THE ACADEMY OF ECONOMIC STUDIES BUCHAREST
DOFIN - DOCTORAL SCHOOL OF FINANCE AND BANKING

DISSERTATION PAPER

DETERMINANTS OF SPREADS OF ROMANIAN SOVEREIGN BONDS

- an application on the EMBIG spreads –

Student: BERBECARU CLAUDIA-FLORIANA
Supervisor: Professor MOISĂ ALTĂR

Bucharest, July 2008

Abstract

There was a rapid compression in the spreads of Romanian sovereign bonds in last years, to a record low level reached in the summer of 2007. We show that the developments in the domestic fundamentals and in the risk appetite of foreign investors on the international markets explain the developments in the spreads. Using data for EMBIG spreads for Romania and other ten Emerging Economies, we find a long-run relationship between the spreads on the one hand and a Credit Rating Outlook Index (CROI) and the volatility index VIX on the other hand. The CROI is a proxy for the developments in the domestic fundamentals, while the VIX is a proxy for the risk appetite of the international investors. To estimate the long-run relationship, we use both a pool equation with fixed effects and the pooled mean group (PMG) estimator of Pesaran, Shin, and Smith (1997). There is a large similitude between the deviations of spreads from the level implied by the long-run relationship in the case of Bulgaria and Romania, which we explain by the EU accession process of these two countries. We find also a comovement in the volatility of daily returns of CEE sovereign bonds, with spillover effects especially between Bulgaria and Romania. The domestic fundamentals were the main drivers of the cumulated change in the equilibrium level of spreads for Romanian sovereign bonds between May 2002 and April 2008.

Contents

I. Introduction

The spreads between the yields of sovereign bonds issued by Emerging Economies and the yields of bonds with the same characteristics but issued by a developed benchmark economy (which virtually is free of default risk) are commonly perceived as reflecting market perceptions of the risks of default of these less developed economies. The yields spreads measure the premium required by investors to hold such securities and they are a component of the costs these less developed countries should pay when borrowing on the external markets.

There was a rapid decrease in the spreads of sovereign bonds for emerging countries in last years. For most of the Emerging Economies, the spreads reached a record low level in the 2007 summer, slightly before the US subprime crisis hit the international financial markets. For instance, EMBIG spreads for Romania decreased from 355 bp in May 2002 to only 26 bp in May 2007. Also, the spreads for emerging markets measured by the EMBIG Composite Index decreased from 370 bp in May 2002 to 53 bp in May 2007.

Clearly, the compression in the spreads has came hand in hand with an improvement in the "real" domestic fundamentals (e.g. decrease in the inflation rate, high GDP growth rates, lower external imbalances) for most of the emerging economies. The improvement in the sovereign ratings of international rating agencies for these countries could be considered as reflecting the progresses recorded by these countries. For instance, the S&P rating for Romania long term foreign currency debt improved from B+ Positive in May 2002 to BBB- Stable in April 2007. And similar improvements were recorded in the case of most of the emerging countries. At the same time, the accession to the European Union was an important driver of the structural reforms of the economic progresses recorded in Emerging Countries from Europe (Poland, Hungary, Slovakia, Czech Republic, Romania, Bulgaria).

But the compression in spreads was due not only to domestic fundamentals, but also to external factors. Starting 2002, the risk appetite of investors on the international markets increased rapidly. For instance, the volatility index VIX which is thought to be a good measure of investors' risk appetite on the international markets was on a downward trend This developments were supported by the abundance of the liquidity in the markets as interest rates in major industrialized countries (US, Euro Area, Japan) were at historically low levels. The recent crisis which hit the worldwide financial markets in the summer of 2007 revealed that investors generally under-evaluated the price of the risk. EMIBIG spreads for many of the Emerging Markets increased when the crisis amplified.

Figure 1. EMBIG Romania and EMBIG Composite

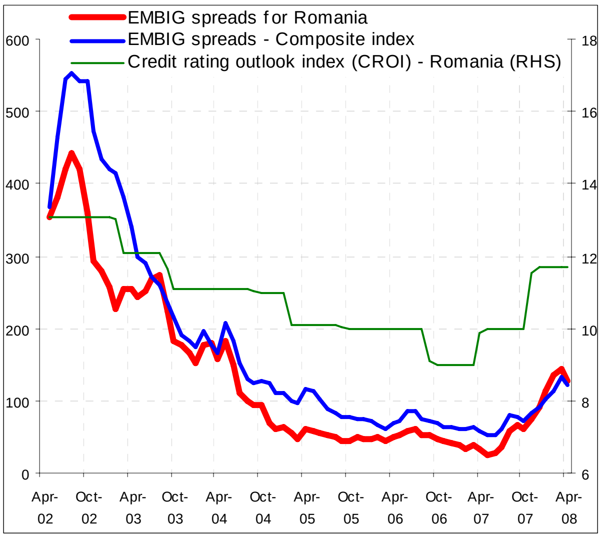
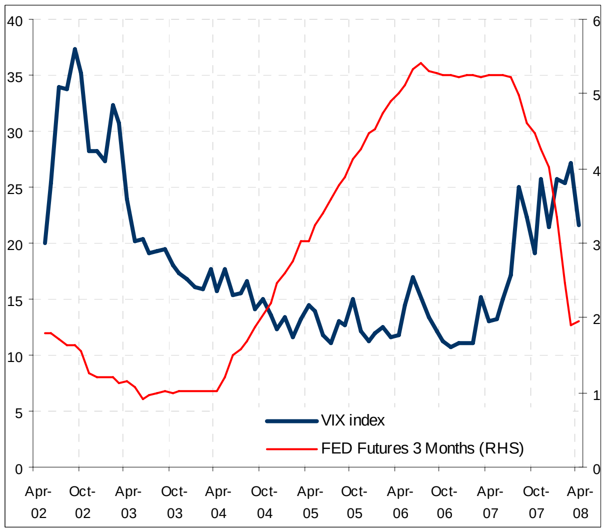


