - s_t &= \alpha + \beta(i_t - i_t^*) + \gamma\sigma_{t+1} + \varepsilon_{t+1} \\ q_{t+1} &= \varphi_1 + \varphi_2(q_t - \varphi_1) + \varphi_3(\varepsilon_t^2 - \sigma_t^2) \\ \sigma_{t+1}^2 &= q_{t+1} + \varphi_4(\varepsilon_t^2 - q_t) + \varphi_5 D_t(\varepsilon_t^2 - q_t) + \varphi_6(\sigma_t^2 - q_t) \end{aligned} \tag{9}$$

where D_t is a dummy variable for the asymmetric effect indicating unexpected exchange rate appreciation, $D_t=1$ for $\varepsilon_t<0$, $D_t=0$ otherwise, q_{t+1} is the long-run component of the conditional variance which reflects shocks to economic fundamentals and converges to the long-run time-invariable volatility level φ_1 with a magnitude of φ_2. This permanent component thus describes the long-run persistent behavior of the variance. The long-run time-invariant volatility level φ_1 can be viewed as the long-run level of returns variance for the relevant sector when past errors no longer influence future variance in any way and (σ_{t+1}^2-q_{t+1}) is the short-run component which is more volatile and driven by market sentiment. In the long-run component of volatility equation, the AR coefficient (φ_2) of permanent volatility should exceed the coefficients (φ_4+φ_6) in the transitory component which then implies that the model is stable and short-run volatility converges faster than the long-run. The closer the estimated value of the φ_2 is to one the slower q_{t+1} approaches φ_1, and the closer it is to zero the faster it approaches φ_1. The value of φ_2 therefore provides a measure of the long-run persistence. The coefficient of the forecast error φ_3 shows how shocks affect the permanent component of volatility.

In several previous instances[5], the coefficient of the autoregressive term in the long-run trend equation is equal to or very close to one. We include an asymmetric term in

[5] These include the previous studies of Black and McMillan, 2004; Byrne and Davis, 2005 and Pramor and Tamirisa, 2006.

the transitory variance model allowing asymmetric impacts of past shocks (relative to long-run volatility) on the short-run volatility. If the coefficient φ_5 is less than zero, then the impact of negative shocks (unexpected domestic currency depreciation, $\varepsilon_t>0$) on short-run volatility is greater than the impact of positive shocks (unexpected appreciation). In other words, the unexpected depreciation increases short-run volatility. The reason we add an asymmetric term into the CGARCH-M model follows the work of Guimarães and Karacadag (2004), who modeled the exchange rate volatility in emerging market currencies and find significant asymmetric effects for the Mexico peso and Turkish lira, while Byrne and Davis (2005) and Pramor and Tamirisa (2006) also find significant negative asymmetric effects.

4. Data and Empirical Analysis

The data set consists of monthly exchange rates and interest rates[6] for developed and emerging countries, which has been collected from Datastream and the Bank for International Settlements (BIS). The exchange rate data is the first day's value for each month. The 1 month interest rate data are collected from Datastream. The time periods for the various countries are different due to data limitations. Table 1 includes details on the samples, exchange rate regimes and dates of any financial crisis outside of the recent 2008 crisis.

Figure 1 represents the measures of risk for each country over the sample through the conditional standard deviation plots and as mentioned above they tend to follow individual country events. For the UK the main jumps in volatility occur during the membership of the European Exchange Rate Mechanism (ERM) and its exit in

[6] Both domestic and foreign interest rates are annualized.

September 1992, as well as the 2008 financial crisis. Similarly with Brazil, there are spikes for the foreign exchange crisis in 1999 when the Brazilian Real left its soft peg and moved to a managed float as well as 2008 again. The main increases in risk in other countries also follow moves away from soft peg exchange rates, such as Thailand 1997 and 1998 and Russia in 1998. The other countries follow similar patterns, although the non-American emerging economies do not suffer such high volatility in 2008 as the financial crisis was not as severe in these countries, as there banking systems was better mostly capitalized. Also Japan and Switzerland have the most stable plots of both long and short run risk as they have not been involved in currency based crises and in the late 2000s had largely escaped the worst of the financial crisis.

