

Common and Idiosyncratic Factors of the Exchange Risk Premium in Emerging European Markets*

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Abstract:

Existing empirical evidence suggests that the Uncovered Interest Rate Parity (UIRP) condition may not hold due to an exchange risk premium. For a panel data set of eleven emerging European economies we decompose this exchange risk premium into an idiosyncratic (country-specific) elements and a common factor using a principal components approach. We present evidence of a stationary idiosyncratic component and nonstationary common factor. This result leads to the conclusion of a nonstationary risk premium for these countries and a violation of the UIRP in the long-run, which is in contrast to previous studies often documenting a stationary premium in developed countries. Furthermore, we report that the variation in the premium is largely attributable to a common factor influenced by economic developments in the United States.

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1. Introduction

Emerging countries are typically perceived to be susceptible to economic risk and uncertainty to a greater extent than industrial countries. In particular, external shocks create a ‘fear of floating’ exchange rate regime in emerging economies (Calvo and Reinhart 2002). Given the importance of external volatility this paper empirically analyzes the behaviour of the foreign exchange risk premium for emerging European markets. While many of these countries are already members of the European Union (EU),¹ they vary in the degree of economic development, integration and progress to membership of the European single currency union, the euro zone. Some countries have already abandoned their own currency and adopted the euro while others are in the process of joining the euro zone.² The reduction in exchange rate uncertainty, particularly among member countries, is one motivation for a candidate country to join the euro zone (see Darby et al. 1999 and Byrne and Davis 2005).

Our definition of the exchange risk premium is closely related to international parity conditions which dominate the literature in international finance. For example, the expected change in the exchange rate should be equal to the interest rate differential according to the Uncovered Interest Rate Parity condition (UIRP). However, evidence suggests that the UIRP does not always hold (see Lewis 1995, Engle 1996, and Chinn 2006) and one pervasive explanation of the failure of this relationship is the existence

¹ There are 27 EU member countries. These are Austria, Belgium, Bulgaria, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Luxembourg, Malta, the Netherlands, Poland, Portugal, Romania, Slovakia, Slovenia, Spain, Sweden, and the United Kingdom. There are three EU candidate countries (Croatia, Former Yugoslav Republic of Macedonia, and Turkey).

² There are 15 member states of the EU using the euro (Austria, Belgium, Cyprus, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Malta, the Netherlands, Portugal, and Slovenia, and Spain).

of a risk premium and risk averse investors. Frankel (1982) was an early attempt to model risk in the foreign exchange market using an extended static CAPM approach.³ More recently, Carriero (2006) finds evidence of a stationary but time varying risk premium between the UK and the US when testing UIRP.

In contrast, evidence in Bansal and Dahlquist (2000) suggest that the UIRP relationship fails mainly for developed economies. UIRP holds to a greater extent for countries with a lower per capita incomes, higher inflation uncertainty and lower credit ratings. In related literature Frankel and Poonalwala (2006) find evidence that the forward exchange rate is a less biased predictor of future spot rates in 14 emerging countries compared to industrial countries.⁴ Furthermore, some evidence is obtained that the UIRP tends to hold with longer maturity returns (Flood and Taylor 1997, Meredith and Chinn 1998, and Chinn 2006) although such long maturity assets are often not available in emerging markets. Finally, Bekaert et al. (2007) suggest the evidence against the UIRP is mixed and depends upon the currency. Given this suggestion, this paper seeks to examine the nature of the risk premium in emerging market economies.

This paper's main innovation is to decompose the foreign exchange risk premium in emerging market economies into the common and idiosyncratic (country-specific) components following the approach of Bai and Ng (2004), and identify economic factors influencing the common factor. The idiosyncratic component of the risk premium is unique to each country and therefore likely explains the heterogeneity in the risk premium across countries. If the country-specific component in the risk premium is

³ Cochrane and Piazzesi (2005) identify a time varying risk premium in the bond market using a single factor model which also has information useful for predicting bond excess returns using one-to five-year maturity bonds.

⁴ See also the discussion in Chinn (2006).

relatively smaller than the common component, it follows that a country may have less opportunity to control the overall risk premium by itself.

On the other hand, the common factor is the risk premium prevailing among a group of countries and may be significant since most countries selected for this study are already members of the EU and their economies influence each other through international trade, financial integration and immigration. However, some countries in our study have only recently joined the EU or are not members (e.g. Bulgaria, Romania, Russia, Turkey and Ukraine) and additionally we seek to examine whether they display heterogeneous behaviour.⁵ Indeed, our methodology is sufficiently flexible to provide insight into the relative importance of common and idiosyncratic shocks in each country.