Figure 2. Volatility index VIX and key interest rate in United States



The dramatic spreads decrease in last years has renewed attention in this subject, with empirical analysis trying to explain the factors behind this evolution. The empirical studies try identify how important the contribution of domestic fundamentals was and how much the spreads were driven by external factors. This is also the subject of our paper. We will identify the contribution of the domestic and external factors for the dynamics of the spreads of the Romanian sovereign bonds. The analysis is performed in a multivariate framework, taking into account the developments in the spreads of other ten emerging countries. We look also for a common pattern in the volatility of the returns for sovereign bonds issued by countries from Europe and for spillover effects between these volatilities.

The paper is organized as follow. Section II presents a short review of the literature on the determinants of the spreads for sovereign bonds. Section III presents the framework usually used to conduct analysis regarding the determinants of spreads for sovereign bonds. Section IV includes an empirical analysis for Romanian sovereign bonds. At the beginning (IV.A) the data used in the analysis are presented. In the second part (IV.B) two panel data estimation methods are used to estimate the equation for dynamics of the spreads for Romanian sovereign bonds. A long run relationship between the spreads of Romanian sovereign bonds on the one hand and the Credit Rating Outlook Index (computed on the base of the sovereign ratings of Standard and Poor's) and the volatility index VIX on the other hand is firstly estimated using a pool model with fixed effects and data for 11 countries. The same long run relationship is then estimated using the pool mean group (PMG) estimator introduced by Pesaran, Shin, and Smith (1997). Last part of the section IV (IV.C) is focusing on the existence of common pattern in the daily volatility of the subset of the Central and Eastern European sovereign bonds. Long-run components for the volatility of daily return for sovereign bonds of five European countries are estimated using Component-GARCH models. We test also for spillover effects between these components. The final section of the paper (V) concludes and presents some directions to be followed in order to improve the current analysis. The results of the estimation are included in the Appendix.

II. Literature review

Financial markets have become more and more globalized in last years. The globalization of financial markets is a part of a wider phenomenon of globalization of the national economies. Increase in the international trade in goods and a service was one reason for the increase in the financial flows between countries. Liberalizations of capital accounts in less developed countries was also a factor which boosted financial flows towards these countries, given that these countries usually need important financial resources in order to sustain the real convergence process. In fact, one of the most important benefits of the financial globalization is that globally integrated financial markets provide more flexible ways of both financing current account deficits and recycling current account surpluses. Moreover, the free play of market mechanisms should tend to ensure that both borrowers and lenders do not knowingly take excessive risks (Obsfeld 1994[1]). At the same time, the entry of foreign financial institutions into domestic financial markets can bring sizeable benefits, as increased competition can help to enhance efficiency in the financial sector. Emerging economies in Latin America, Asia, Africa and Europe have been an important recipient of funds provided by developed countries. Both foreign direct investments and portfolio investments in these countries increased rapidly. Also, the volume of bonds issued by developing countries has risen significantly since 1990, especially in the case of countries from Latin America.

Yields on bonds issued by the Emerging Economies on the external markets are important as they reflect the cost these countries should pay to finance their economic development and they capture the default risk of these countries. A key question is whether the borrowing cost for a country can be associated with its domestic economic fundamentals

[1] Obstfeld, M. (1994), *"International capital mobility in the 1990s"*, CEPR discussion paper no. 902

or that there are other factors which might be also important. Another question is weather there is or not a co-movement in the yields of different countries with different domestic economic fundamentals. If the borrowing cost is driven mainly by domestic fundamentals, then countries implementing sound macroeconomic policies should benefit from better financing conditions. On the other hand, if the borrowing costs are driven by external factors, then there are risks that developing countries would be vulnerable to shocks located in the developed economies.

Empirical studies focusing on the determinants of borrowing costs encountered by emerging countries on the external markets usually are using in the analysis the Emerging Markets Bond Index Global (EMBIG) spreads computed by JP Morgan. The spreads reflect the difference between the yields of emerging country's sovereign bonds and yields of bonds with identical maturity and issued by the US government (in the case of EMBI spreads for bonds denominated in USD) or by German government (in the case of EMBI spreads for bonds denominated in euro).