Table 2 presents the OLS results based on equation (2)[7]. The range for the β coefficient is from -2.1875 to 1.0768. The results for the developed countries are quite similar to previous empirical studies which have negative and insignificant β coefficients. The β coefficients from emerging countries are positive but mostly insignificant. Only Russia gives us positive and significant results. The p-values from the Wald test demonstrate that UIP is valid for the UK, Canada, Switzerland and Thailand at the 10% level, but the coefficient β is insignificant. The misspecification and diagnostic tests show that OLS is not the best model for UIP testing because there

[7] There tends to be a negative correlation between the interest differential and risk measure for the developed countries but not the developing ones, which supports the finding that UIP tends to hold in developing countries but not the more developed ones. The correlation between the interest rate differential and risk premium as follows: UK (0.3205), Australia (-0.0209), Japan (-0.0588), Canada (-0.2409), Switzerland (-0.2309), Brazil (0.0660), Mexico (0.5391), Malaysia (0.2487), Thailand (0.3707) and Russia (0.7419)

are problems such as serial correlation[8] and the ARCH effect[9]. In the following sections we will test GARCH class models to decide whether the risk premium affects the UIP condition.

Table 3 displays the unit root test results for the exchange rate change and interest rate differentials. The Augmented Dickey-Fuller (ADF) test provides strong evidence that the exchange rate change is stationary at the 1% confidence level. However the unit root results for the interest rate differential are contradictory. From the ADF test, six out of ten countries are stationary at the 10% level, but with the DF-GLS test it is only three countries and with Ng-Perron it is five. This result indicates that the interest rate differential is more persistent than the exchange rate change. It is also an interesting finding that the exchange rate change is always stationary but the interest rate differential is nonstationary which is consistent with other studies (de Brouwer, 1999 and Goh, Lim and Olekalns, 2006). Further unit root research on the interest rate differential based on the Zivot-Andrews test and TAR/M-TAR tests find that the interest rate differential is stationary with the structural break or the asymmetric adjustment[10].

Table 4 illustrates the estimated risk-adjusted UIP results[11] from a CGARCH-M

[8] The result of the Q-statistic for serial correlation is not shown in Table 1. The p-value for all 36 lags is nearly zero which means that there is autocorrelation between the residuals.

[9] We also carried out tests for the ARCH effects, but the results were mixed. We have not placed much emphasis on this as the test is for only the basic ARCH effect, so failure to find any evidence of it doesn't mean more complex forms of ARCH effect do not exist, such as ARCH with asymmetric adjustment.

[10] The results from Zivot-Andrews test and TAR/M-TAR tests are available from the authors on request.

[11] The GARCH-M results seem to have a change in standard errors rather than a change in the magnitude of the coefficients.

model[12] from equation (9) under the generalized error distribution (GED) using maximum likelihood estimation (MLE). The intercept is significant at the 10% level except Australia, Canada, Brazil and Mexico, which means that there is a constant risk premium. This intercept takes account of each country's specific financial systematic risk such as the liquidity of the foreign exchange market. It is apparent that the β coefficients of the interest differential from developed countries are all negative and significant except Canada. This negative coefficient means that an increase in the interest differential will lead to a decrease in the expected change of the exchange rate. This is consistent with previous literature (Froot and Thaler, 1990). However, the β coefficients are positive and significant for three of the five emerging countries. Compared with the results from the OLS, the β coefficients for Thailand and Russia have increased and are close to one. The coefficients of the time-varying risk premium are significant in seven out of ten cases. The negative γ coefficient corresponds to the mean-variance theory or alternatively expected utility theory. It implies that when there is an increase in risk, the deprecation of the home currencies decreases, and the expected return from holding this home currency increases. The risk averse investors required more return when they face higher risk. This negative coefficient for the risk premium is also found by Melander (2009) when testing the UIP condition with a basic GARCH model in Bolivia. The Wald test for no risk premium is rejected for all countries except for Australia, Canada and Mexico at the

[12] We have also measured the risk premium based on the GARCH-M model, but the results are similar to CGARCH-M model and the CGARCH-M model is better overall than the GARCH-M model. The results from the GARCH-M model are not included but are available from the authors on request. To confirm that the CGARCH-M results are better than the GARCH-M, if both $\varphi_2=\varphi_3=0$ (equation 9), the CGARCH model will reduce to the standard GARCH(1,1) model with a constant long-run volatility trend. In the Wald test results in Table 4, only Japan fails to reject the null hypothesis and we conclude that the CGARCH-M model is better than the GARCH-M model and there exists time-varying long-run volatility.