Furthermore, in our research setting, the US is the benchmark country and thus her economic and financial developments may become exogenous shocks common to the emerging markets. Using a VAR and theoretical general equilibrium modeling, Uribe and Yue (2006) identified that the US interest rate impacts upon interest rate spreads in emerging market economies, which consequently have real effects in these economies. Neumeyer and Perri (2005) propose that monetary conditions in emerging economies are dependent upon US interest rates and international factors drive country risk in emerging market economies. Monetary conditions in the US would appear to be important for emerging market economies. The significant size of the common component may also become a measure of economic and financial integration because it likely indicates that the economy is highly open to other countries.

⁵ The Czech Republic, Estonia, Latvia, Lithuania, Slovenia and Poland became EU members in 2004, and Bulgaria and Romania joined in 2007. Non-members, Croatia and Turkey initiated negotiation in 2005 for EU accession.

For this purpose, we employ a nonstationary panel econometric approach which pools the time-series data of emerging European countries. This approach enables us to distinguish the common and idiosyncratic foreign exchange risk premiums, and sets this paper apart from previous studies since they often investigated the total risk premium in a univariate (or time-series) context without considering commonalities with other countries (see Section 2). In addition to the technical issues, our paper is innovative since as far as we are aware we are the first to apply this methodology to the analysis of the UIRP and we seek to provide additional evidence with respect to the UIRP. Finally we consider whether the common element in emerging economies' foreign exchange risk premium is related to US monetary policy, consistent with the evidence in Uribe and Yue (2006), that US interest rates have an impact on emerging countries' economy.

The rest of the paper is organized as follows. Section 2 reviews relevant literature and explains the definition of the exchange risk premium employed in this study. Section 3 describes our data set and conducts preliminary analysis. The decomposition of the risk premium into common and idiosyncratic components is carried out in Section 4 using the recently developed nonstationary panel econometric approach (Bai and Ng 2004). This section also analyzes economic factors influencing the common factor of the risk premium. Finally, our main findings are summarized in Section 5.

2. The Exchange Risk Premium

The definition of the exchange risk premium may differ somewhat depending on the researcher. In the absence of data on the forward exchange rate and survey-based expectations on the exchange rate, we derive the exchange risk premium using the UIRP condition.⁶ In this section, we shall explain the risk premium focusing on the statistical characteristics of the premium.

Let us begin with the Covered Interest Rate Parity (CIRP) condition, a longstanding and venerable concept in international finance. The CIRP utilizes the forward market which provides investors with an opportunity to hedge against risks of currency fluctuations. Because hedging risks is important for traders in flexible exchange rate regimes, much research has been conducted using industrial countries and equation (1) below, particularly since the breakdown of the fixed exchange rate regime in the early 1970s.

$$1 + i_t = (1 + i_t^*)F_t/S_t \quad (1)$$

where F_t and S_t are the current period t forward and spot exchange rates, whilst i_t and i_t^* are the domestic and foreign interest rates respectively.⁷ Thus, according to the CIRP, the forward premium, which is the difference between the forward and spot rates, is explained by the interest rate differential.

What happens when the foreign exchange rate risk is not covered? This leads us to the concept of the Uncovered Interest Rate Parity (UIRP) hypothesis. The UIRP shows that the risk-neutral investor is indifferent to investing in identical financial assets except for

⁶ See Isard (1992).

⁷ In this study the exchange rate is defined as domestic currency units per unit of foreign currency. The US is the foreign country hence an asterisk denotes US interest rates.

currency denomination, and such a relationship can be summarized as follows:

$$1 + i_t = (1 + i_t^*)E_t S_{t+k}/S_t \quad (2)$$

Here the forward rate in equation (1) is replaced with the expected spot rate, $E_t S_{t+k}$, which is the expected value of the spot exchange rate at time $t+k$ given the information available at time t . Expressing this equation in log form, $\ln(S_t) = s_t$, where $\Delta E_t s_{t+k}$ is the expected change in the spot rate, and ignoring Jensen's inequality term, we have:

$$i_t - i_t^* - \Delta E_t s_{t+k} = 0 \quad (3)$$

When equation (3) holds, the asset portfolio is in equilibrium and there is no capital movement across countries. However, this condition may not hold due to deviations from UIRP. In that case, equation (4) may be more useful.

$$rp_t = \Delta E_t s_{t+k} - i_t + i_t^* \quad (4)$$

In equation (4) rp_t measures the deviation from the UIRP. If $rp_t < 0$, the home country experiences capital inflow. On the other hand, if $rp_t > 0$, the home country faces capital outflow. Thus, deviation from the UIRP is often used to measure international capital mobility or capital market integration across countries (see Obstfeld and Taylor 2004). As discussed in the introduction, a number of researchers have investigated the UIRP and provided evidence against this condition (Engle 1996 and Chinn 2006). The most common explanation of the deviation from the UIRP is the time-varying risk premium that separates the spot and forward rates. Other factors contributing to the UIRP violation may include political risk, default risk, differential tax risk and market liquidity—which make financial assets in two countries imperfect substitutes (Hallwood and MacDonald 2000).