Many authors choose to use EMBI spreads (secondary market yields) in the analysis and not primary yields because the latter ones may lead to sample selection biases. Eichengreen and Mody (1998a)[2] noted that when secondary spreads rise due to poor market conditions, primary yields do not rise proportionally, and in some cases they even fall. In some circumstances the perceived risk of emerging market debt may deteriorate leading to raising secondary market spreads. However, this may have an opposite effect on launch spreads because the factors that increased the perceived risk of emerging markets may ration out of the market the riskier investors, leaving only low risk borrowers to launch new issues. Using primary yields as a measure of risk Eichengreen and Mody (1998b)[3] find that changes in macroeconomic fundamentals explain only a fraction of spread evolution.

The empirical studies reveal that both the variables measuring policies and economic performance (fundamentals) of a country and the external variables (like international interest rates, global liquidity conditions and the risk appetite on the international markets) are drivers of spreads for sovereign bonds. In many cases, the external factors have almost the same importance than the domestic ones and in some periods they are becoming the main driver of spreads (Gonzalez-Rozada and Levy-Yeyati (2006), Hartelius, Kashiwase and Kodres (2008)). Hauner and others (2007) show that the Emerging Economies from Europe enjoyed

---

[2] Eichengreen, B and A. Mody (1998a), "Interest rates in the north and capital flows to the south: is there a missing link?", NBER Working Paper, No. 6408

[3] Eichengreen, B and A. Mody (1998b), "What explains changing spread on EM debt: fundamentals or market sentiment?", NBER Working Paper, No. 6408

higher policy credibility than other Emerging Economies due to the accession process to the European Union and they have also lower spreads. Luengnaruemitchai and Schadler (2007) also suggest the existence of a EU “halo effect” for these countries.

As regards the most widely used techniques in literature for analyzing sovereign bond spreads determinants we can mention the conventional panel estimation techniques, the panel mean group estimation procedure proposed by Pesaran, Shin and Smith (1999), Vector Autoregressive Models (VAR).

Among the papers which employ conventional panel estimation techniques we can mention those of Hartelius, Kashiwase and Kodres (2008) and Luengnaruemitchai and Schadler (2007). Hartelius, Kashiwase and Kodres (2008) model the EMBI spreads as a function of two important factors: fundamentals and liquidity. In comparison with other papers which use different macroeconomic variables (Goldman Sachs (2000)), in their paper the above mentioned authors use as a proxy for macroeconomic variables a constructed credit rating outlook index which takes into account the non-linear relation which exists between spreads and rating. Luengnaruemitchai and Schadler (2007) model EMBI spreads in a similar way with Hartelius, Kashiwase and Kodres (2008). They also analyse spreads determinants, but they try to find some aspects which distinguish the new members of EU from another emerging markets. Eventually they reach the conclusion that the apparent advantage of these countries can not be explained by quantifiable factors but rather reduced risk aversion due to the new EU membership.

One of the papers which study the determinants of sovereign spreads using secondary market yields and the panel mean group estimation technique is that of Goldman Sachs (2002). They estimate a long run equilibrium model of emerging market spreads using the pool mean technique developed by Pesaran, Shin and Smith (1999). This technique involves defining a dynamic, error correction panel where short run parameters are allowed to vary by cross sections while long run elasticities are restricted to be identical across groups. Panel mean group estimator is also used by Ferrucci (2003) for investigating the relationship between emerging market spreads and a set of common macroeconomic variables. He concludes that market do take into account the macro fundamentals when pricing sovereign risk but non fundamental factors also play an important role. He compares market based spreads with model based ones and finds that spreads trade at a level which is close to the theoretical equilibrium level explained by fundamentals. He concludes assuming that the misalignments may de due to capital market imperfections or higher investor risk appetite.

There are also some empirical studies in which the spreads are treated as being endogenous variables, affecting and being affected by domestic and international macroeconomic conditions (Uribe and Yue (2003)[4]). In these case, they uses Vector Autoregressive Models (VAR).

III. Determinants of the spreads for sovereign bonds

The typical assumption is that the spreads of sovereign bonds yields for emerging countries against the yield of a developed reference country reflect the default risk of the country. Accordingly, the conventional approach should be to model the spreads of sovereign yields as a function of the probability of default and of the loss given the default (ot the expected recovery). Models can be classified in two categories: structural and reduced form. Most of the empirical analyses of the spreads of sovereign bonds are using reduced-form models.

From an analytical point of view, in a simpler form, the relationship between the yields r of domestic bonds (which have a default risk) and the yields of foreign risk-free government debt yields r_f in the presence of risk-averse international investors can be written as:

$$(1-p)(1+r)+pRV=(1+r_f) \tag{1}$$

where p is the expected probability of default, RV is the recovery value. Assuming that the probability of default has a logistic form, Edwards (1994) obtained a simple log-linear relationship between the spreads of sovereign bonds and their potential determinants:

$$\log \mathrm{s}_\mathrm{t} = \alpha_0 + \sum_{i=1}^{k} \alpha_i x_{i,t} + \varepsilon_t \tag{2}$$

where $s = r - r_f$ is the sovereign bond spreads and x_i $i = 1,k$ is a set of macroeconomic fundamentals which the probability of default of the country depends on and ε is an error term.