10% level. The lack of a risk premium was also found by Domowitz and Hakkio (1985), who could not reject the null hypothesis of no risk premium for the currencies of five industrial countries using an ARCH-M model and with Baillie and Bollerslev (1990) who fail to find a time-varying risk premium for four European countries based on a multivariate GARCH model. The insignificant risk premium coefficients of Australia, Canada and Mexico may result from either a poor measure of risk or the misspecification of the model. In other words, the conditional standard deviation may not be the proper measure of risk or the univarite GARCH-M model is not an appropriate econometric model to estimate the risk premium. But there is still an improvement when testing UIP including a risk premium using a CGARCH-M model rather than the basic OLS model.

In the long-run trend equation, we find all countries have a positive and significant long-run average volatility φ_1 except for Canada, Brazil and Thailand, but the magnitude is extremely small and nearly zero. The coefficient φ_2 of the lagged permanent volatility is large and highly significant at the 1% level for all countries except for Japan, confirming the presence of long-run volatility persistence. In particular, the coefficient is close to one which means the trend persistence is very high and the permanent volatility converges to its mean level slowly. These results indicate that permanent conditional volatility exhibits long memory. Half of the forecast error parameters (φ3) are significant, capturing the influence of the time-dependent movement of the permanent component.

As for the transitory components, the coefficient φ_4 measured the initial impact of a shock to the transitory component of the CGARCH model, and it is significant in three out of the ten countries. The coefficient φ_6 indicates the degree of memory in the transitory component, which is significant in two out of ten cases. While the shock persistence in the transitory components measured by the sum of the (φ_4+φ_6) is lower than the coefficient (φ_2) of the lagged permanent volatility, which implies that our model is stable and mean reversion is slow in the long run. Therefore, the long-run volatility is more persistent than the short-run. The larger long-run volatility component indicates that the risk premium is mainly driven by shocks to economic fundamentals rather than shifts in market sentiment[13]. This result is similar to previous literature, such as Black and McMillan (2004), Guimarães and Karacadag (2004), Byrne and Davis (2005) and Pramor and Tamirisa (2006). The asymmetric coefficient (φ_5) is negative and significant in the UK, Australia and Brazil which implies unexpected domestic currency depreciations have a larger effect on volatility than unexpected appreciations. This is consistent with the currency crisis explanation and this finding is in line with the findings of Byrne and Davis (2005). But in contrast, Japan, Switzerland, Mexico, Malaysia and Thailand have significantly positive coefficients indicating that an unexpected appreciation has a larger effect on volatility. The explanation for the Japanese yen and Swiss francs is that they are funding currencies for the carry trade and appreciate during periods of high volatility due to the unwinding of the trade. The finding of a positive sign on the asymmetric term could be explained by studies on emerging stock markets. For instance Koutmos *et al*. (1993) suggested that positive shocks have a greater effect on volatility than negative

[13] We assume that fundamentals drive long-run volatility and market sentiment the short-run, as with the other studies mentioned, but measuring this specifically is beyond the scope of this study. In addition it should be noted that these fundamentals may experience temporary changes. As an anonymour referee noted, these could include oil price changes due to military events or supply –chain interruptions due to Tsunamis. These are different to the types of change indicated by the term market sentiments.

shocks for the Athens stock market and could possibly be explained by investors perceiving excessive rises in asset prices as evidence of a speculative bubble, facilitating a rise in uncertainty and an associated increase in volatility. This scenario could equally arise in the emerging foreign exchange markets.

Figure 1 shows the estimated transitory and permanent component of volatility based on the CGARCH-M model[14]. The transitory component of the volatility is much smaller than the permanent volatility for all the countries. And in most countries, the transitory component is much more volatile than the permanent component. The transitory volatility is driven by market sentiment which is related to short-run speculative pressures. The permanent volatility is based on the fundamentals of the macro economy, such as the goods markets, where it is assumed adjustment takes a longer while than with the transitory volatility, due to the usual inertia in such markets. This implies that transitory shifts in financial market sentiment tend to be less important determinants of volatility than shocks to the underlying macroeconomic fundamentals. It is consistent with the results from Black and McMillan (2004), Byrne and Davis (2005) and Pramor and Tamirisa (2006), who analyze developed countries exchange rate volatility. In Figure 1 the times of particular volatility increases have been identified and as expected they occur mostly at times of financial crises, especially the recent financial crisis which culminated in the collapse of Lehman

[14] The conditional standard deviation graphs and estimation results based on the GARCH-M model are available from the authors on request, but have been omitted as they are generally similar to the C-GARCH-M results. Canada and Russia are two special cases for the GARCH-M graphs. Canada appears to be on an exploding path because this model could not satisfy the non-negativity and stability constraints of the GARCH-M models. (the α coefficient is negative, β coefficient is greater than one and sum of them is larger than one). Russia looks relatively smooth except a huge spike during the 1998 crisis due to the negative α coefficient.