Since the expected value of exchange rate changes is unobservable, previous studies have transformed equation (4) to account for this as follows

$$risk_t = rp_t - \varepsilon_t = s_{t+k} - s_t - i_t + i_t^* \quad (5)$$

where ε_t is an expectations error ($\varepsilon_t = E_t s_{t+k} - s_{t+k}$) which follows a white noise process ($\varepsilon_t \sim iid(0, \sigma)$) when the investors form rational expectations. In the absence of data on the forward exchange rate and survey-based expected exchange rate in most emerging markets, we will employ equation (5) and use $(rp_t - \varepsilon_t)$ as a proxy for the foreign risk premium.

One may expect that our proxy for the foreign risk premium follows a stationary process since there is mounting evidence of a stationary risk premium. For example, using the data of industrial countries, Taylor (1987) and Carriero (2006) show that a combination of the expected exchange rate change and the interest rate differential yields a stationary time-varying risk premium. Furthermore, Kasman et al. (2008) provide evidence of stationary interest rate differentials between Germany and emerging markets such as Croatia, Estonia and Turkey. While the stationarity of interest rate spreads does not ensure the stationary premium, this becomes evidence of the high level of financial market integration.

3. Data and Preliminary Study

The data used in this paper is obtained from the International Monetary Fund's *International Financial Statistics (IFS)* for the sample period January 1998 to February 2008. This sample period is determined by data availability, and monthly frequency is chosen since lower frequency data (i.e., quarterly and annual data) provides us with a

smaller number of observations to conduct the univariate (or time-series) analysis of the common factor. Our data set includes the following emerging European economies: Bulgaria, Croatia, Czech Republic, Estonia, Latvia, Lithuania, Romania, Russia, Slovenia, Turkey and Ukraine.⁸

The composition of our group of countries is a unique aspect of this research. Many emerging economies in this study only recently became EU members, and incidentally, these countries changed their exchange rate regimes during our sample period. Thus, in order to better interpret our empirical results, we shall very briefly review their regimes here, which can be categorized broadly into three groups: floating, intermediate and fixed.⁹

For example, Bulgaria adopted a floating exchange regime (1990-1997) and shifted to a currency board arrangement. Estonia, Latvia, and Lithuania had fixed their exchange rate during our sample period. Notably, Lithuania's exchange rate (the litas) was fixed against the US dollar until February 2002. When the euro was introduced, these four countries fixed their exchange rate against the euro with the exception of Latvia whose exchange rate was fixed against the SDR. Among the other countries, the Czech Republic adopted intermediate exchange rate regime until 1997 and shifted to a floating regime thereafter. Croatia's exchange rate regime is also categorized as an intermediate regime. Ukraine's rate was pegged to the US dollar, but became more free floating since 1999. Due to the regular interventions however, the regime is classified as a conventional pegged arrangement from 2001. There are countries that have not changed

⁸ We remove significant outliers for Russia for 1998M9, for Turkey 2000M12 and 2001M2 and 1998M8 for the Ukraine.

⁹ The IMF provides detailed classification of exchange rate regimes on its homepage (<http://www.imf.org>).

their exchange rate regime during our sample period, which include Russia, Romania, Slovenia and Turkey that employ floating exchange rate regimes. Thus, a different exchange rate arrangement may reflect heterogeneity in the risk premium fluctuation.

We utilize monthly data on market interest rates and the bilateral exchange rates (defined as domestic currency units per unit of US Dollar) (AA..ZF).¹⁰ Since the market for long-term government securities is typically illiquid in emerging markets, interest rates are short-term money market rates (60..ZF).¹¹ Here, interest rates and exchange rate changes are expressed in annual percentages. Based on equation (5), we derive the risk premium ($risk_{it}$) in a panel data set for country i and time t as follows.

$$risk_{it} = (s_{it+1} - s_{it}) - i_{it} + i_{it}^*, \quad i=1, \dots, 11, t=1998M1, \dots, 2008M2 \quad (6)$$

The asterisk indicates the benchmark country which is the US in this study. As mentioned in the previous section, our definition of $risk_{it}$ includes the expectations error of investors, and the long-run UIRP requires $risk_{it}$ to be stationary.

<<Table 1>>

Table 1 provides the basic statistical summary of our foreign exchange risk premiums and their components; exchange rate changes ($s_{it+1} - s_{it}$) and interest rate differentials ($i_{it} - i_{it}^*$). This table indicates that the mean and standard deviation of the bilateral dollar exchange rates differ according to the country. More than half of our countries have negative average exchange rate growth, hence they experienced currency appreciation (revaluation) during our sample period. We note that the level of currency appreciation

¹⁰ We obtained data based on SDRs and converted this into US Dollars.