The set of macroeconomic variables used in the empirical studies as determinants of the sovereign bonds spreads refer mainly to liquidity and solvability indicators which reflect the sustainability of the existing debt stock (both domestic and external). The country must

[4] Uribe, Martin, and Zhanwei Vivian Yue (2003), Country Spreads and Emerging Countries: Who Drives Whom?”,

be able to generate enough foreign exchange resources in order to service its external obligation. The assessment of debt sustainability takes into account indicators as: the economic growth rate, the inflation rate, the public budget balance (as percent of GDP), the external debt (as percent of GDP), the current account balance (as percent of GDP), the official foreign exchange reserves (as percent of GDP or in months of exports), the real exchange rate (as a measure of external competitiveness of the country, the degree of openness of the economy. At the same time, political factors might be an important factor of sovereign spreads although they are more difficult to quantify.

However, the empirical studies revealed that not only the domestic fundamentals are important in explaining the spreads of sovereign bonds. Alongside domestic fundamentals, external factors are also very important. The spreads of sovereign bonds of emerging markets captures the risk premia attached to particular countries, but they reflect not only the default risk of the country but also the degree of unwillingness to buy that country's debt. This might be of particular interest because the unwillingness of foreign investors to buy bonds issued by an emerging country may be unrelated to the actual default risk, but instead it might reflect factors such as the financial position of investors, liquidity risk in financial markets, or other factors which are related to the investor's risk appetite. In these case, the relation (1) might be augmented with a risk premium φ which depends on the foreign degree of risk aversion and possible on the probability of default p.

$$(1-p)(1+r)+pRV=(1+r_f)+\phi \tag{3}$$

The external factors might become an important driver of the sovereign spreads during period of stress: "*When U.S. stocks are volatile, EMBI spreads widen. They narrow again when U.S. calm down. That suggests that emerging market debt is not being driven by judgement of governments' creditwortnisess.*" [Financial Times, 26 October 2007]. Developments in spreads of sovereign bonds might deviate from the level implied by domestic fundamentals for a long period of time and not only in short term.

Effective developments in the EMBI spreads for Hungarian bonds clearly provide such an example. Between May-2002 and January 2006 there was no change in the S&P long-term foreign currency of Hungary. The credit rating outlook index that translates the S&P rating on a numerical scale remains unchanged at 7 (which corresponds to A- with stable outlook)[5]. At the same time, the S&P change the rating outlook to negative in January 2006 and even decrease the rating to BBB+ in 2006. Rating developments could suggest that there

[5] More information regarding the credit rating outlook index are presented in the section desrcibing the data

was no major improvement in the Hungarian economy between 2002 and 2007 and, on the contrary, the things could even became worst (in fact the large budget deficit became larger and larger and the stabilization plan implemented in 2006 triggered a sharp increase in inflation rate and a slowdown in the economic growth). But the spreads for sovereign Hungarian bonds continued to decrease between 2002 and 2006.

Romania and Bulgaria offer also an interesting example. Spreads of Romanian and Bulgarian sovereign bonds compressed rapidly since 2007. Much more, they reached a record low level in the summer of 2007, going even below the level for the other regional countries with much better fundamentals. For instance, spreads for Bulgarian sovereign bonds stood only at 20 bp in May 2007 and at only 18 bp in June 2007, which were the lowest level among the EU member countries despite the fact that Bulgaria had poorer economic fundamentals (as reflected also by S&P ratings). We had the same story in the case of Romania. In May 2007 EMBI spreads for Romanian bonds only at 26 bp, below the levels of spreads for Hungary and Poland. And this despite the fact that fundamentals of Romanian economy were clearly poorer than ones of the other countries (current account deficit in Romania climbed to a record level of around 14% of the GDP in the early 2007). The two countries became a full member of the European Union on 1 January 2007 and this might be a factor which might explain the compression in spreads.