Brothers in September 2008. During the crises periods the short-run volatility approximates much more closely to the long-run which reflects the importance of short-term turbulence in the international financial markets, such as Brazil in 1999 M2, Australia in 2008M10.

We also find that developed countries have relatively low conditional standard deviation compared with other emerging countries. Emerging countries are more volatile, especially during the crisis period. This is reasonable because of their limited ability to conduct counter-cyclical monetary and fiscal policy and non-credible monetary institutions and weak fiscal position. (Hausmann *et al*., 2006). Combined with the results of the β coefficients, we find that UIP works differently for countries whose exchange rate and interest rate display high or low standard deviation. The result found that UIP is more likely to be held when there is high exchange rate volatility, which is consistent with Mahieu (2007) and Ichiue and Koyama (2008).

5. Conclusion

The main finding of this paper is that the risk premium is significant in most countries. Including the risk premium in the UIP condition improves on the original model, as the β coefficient is much more significant with a risk premium included in the model than in the basic OLS model, although UIP still does not hold in many countries. This result suggests that risk is an important part of modeling the exchange rate and needs to be considered in both empirical and theoretical models. In addition the risk needs to be considered in terms of the permanent and transitory components, where the permanent component is found to have the greatest effect, suggesting it is volatility from the macroeconomic fundamentals that are the primary determinant of exchange rates. This study also finds that in general emerging countries work better in terms of UIP and the inclusion of the risk premium than developed countries. The β coefficients in emerging countries, such as Brazil, Russia and Thailand are positive and close to unity. Moreover, the CGARCH-M model works better with UIP, in terms of modeling the risk premium as it considers both the long-run and short-run volatility components. However other aspects of these results are more disappointing in that the addition of the risk premium does not improve the signs or magnitudes of the interest differential in many of the countries, which suggests that the addition of a time varying risk premium doesn't completely solve the UIP puzzle.

Further research could incorporate a longer data span as more data becomes available. It could also develop the CGARCH-M model by including both long-run and short-run volatility in the mean equation and not just consider the conditional standard deviation. It would then be interesting to test the validation of the UIP condition with both a short-run and long-run risk premium.

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Table 1

Country	*Data Sample begins*	*Crisis years (all + 2008)*	*Exchange Rate regimes (pre and post crisis)*
UK	1986	1992	Soft peg /managed float
Australia	1986	-	Managed float
Japan	1986	-	Managed float
Canada	1990	-	Managed float
Switzerland	1986	-	Managed float
Brazil	1994M11	1999	Soft peg /managed float
Mexico	1994M4	1994/95	Soft peg /managed float
Malaysia	1985M11	1997/98	Soft peg
Thailand	1992M2	1997/98	Soft peg /managed float
Russia	1995M1	1998	Soft peg /managed float

Table 2 OLS Results

			Wald test	
	α	β	χ^2	***p-value***
UK	-0.0002	-0.0295	1.3946	0.4979
Australia	0.0012	-0.8156	5.1147	0.0775
Japan	-0.0070**	-2.1875**	9.8570	0.0072
Canada	-0.0001	-0.0485	2.0281	0.3627
Switzerland	-0.0031	-1.3372	3.7904	0.1503
Brazil	0.0031	0.1571	12.0729	0.0029
Mexico	0.0058	0.1821	5.5142	0.0635
Malaysia	0.0015	0.4763	6.3549	0.0417
Thailand	-0.0010	1.0768	0.2349	0.8892
Russia	0.0023	0.5836***	5.5384	0.0627

Note: The OLS result is run by the equation (2). The Wald test is a joint test of null hypothesis H_0: $\alpha=0$, $\beta=1$. ***, ** and * denote statistical significance at 1%, 5% and 10% level.