¹¹ Estonia has missing data in 1999M3 and 1999M10 to M12 which are created by linear interpolation.

(revaluation) differs substantially among countries, with a range of one to eight percent per year. In contrast, four countries (Romania, Russia, Turkey, and Ukraine) experienced currency depreciation (devaluation) of an average magnitude of more than nine percent per year. Those four countries also experienced higher exchange rate volatility, measured by a standard deviation.

<<Figures 1, 2 and 3>>

Interest rate differentials are positive in eight out of the 11 countries. According to our definition of interest rate differentials ($i_{it} - i_{it}^*$), their typically positive value suggests relatively higher interest rates in emerging markets compared with the US rate. The interest rates of Romania, Turkey, and Ukraine were considerably higher than the benchmark rate. Those countries also experienced high volatility in interest rates, which reflects their economic and financial difficulties during our sample period. Interestingly, countries with high exchange rate volatility also experienced high interest rate volatility. The interest rate spread and US interest rate are plotted in Figures 1 and 2. Clearly there is a high degree of commonality in interest rate spreads and this would appear, at least graphically, to be related to the US rate.

Table 1 also shows that the risk premium is negative on average in most countries. Appreciating currencies and high interest rates in emerging markets attribute to this negative risk premium. Not surprisingly, countries experienced high volatility in exchange rates and interest rates exhibited high volatility in the risk premium, and seem to be ones that are not EU members and/or have implemented a flexible exchange regime.

A preliminary analysis is also conducted to examine if the UIRP holds for a pooled regression (see Figure 3). Essentially, this sets out a scatter plot illustrating the bivariate relationship between the annualized percentage change in the exchange rate (Δs_{it+1}) and interest rate differential ($i_{it} - i_{it}^*$). This indicates that the interest rate differential is positively related to exchange rate changes, in contrast to the UIRP, and unfortunately the estimated coefficient is significantly different from one, suggesting a possible violation of the UIRP.¹²

4. The Decomposition of the Exchange Risk Premium

Previous studies rarely attempted to decompose the risk premium into common and idiosyncratic components although their research target was industrial countries which can be characterized as open economies. If there is a degree of cross correlation in any panel and one is interested in the time-series properties of this data, then it is sensible to consider their stationary properties.

For this purpose, we employ the PANIC approach (Bai and Ng, 2004). This method utilizes a factor structure to model the nature of the nonstationarity in large dimensional panels. This is set out for the case where only an intercept is included:

$$risk_{it} = c_i + \Lambda_i' F_t + e_{it}. \quad (7)$$

¹² The pooled estimated regression suggested the estimated coefficient on interest rate differentials in the UIRP regression was 0.58(t-statistic=10.71), hence this is significantly different from one, and represents a failure of UIRP although not to the extent suggested in the survey by Froot and Thaler (1990). A preliminary analysis indicates a substantial proportion of the bivariate correlations using Ng (2006) indicated that these were insignificantly different from zero. Furthermore, a comparison of the standard deviation of the idiosyncratic to the data indicated that a substantial proportion of the variability of Romania, Russia, Turkey and the Ukraine was explained by the idiosyncratic component. This suggests that a common factor is more likely to exist for $N = 7$ which is also consistent with an increase in the eigenvalue of this smaller panel of countries.

The series $risk_{it}$ is a sum of a cross-section specific constant (c_i), a common component $\Lambda_i'F_t$, where Λ_i is a corresponding matrix of factor loadings and F_t are the factors, and an error, e_{it} , which is the idiosyncratic component. The panel time series $risk_{it}$ is nonstationary if the common factors or the idiosyncratic component, or both, are nonstationary. In this connection, the PANIC allows us to identify whether nonstationarity is pervasive or series-specific. Bai and Ng (2004) propose the method of principle components to obtain the common factors, and the appropriate number of factors is determined by the information criteria developed by Bai and Ng (2002). The PANIC does not assume that only the idiosyncratic component may be nonstationary, unlike Moon and Perron (2004) and Pesaran (2006). The PANIC determines explicitly whether the nonstationarity in a panel time series is pervasive or variable-specific.

We make use of two test statistics from Bai and Ng (2004): an Augmented Dickey Fuller (ADF) test on the common factor ($ADF_{\bar{F}}^c$) and a Fisher-type pooled ADF test on the idiosyncratic individual errors ($ADF_{\bar{\varepsilon}}^c(i)$). The test statistic on the idiosyncratic element is distributed as standard normal as follows:

$$P_{\bar{\varepsilon}}^c = \frac{-2 \sum_{i=1}^N \log p(i) - 2N}{\sqrt{4N}} \rightarrow N(0,1). \quad (8)$$

where $p(i)$ is the p -value associated with ($ADF_{\bar{\varepsilon}}^c(i)$) of the ADF test for the i cross section, and ρ_i is the autoregressive parameter of the independent error processes. The test statistic examines whether $H_0: \rho_i = 1 \quad \forall i$ against $H_0: \rho_i < 1$ for some i . Thus, under the null hypothesis, all cross-sections are nonstationary and the alternative is that some may be stationary.