Figure 3 EMBIG spreads and Credit rating outlook for Hungary

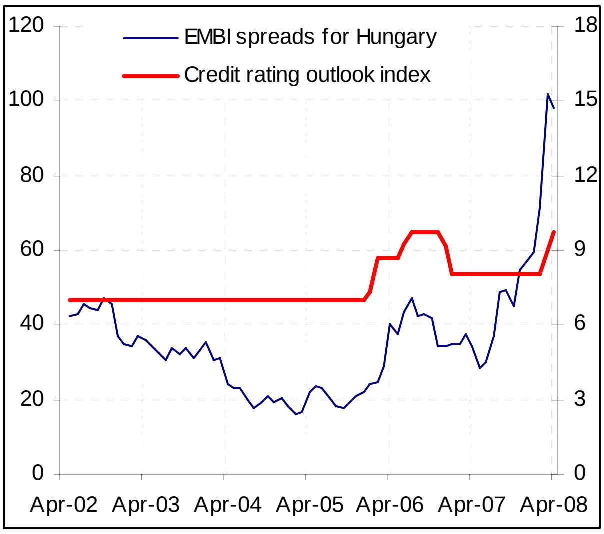


Figure 4 EMBIG spreads for CEE countries

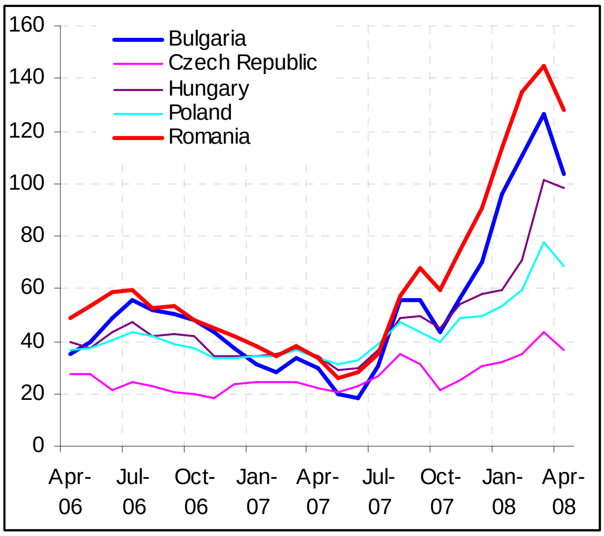


Table 1 EMBIG spreads for CEE countries at the middle of 2007

	Country	Spreads in May 2007	CROI	S&P Rating
1	Slovakia	19.6	6	A, Stable
2	**Bulgaria**	**19.7**	**8**	**BBB+, Stable**
3	Czech Republic	20.4	6	A-, Pozitive
4	Lithuania	20.9	6.52	A, Negative
5	**Romania**	**26.1**	**10**	**BBB-, Stable**
6	Croatia	27.8	9	BBB, Stable
7	Hungary	28.6	8	BBB+, Stable
8	Poland	31.6	7	A-, Stable

	Country	Spreads in June 2007	CROI	S&P Rating
1	**Bulgaria**	**18.3**	**8**	**BBB+, Stable**
2	Croatia	21.5	9	BBB, Stable
3	Lithuania	21.7	7.7	A, Negative
4	Slovakia	22.2	6	A, Stable
5	Czech Republic	22.7	6	A-, Pozitive
6	**Romania**	**28.0**	**10**	**BBB-, Stable**
7	Hungary	30.1	8	BBB+, Stable
8	Poland	33.0	7	A-, Stable

The influence of external factors on the spreads of bonds for emerging markets was also proved by the increase in these spreads in the second half of 2007, shortly after the beginning of US subprime crisis. Spreads of some emerging markets, especially the ones for countries from Europe like Slovakia or Poland, increased in last months despite the strong fundamentals of these economies.

Given this situation, in order to study the determinants of the spreads for sovereign spreads the empirical studies make use of the following reduced-form equation:

$$\log(s_{i,t}) = \alpha_i + \beta_i X_{i,t} + \gamma_i Z_t + \varepsilon_{i,t} \tag{4}$$

where $s_{i,t}$ is the spread for country i at t, $X_{i,t}$ is a set of domestic fundamentals for country i at t (likes ones previously presented), Z_t is a set of external factors reflecting the degree of risk appetite of international investors and with an potential impact on the spreads , β_i is the vector of coefficients for domestic fundamentals for country i, γ_i is the vector of coefficients for external factors for country i, α_i is an intercept, and $\varepsilon_{i,t}$ is an error term.

Instead of using a set of variables for domestic fundamentals, many empirical studies are using a index of cardinal numbers assigned to the sovereign long-term credit ratings of the country from one of the international rating agencies (Standard & Poor's, Fitch, Moody's). The implicit assumption is that the developments in the credit ratings are a good proxy for the developments in the fundamentals of the country (and this should be the case given that the international rating agencies are basing their credit ratings on the developments of fundamentals in the each country).

The set of external factors ($Z_{i,t}$) includes the variables which measures the degree of risk appetite of international investors. The risk appetite of international investors is not direct observable. Gonzalez-Hermosillo (2008) considers that four different global market risk factors are assumed to reflect the degree of risk appetite: (1) the funding liquidity premium which might be proxy by monetary condition, (2) the default risk, (3) the market liquidity risk which takes into account the preference of investors for liquid instruments, and (4) the market volatility premium. From a practical point of view, there are some market indicators which might be considered a proxy for the degree of risk appetite. The 3-months-ahead federal funds futures rate is usually used to measure the global funding liquidity risk and the credit availability in the global financial system. The market volatility is usually measured by the Chicago Board of Option Exchange (CBOE) Volatility index. The credit risk premium could be measured by a spread between the credit swap rate and the Treasury bond yield both for a long-term maturity (10 years). The market liquidity premium could be proxy by the difference between the yields for government securities with long term maturities and the yields for government securities with shorter maturities. In most of cases, the previous variables refer to the US economy.

Credibility of polices pursued by a country might be also an important factor for sovereign bonds dynamics. If policies are “good”, they will presumably reduc borrowing costs more if markets believe they would remain good in the future. If policies have been good but the government announces that it would temporarily deviate from past policies, e.g. to counteract a severe economic shock, credibility can help long-term market expectations despite the temporary deviation from the norm. Hauner and others (2007) consider the EU new member countries as an interesting case study of the effect of policy credibility on borrowing costs. EU accession has improved policy credibility, at least initially, in these countries.