Table 3 Unit Root Tests

	Exchange Rate Changes				***Interest Rate Differentials***			
	ADF	***DF-GLS***	***Ng-Perron***		***ADF***	***DF-GLS***	***Ng-Perron***	
	Intercept	*Intercept*	MZ_α	MZ_t	*Intercept*	*Intercept*	MZ_α	MZ_t
UK	-7.6202***	-1.2332	-2.6495	-1.0963	-2.6211*	-1.2106	-5.1906	-1.4859
Australia	-10.496***	-2.2835**	-3.9309	-1.1178	-1.6213	-0.3390	-0.2336	-0.2055
Japan	-3.8790***	-1.0809	-1.3892	-0.8327	-2.7543*	-1.8148*	-7.6415	-1.8999
Canada	-3.8679***	-5.3184***	-44.293	-4.4080	-2.8524*	-0.4195	-0.7065	-0.4507
Switzerland	-11.315***	-1.7344	-3.6724	-1.3312	-2.4776	-1.7575	-9.9998	-2.1697
Brazil	-10.987***	-10.982***	-83.901	-6.4612	-3.8831***	-0.1359	-0.1915	-0.1401
Mexico	-4.2313***	-3.0826***	-11.089	-2.3265	-1.8171	-1.6341*	-6.6102	-1.7619
Malaysia	-4.7127***	-1.3354	-2.0906	-0.9565	-2.9323**	-1.7024*	-5.8248	-1.7063
Thailand	-5.6943***	-4.8595***	-32.774	-4.0451	-2.1471	-1.4932	-6.5945	-1.7772
Russia	-5.6917***	-1.6421*	-4.7981	-1.4227	-4.3745***	-0.1886	-0.1083	-0.0951

Note: ADF test use the general to specific approach to select the number of lags. DF-GLS and Ng-Perron tests use modified information criteria (MIC). ***, ** and * denote statistical significance at 1%, 5% and 10% level.

Table 4 CGARCH-M Model

	UK	*Australia*	*Japan*	*Canada*	*Switzerland*
α	-0.0381*	0.0016	0.0569***	0.0052	0.0722*
β	-2.3395***	-1.4183**	-2.9683***	-0.7102	-2.0599**
γ	1.8824**	0.0105	-2.5817***	-0.3066	-2.7819*
φ_1	0.0005***	0.0007**	0.0006***	0.0025	0.0007***
φ_2	0.7055***	0.9639***	0.1393	0.9992***	0.6250***
φ_3	0.1136	0.0816	0.0021	0.0395*	-0.1017
φ_4	0.0964	0.2624*	-0.1195	0.0043	0.0458
φ_5	-0.2666**	-0.3783*	0.2271***	0.1407	0.2800**
φ_6	-0.1941	0.2803	0.0964	0.2125	-0.0357
$\varphi_2=\varphi_3=0$	0.0000	0.0000	0.9999	0.0000	0.0000
$\beta=1$	0.0000	0.0001	0.0000	0.0063	0.0004
$\alpha=\gamma=0$	0.0970	0.6589	0.0000	0.3127	0.0021
	Brazil	***Mexico***	***Malaysia***	***Thailand***	***Russia***
α	0.0036	-0.0020	-0.0046**	-0.0082***	0.1315***
β	0.2574***	-0.0862	0.2143	1.1516***	0.6108***
γ	-0.2133***	0.1317	0.3885***	0.3569***	-2.3717***
φ_1	0.0010	0.0010**	0.0002***	0.0005	0.0037***
φ_2	0.9778***	0.8548***	0.8859***	0.9883***	0.9123***
φ_3	0.2767***	0.1935	0.1887***	0.1595*	-0.0044**
φ_4	0.2495**	0.0630	0.1101**	0.0932	-0.0006
φ_5	-0.3938**	0.4567**	0.1037**	0.3445*	0.0521
φ_6	0.4170	0.2986	-0.6065***	-0.2535***	0.7587
$\varphi_2=\varphi_3=0$	0.0000	0.0000	0.0000	0.0000	0.0000
$\beta=1$	0.0000	0.0000	0.0210	0.5061	0.0269
$\alpha=\gamma=0$	0.0178	0.3685	0.0000	0.0000	0.0000

Note: the CGARCH-M model is expresses in equation (9). The p-value of the Wald tests of $\beta=1$ and $\alpha=\gamma=0$ and $\varphi_2=\varphi_3=0$ are in the table. ***, ** and * denote statistical significance at 1%, 5% and 10% level.

Figure 1 The Conditional Standard Deviation using CGARCH-M