The stationarity of the common factors is individually examined using the ADF test. With one common factor, this test becomes identical to the original ADF test. Thus, this test is based on the following specification with the null of $\theta = 0$ against the alternative of $\theta < 0$.

$$\Delta \hat{F}_t = \alpha + \theta \hat{F}_{t-1} + \delta_i \sum_{i=1}^p \hat{F}_{t-i} + \varepsilon_t \quad (9)$$

where \hat{F}_t is the estimate of common factors. The PANIC results are summarized in Table 2. In order to see the robustness of our findings, we conduct a panel data analysis for the full-sample ($N=11$) and sub-sample of countries ($N=7$). First, this table shows three information criteria (*IC1* to *IC3*) to determine the number of common factors in our data.¹³ These information criteria produce somewhat mixed results. While two criteria (*IC1* and *IC2*) suggest five common factors, one criterion (*IC3*) raises evidence of one common factor. The subsequent part of our analysis is based on one factor since although the first two criteria may be more reliable in a large data set (i.e., large N and T), they tend to overestimate the number of true common factors in a finite sample context (Bai and Ng 2002).

<<Table 2>>

Table 2 shows the common factor is nonstationary (i.e. $ADF_{\hat{F}}^c < -2.86$) and the

¹³ There three information criteria are from Bai and Ng (2002). Where $\hat{\sigma}^2(k)$ is based on the residuals from a regression of the first differenced data on k principal components, the first information criterion can be expressed as $IC_1(k) = \ln \hat{\sigma}^2(k) + k \ln(NT/(N+T))(N+T/NT)$. The second information criterion is $IC_2(k) = \ln \hat{\sigma}^2(k) + k \ln(N+T/NT) \ln[\min\{\sqrt{N}, \sqrt{T}\}]^2$. Bai and Ng (2002) however suggest that a third information criterion is to be preferred with panel cross sectional correlation: $IC_3(k) = \ln \hat{\sigma}^2(k) + k \hat{\sigma}^2(N-T-k) \ln(NT)/NT$.

idiosyncratic is stationary (i.e. $P_e^e > 1.64$). This result remains valid even when the abovementioned four countries (Romania, Russia, Turkey and Ukraine) are excluded from the eleven country analysis. This also implies that the exchange risk premium is nonstationary for these countries, and indicates the UIRP does not hold even in the long-run. The poor performance of the UIRP here may result partly from the use of short maturity returns.

In order to check our conclusion as regards the nonstationary common factor, we examine the stationarity of the common factors based on 5 factors which *IC1* and *IC2* suggest. The five common factors are $[-2.505, -3.728^*, -3.731^*, -3.023^*, -2.542]$ where the asterisk indicates that the null of the ADF test can be rejected at the 5 percent level. Next, in the presence of multiple nonstationary common factors, the multivariate Johansen cointegration test is conducted using the first and fifth common factors. The results are reported in Table 3 with evidence that they are not cointegrated. In other words, even though the analysis is conducted under an assumption of five common factors, the conclusion of nonstationary common factors is valid.

<<Table 3>>

This result is consistent with the univariate analysis of each country's risk premium (Table 4). This is conducted by the standard unit root tests (ADF and ADF-GLS (Elliott et al. 1996) tests), and the results in Table 4 indicate that there is some stationarity within the panel data set, but the PANIC test is robust to this and highlights pervasive nonstationary although countries will be affected differently. Indeed, these two unit root tests suggest that four out of eleven countries (Croatia, Lithuania, Russia, and Turkey)

have a stationary risk premium.¹⁴

Furthermore, we examine in Table 4 the relative importance of the common factor using two statistical ratios: firstly, the ratio of the standard deviation of the idiosyncratic residual in equation (7) to the differenced risk data (i.e. $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$) and secondly the ratio of the standard deviation of the common to the idiosyncratic component (i.e. $\sigma(\Lambda_i'F_t)/\sigma(e_{it})$). The former should be equal to one and the latter equal to zero if the idiosyncratic dominates. According to $\sigma(\Lambda_i'F_t)/\sigma(e_{it})$, the common component of the risk premium is relatively high in Bulgaria, Estonia, and Slovenia. In contrast, the idiosyncratic dominates the variation of the risk premium in Ukraine. This is confirmed by the ratio of the standard deviation of the idiosyncratic component to the differenced data (i.e. $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$). A high $\sigma(\Lambda_i'F_t)/\sigma(e_{it})$ ratio may be an indication of a nonstationary premium, but this is not the case. This finding motivated us to re-examine the premium by reducing the number of countries.