There are situations in which an increase in the spreads for a country where fundamentals have deteriorated or are perceived to be weaker than expected due to a change in the sentiment of international investors about that country triggers also a increase in the spreads of other countries with good economic fundamentals. The discovery of a bad news about one country may cause investors to revise their expectations about the fundamentals of other specific countries which share similar features. This might happen for instance because the international investors own in their portfolio debt instruments issued by more countries and they have to rebalance their portfolio. In order to asses the spillover effects from one country to the other countries a usual approach is to evaluate the co-movements in the volatility of returns of financial instruments (in these case the returns for sovereign bonds). The volatilities are estimated by different specification of GARCH models, both unvaried and multivariate.

IV. Empirical analysis

We are interested in evaluating the determinants of EMBIG spreads for the Romanian bonds. Also we are interested in assessing the existence of a common pattern in the volatility of daily returns of sovereign bonds issued by the European Emerging Countries and testing for spillover effects among the countries.

In the first part of the analysis will we use data for 11 Emerging Economies and we will estimate a reduced form equation as in (4) :

$$\log(s_{i,t}) = \alpha_i + \beta_i X_{i,t} + \gamma_i Z_t + \varepsilon_{i,t}$$

where $s_{i,t}$ are the EMBIG spreads of country i at t, $X_{i,t}$ is a credit outlook index (CROI) for country i at t which is computed from S&P sovereign ratings and Z_t is a set of variables which reflects the risk appetite of foreign investors on the external markets. In line with other empirical studies, we begin by assuming that Z_t includes ones of the following indicators: the volatility index VIX, the 3-months FED funds future rate, the volatility of the deviation of the FED funds future rate from the FED funds rate.

Statistical test reveals that all the time series are not stationary, but that there are some cointegration relations between variables. We estimate the equation in (4) using two estimation methods: (1) a fixed effects pool model and (2) the poll mean group (PMG) estimator of Pesaran, Shin, and Smith (1997) which allows for an explicit long-term relation in the variables. Given that the variables are I(1) and cointegrated, the residuals from the first model could be considered as a deviation from the lon-run equilibrium. We compare them with the deviation implied by long-run relationship in the second model.

In the second part of the analysis, we are interested to find out if there is a co-movement or if there are spillover effects between the volatilities of prices of Romanian Eurobonds and prices of other emerging markets' bonds. We estimate Component-GARCH models for the volatility of daily returns of 5 sovereign bond prices for Emerging Countries from Europe (including Romania), and we test for a common pattern and volatility spillovers between the estimated volatilities.

This section of the paper has three parts. In the first part we present the data used in the analysis and perform statistical tests of stationary. In the second part, we focus on the determinants of the EMBIG spreads and we estimate the two panel models. In the final part, we estimate the Component-GARCH models and we test for a common pattern and spillover effects between volatilities.

IV. I. Data used in analysis

EMBIG spreads

In the analysis we use as dependent variable the euro denominated Emerging Markets Bonds Index Global (EMBIG) spreads computed by J.P. Morgan. For each country the index tracks the weighted averages of yield spreads over the German reference rates of external debt instruments denominated in euro. Emerging Market Bond Index Global (EMBIG) was launched in 1998 due to investors' requirement for a benchmark that includes a broader number of countries. Before the lunch of the EMBIG, J.P. Morgan computed only the (merging Market Bond Index Plus (EMBI+). Selection criteria for including an instrument in EMBIG are less restrictive than ones for the EMBI+. The instruments included in the index do not have to satisfy additional liquidity criteria such as a minimum bid/sell price and a specific number of interbank quotations. Also, the sfere of instruments included in the EMBIG are larger than one for EMBI.

In order to be included in the index one country has to fulfill two types of requirements:

A. Country admission requirements (criteria which determine if a country is defined as an emerging market) :
   - The country has to be classified as having low or middle per capita income according to World Bank, *or*
   - It has restructured external or local debt in past ten years, or
   - Currently has restructured the external or local debt outstanding;

B. General instrument admission requirements:
   - The face amount of outstanding debt of at least 500 million euro, and

- When added to the index the instruments should have at lest 2 ½ years till maturity, and
- Daily price must be available either from J.P. Morgan or from an outside source.

The country weights attributed to each country are computed by the aggregation of the weights of all its instruments included in the index. The weight of each instrument in the EMBIG index is calculated by dividing the issue's market capitalization by the total market capitalization for all the instruments included in the index. The market capitalization for each instrument is computed by multiplying its outstanding face value amount by the bid side of the settlement price.

We retain into analysis the spreads for 11 countries for which we have continuously observation from May 2002 to April 2008. Five of the eleven countries used in the analysis are currently members of the European Union: Poland, Hungary, Slovakia, Romania, and Bulgaria. The other countries included in analysis are: Croatia (currently a candidate country to the European Union), Turkey, South Africa, Brasil, Mexic, and Venezuela.

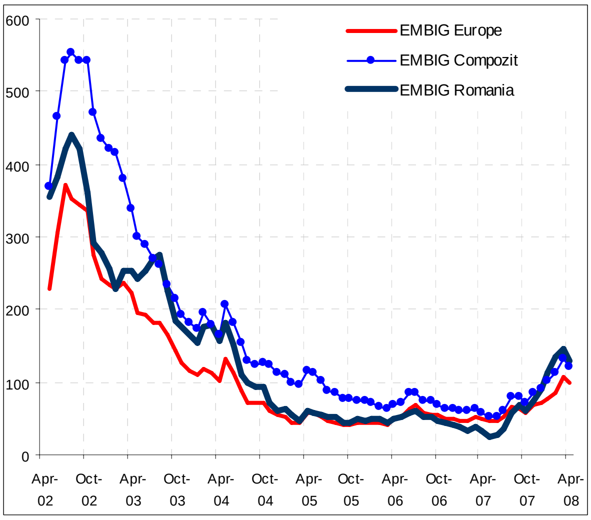
The data used in analysis start in May 2002 and end at the beginning of May 2008. We assess the determinants of the level of EMBIG spreads we use data with monthly frequency computed as simple average of daily observations. There are 72 observations available for each of the 11 countries.

In the second part of the analysis (estimation of the Component-GARCH models) we are using the EMBIG price indexes which are used to derive the performance of a portfolio. We compute the daily return for each of the eleven bonds retained in the analysis. This time, we have 1574 observations for each country.

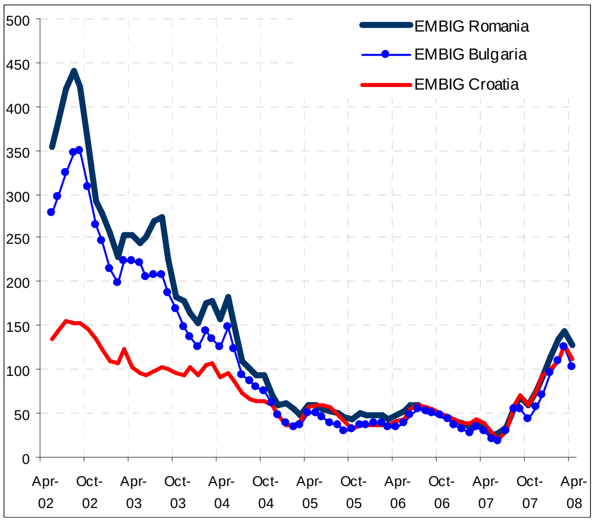
Dynamics of the EMBIG spreads for the 11 countries retained in the analysis, at monthly frequency, is presented in Figure 5, panels A to D. The Figure 5 reveals the decrease in the EMBIG spreads for all 11 countries between 2002 and 2007. The countries which became a full member of the European Union in 2004 have the lowest spreads, while countries from the Latin America have the largest spreads. However, EMBIG spreads for these countries decreased rapidly since 2002. There is a strong correlation between the moves in the EMIG spreads for Romania and the EMBIG spreads for Bulgaria, the countries which have become a full member of the European Union in January 2007. There is also a strong correlation between the EMBIG spreads for Romania and the EBIG spreads for Hungary.

Figure 5. Dynamics of EMBIG spreads for the Emerging Markets

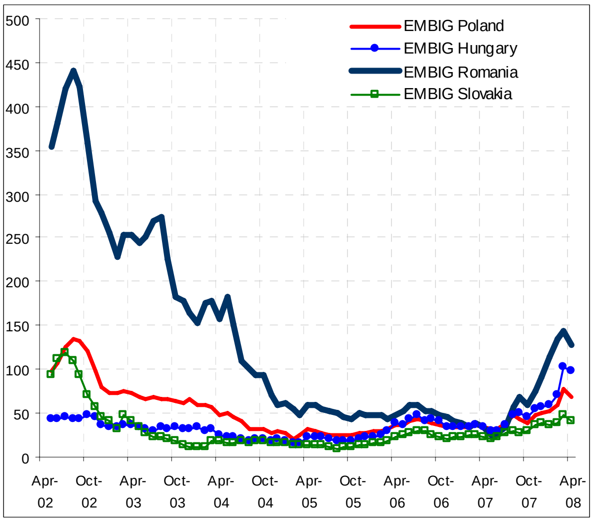
Panel A. EMBIG spreads for Romania, for countries from Europe and Composite



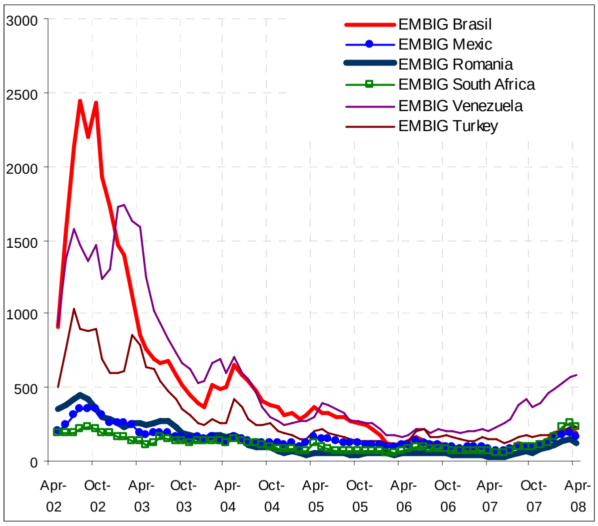
Panel B. EMBIG spreads for Romania, Croatia and Bulgaria