<<Table 4>>

The seven country analysis yields a very similar outcome (Table 5), but we can see more clear characteristics of the risk premium. The four countries dropped are Romania, Russia, Turkey and Ukraine that experienced significantly high exchange rate depreciation and interest rate spreads. Interestingly, those countries (e.g., Estonia and Slovenia) with the smallest $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$ in Table 5 are also the ones with nonstationary risk premiums. This is consistent with a high $\sigma(\Lambda_i'F_t)/\sigma(e_{it})$ ratio and our finding of a nonstationary common factor. Thus, the risk premium of Estonia and Slovenia are dominated by the common factor, which suggests the sensitivity of these

¹⁴ Here we suggest the time-series is stationary if both tests reject the null hypothesis of the unit root.

countries to economic and financial changes in foreign countries. By contrast, of the 7 countries, Lithuania with the highest $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$ and the lowest $\sigma(\Lambda_i'F_t)/\sigma(e_{it})$ appears to have a significant country-specific element in her risk premium and shocks are temporary to her risk premium. This may be an indication that large economies tend to have more influence on their own risk premium. Also the panel will present different evidence of nonstationarity and there appears to be pervasive nonstationarity in the system, as is indicated by the nonstationary factor.

<<Table 5>>

In terms of associating the pervasive risk premium with potential determinants, we estimate the long-run relationship between the common factor and both US inflation and industrial production using Phillips and Hansen (1990) Fully-Modified-OLS. We chose these explanatory variables based on a Taylor type monetary policy rule and our data frequency (i.e., monthly data).¹⁵ Additionally, Uribe and Yue (2006) demonstrate that the US interest rate explains about 20 percent of variation in economic activity in emerging markets. Our results are reported in Table 6 and confirm that US inflation is consistently significant at the 10 percent significance level in explaining the common risk factor and occasionally at the one percent level. We identify a negative long-run relationship and increases in inflation are associated with a more than proportionate changes in risk.¹⁶ This is consistent with US monetary policy explaining a degree of the risk premium for emerging market economies. In contrast to the US inflation, industrial production is found to be insignificant in our data. Overall this should reflect the

¹⁵ Stock and Watson (2007) suggest that US inflation can be modelled as a nonstationary process. FM-OLS is therefore an appropriate means of examining the long run relationship between a nonstationary common exchange rate risk premium and inflation.

¹⁶ There is unlikely to be any endogeneity between emerging economies risk and US monetary policy responses. In any case FM-OLS is corrected for endogeneity and serial correlation.

inflation preoccupation of the US monetary authorities during our sample period.

<<Table 6>>

5. Conclusion

This paper empirically analyzed the foreign exchange risk premium of eleven emerging European countries. Unlike most previous research, the time-varying premium here is decomposed into common and idiosyncratic factors using the statistical method developed by Bai and Ng (2004). This is a rather different approach to investigating the foreign exchange risk premium because most previous research does not consider any cross-sectional element in the premium. Furthermore, the risk premium of emerging countries in Europe has not been much studied although exchange rate risk is one important concern for most countries considering adopting the euro.

Our results suggest that unlike industrial countries, the foreign exchange premium in emerging markets is not stationary. We reveal that the nonstationary common factor is attributable to the nonstationary risk premium, and the idiosyncratic component of the premium is found to be stationary. This gives rise to evidence of a violation of the UIRP even in the long-run context, and thus it follows that the UIRP should not be viewed as an equilibrium concept. In short, compared with previous studies (i.e. Bansal and Dahlquist 2000 and Frankel and Poonalwala 2006), this paper provides weaker evidence of the UIRP and underscores the importance of the deviation from the UIRP in emerging markets.

Furthermore, we analyzed what economic factors can explain the nonstationary common risk premium for these countries. Our results show that this common factor in

emerging markets reflects economic and financial developments in the US which is our benchmark country. In particular, US inflation has some explanatory power over the common movement in the premiums. This confirms that emerging markets are heavily sensitive to the US monetary policy, consistent with Uribe and Yue (2006) and Neumeyer and Perri (2006).

Finally, this paper assumes a linear characteristic for the foreign risk premium. Although technically it is more complicated, one may wish to consider a possible non-linear characteristic of the foreign risk premium in the panel data context. In a univariate analysis, Mehl and Cappiello (2007) suggest that such non-linearities in the premium are more common in emerging markets than in developed countries.

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TABLES

Table 1: Summary Statistics for the Risk Premium, Exchange Rate and Interest Differential

<i>Country</i>	<i>Risk Premium</i>		<i>Exchange Rate Change</i>		<i>Interest Rate Differential</i>	
	<i>Mean</i>	<i>Standard Deviation</i>	<i>Mean</i>	<i>Standard Deviation</i>	<i>Mean</i>	<i>Standard Deviation</i>
Bulgaria	-2.736	32.471	-3.740	32.107	-1.004	1.807
Croatia	-4.646	32.535	-3.296	32.771	1.351	3.121
Czech Republic	-8.634	41.425	-7.678	41.208	0.955	3.185
Estonia	-4.823	32.033	-3.860	31.825	0.963	2.693
Latvia	-2.771	19.557	-2.808	19.311	-0.036	2.044
Lithuania	-5.537	23.583	-5.929	23.480	-0.393	1.592
Romania	-19.513	34.640	10.325	37.380	29.838	29.498
Russia	6.445	80.438	13.393	84.815	6.948	16.850
Slovenia	-2.891	32.221	-1.256	32.145	1.635	2.004
Turkey	-23.099	64.773	17.785	62.272	40.885	41.843
Ukraine	-1.605	47.068	9.464	49.161	11.069	17.816

Notes: The foreign exchange premium is determined by $1200*(s_{it+1} - s_{it}) - i_{it} + i_{it}^*$. Where s_{it} is defined as domestic currency units for country i per US dollar, i_{it} is the domestic interest rate and i_{it}^* is the US interest rate. The interest rate differential is defined as $i_{it} - i_{it}^*$. We use money market interest rates. Sample period 1998M1 to 2008M2.

Table 2: PANIC Evidence for the Risk Premium: US as Benchmark

	Idiosyncratic $P_{\hat{\epsilon}}^c$	Common factor $ADF_{\hat{F}}^c$	<i>IC1</i>	<i>IC2</i>	<i>IC3</i>
US\$ as Benchmark					
11 countries	3.800*	-2.691 (54%)	5	5	1
7 countries	2.361*	-2.505 (85%)	5	5	1

Notes: Asterisk (*) denotes rejection of the null hypothesis of unit root. In our factor model, $ADF_{\hat{F}}^c$, the factor unit root test, has a 5% asymptotic critical value of -2.86 (see Bai and Ng, p. 1135, 2004). The idiosyncratic unit root test, $P_{\hat{\epsilon}}^c$ is distributed as standard normal, hence the critical value at the 5% level is 1.64. Lag lengths are determined by the formula $4[\min[N,T]/100]^{1/4}$ following Bai and Ng (2004). The data set covers ten countries from 1998M1 to 2008M2. The maximum number of the common factors is equal to five. Four countries, Romania, Russia, Turkey, and Ukraine, are excluded for the seven-country analysis. Eigenvalues in parentheses give an impression of the degree to which the different factors explain overall variation in the panel time series.

Table 3: Cointegration between Common Factors For 7 Countries

$H_0: \text{rank} \leq$	Eigen value	Trace test	P-value
0	0.100	14.448	0.266
1	0.027	3.006	0.588

Notes: The Johansen cointegration test is conducted for 1st and 5th common components which are found to be nonstationary for the group of 7 countries. The constant enters the cointegrating vector.

Table 4: PANIC Evidence for the Risk Premium for 11 Countries

	Unit Root Tests		Factor Importance	
	ADF-GLS	ADF	$\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$	$\sigma(\Lambda_i' F_t)/\sigma(e_{it})$
Bulgaria	-1.997*	-2.381	0.120	3.091
Croatia	-2.476*	-2.941*	0.227	2.219
Czech Republic	-1.535	-3.458*	0.181	2.400
Estonia	-1.646	-2.629	0.097	3.250
Latvia	-1.711	-1.964	0.311	1.545
Lithuania	-2.985*	-2.982*	0.452	1.139
Romania	-0.358	-3.973*	0.769	0.527
Russia	-6.394*	-8.401*	0.969	0.142
Slovenia	-1.984*	-2.460	0.086	3.511
Turkey	-2.148*	-3.948*	0.978	0.141
Ukraine	-1.230	-3.053*	1.000	0.008

Notes: This table is based on PANIC results from Table 3. 5% asymptotic critical value for ADF-GLS is -1.98 from Ng and Perron (2001), in bold and with asterisk when significant. ADF test has a 5% asymptotic critical value of -2.86 from Fuller (1976), in bold and with asterisk when significant. The $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$ is the ratio of the standard deviation of the idiosyncratic to the differenced risk data and $\sigma(\Lambda_i' F_t)/\sigma(e_{it})$ is the ratio of the standard deviation of the common to the idiosyncratic component. The former should be equal to one and the latter equal to zero if the idiosyncratic dominates.

Table 5: PANIC Evidence for the Risk Premium for 7 Countries

	Unit Root Tests		Factor Importance	
	ADF-GLS	ADF	$\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$	$\sigma(\Lambda_i' F_t)/\sigma(e_{it})$
Bulgaria	-1.997*	-2.381	0.110	3.310
Croatia	-2.476*	-2.941*	0.205	2.377
Czech Republic	-1.535	-3.458*	0.172	2.430
Estonia	-1.646	-2.629	0.095	3.405
Latvia	-1.711	-1.964	0.320	1.539
Lithuania	-2.985*	-2.982*	0.468	1.109
Slovenia	-1.984*	-2.460	0.075	3.889

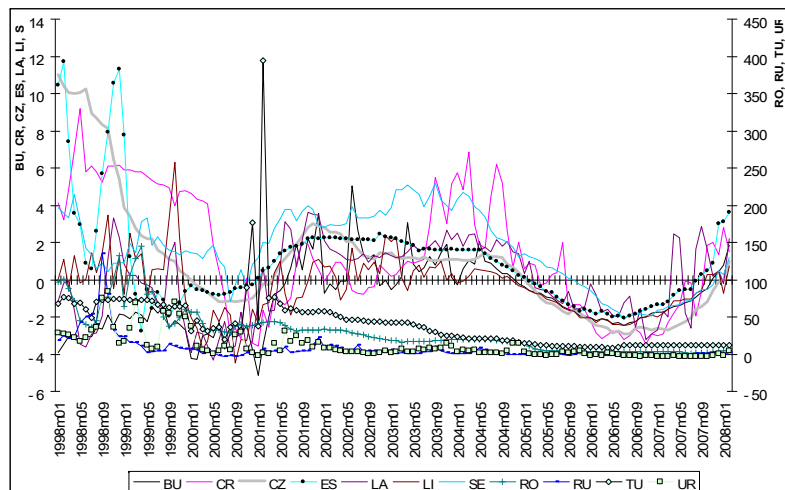
Notes: This table is based on PANIC results from Table 3. 5% asymptotic critical value for ADF-GLS is -1.98 from Ng and Perron (2001), in bold and with asterisk when significant. ADF test has a 5% asymptotic critical value of -2.86 from Fuller (1976), in bold and with asterisk when significant. The $\sigma(\Delta e_{it})/\sigma(\Delta risk_{it})$ is the ratio of the standard deviation of the idiosyncratic to the differenced risk data and $\sigma(\Lambda_i' F_t)/\sigma(e_{it})$ is the ratio of the standard deviation of the common to the idiosyncratic component. The former should be equal to one and the latter equal to zero if the idiosyncratic dominates.

Table 6: Determinant of the Common Risk Factor

	OLS	FM-OLS	FM-OLS	FM-OLS
Constant	-4.788 (t=4.788) ^a	-3.974 (1.122)	-4.707 (1.382)	
US INF	-1.134 (1.700) ^c	-1.344 (1.876) ^c	-1.326 (1.861) ^c	-2.288 (3.730) ^a
US IP	-3.486 (0.681)	-4.787 (0.870)		

Notes: Sample period 1998M3 to 2008M2. We use OLS and Phillips and Hansen (1990) FM-OLS estimation, in which the dependent variable is the common factor from the risk premium (see Table 3). Explanatory variables are US Inflation (US INF) and US industrial production (US IP). Parentheses contain t-statistics: a, b and c indicate estimated coefficients statistically significant at the 1%, 5% and 10% level respectively.

Figure 1. Interest Rate Differentials



Notes: Bulgaria (BU), Croatia (CR), Czech Republic (CZ), Estonia (ES), Latvia (LA), Lithuania (LI), Romania (RO), Russia (RU), Slovenia (SE), Turkey (TU), and Ukraine (UR).

Figure 2. US Interest Rates

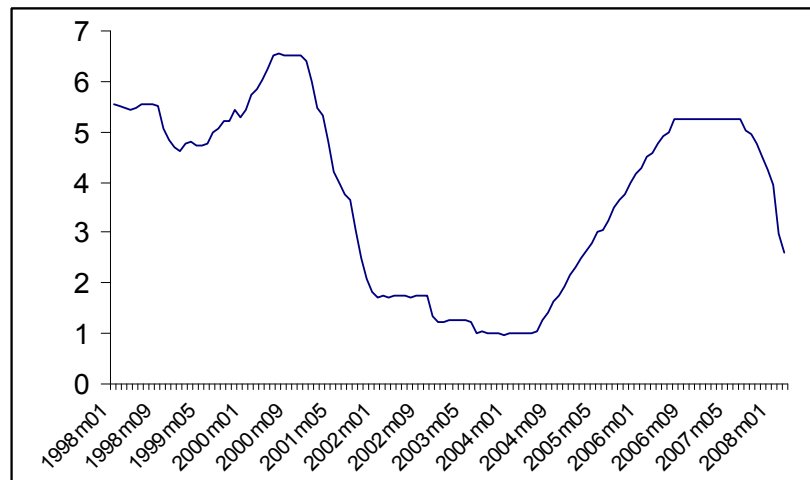


Figure 3. Exchange Rate Change and Interest Rate Differential Relation