Panel C. EMBIG spreads for Romania, Poland, Hungary and Slovakia



Panel D. EMBIG spreads for Romania, Turkey, South Africa, Mexic, brasil, Venezuela



Credit Rating Outlook Index (CROI)

One of the most important determinants of the EMBIG spreads are specific country fundamentals such as exchange rate regime, inflation, GDP, current account performance, external debt, national savings, accumulation of foreign exchange reserves, fiscal policies etc. At the same time, the long term sovereign ratings for each country provided by the international rating agencies could be considered as an aggregate indicator which reflects the developments in the fundamentals of each country. So, the ratings might be used instead of the set of variables which are country's fundamentals. The problem is that the rate scale of the international rating agencies is qualitative-hierarchic and it cannot be directly used in the quantitative estimations. In order to use the sovereign ratings in estimations it is necessary to convert these ratings on a numerical (cardinal) scale.

We decided to use the credit rating outlook index computed by K. Hartelius, K. Kashiwase, L.E. Kodres (2008) on the basis of long term rating in foreign denominated currency and country outlook provided by Standard and Poors rating agency[6]. The CROI index computed by the authors takes into account not only the effective ratings, but also the outlook of the ratings which the authors founded out to provide useful information. In constructing the CROI index, the authors divided the countries in three categories: investment grade (countries with long term rating from AAA to BBB-), noninvestment grade tier 1 (countries with long term ratings from BB+ to CCC+) noninvestment grade tier 2 (countries with long term ratings from CCC to SD). The CROI index vary between 0 and 22, with highest value (22) corresponding to the worst country rating. The relation between the CROI index and the Standard and Poor's sovereign ratings is presented in Appendix 1.

There are three important properties of CROI. Firstly, in the investment grade category we can observe two important aspects. On the one hand, for a country with long term rating and a positive outlook CROI value is lower than for a country with a one notch higher long term rating and a negative outlook. On the other hand, an increase in CROI responding to a negative outlook is greater than reduction in CROI responding to a positive outlook. Secondly, when an outlook improves from stable to positive this change is reflected into a higher reduction in the investment category (1 point) CROI than in noninvestment category tier 1 (0.9 points); also when an outlook changes from stable to negative this deterioration is reflected into an equal increase in CROI for both investment and non investment grade tier 1 countries. Thirdly, there is no distinction in CROI value for countries from the non investment grade tier 2 with the same long term rating different outlooks (positive, stable and negative).

International rating agencies adjust their ratings only at discrete moments in time, while the changes in fundamentals in the economy are continuously. Although in the long run the sovereign ratings would move in line with the fundamentals of the economy, there is the possibility that in the short run (several weeks or months) the developments in ratings would deviate from changes in fundamentals. Accordingly, we decided to smooth the Credit Rating

[6] Instead of using a single credit rating outlook index P.Luengnanaruemitchai and S. Schader (2007) use three indices of fundamentals: economical, political and financial indices.

Outlook Index (CROI) by a Hodrick-Presscot filter with a very low value for the smoothing parameter λ ($\lambda = 15$). The filtered CROI series would be used further in the estimations.

The existing relationships between the S&Pratings for long-term foreign currency and the CROI index (both in the original form but also after the filtering operation) are presented in the Appendix 2. We can see that in the long run there are many similarities between the evolution of CROI and of sovereign ratings.

Volatility index of S&P 500 (VIX)

The Chicago Board Options Exchange Volatility Index (VIX) is a key measure of market expectations of near-term volatility (30 days) conveyed by S&P 500 stock index option prices[7]. The calculation is independent of any model. The index computation is based on a formula which derives market expectations of volatility directly from index options prices rather than an algorithm that implies baking implied volatility out from an option pricing model. The index came to be considered by many to be the world's premier barometer of investor sentiment and global market volatility. The VIX is often referred to as “investors gauge”. The reason for this name is that VIX is based on real time options prices, which reflects investors’ consensus view of future expected stock market volatility. Historically, during periods of financial stress which are accompanied by steep stock market decline options prices rise and also does VIX. Conversely VIX- tend to decline as market sentiment improves. Therefore, VIX may be considered a proxy for investors’ attitude towards risk and appears to explain movements of the emerging markets bond spreads in recent years (K. Hartelius, K. Kashiwase, L.E. Kodres 2008).

We used in the analysis data with monthly fervency (computed as simple average of daily observation) for period May 2002-May 2008. Dynamics of the VIX index is presented in Figure 6.

Fed Fund Futures rate

Following (K. Hartelius, K. Kashiwase, L.E. Kodres 2008) we use implied yield of 3 months ahead 30 days fed fund futures in order to reflect the short term interest rates and market expectations of US future policy rate. Implied yield of 3 months ahead 30 day fed fund futures has become a market wide benchmark for leveraged carry traded investors who

[7] http://www.cboe.com/micro/vix/introduction.aspx

borrow at the short term end of the yield curve to invest in emerging market. The investors all over the financial world keenly watch these interest rates in the periods preceding Federal Open Market Operations Committee (FOMC). Also, this rate has the advantage that it influences interest rates all along the US yield curve.

Volatility in the Fed Fund Futures

Volatility of fed funds futures is used as a measure of the uncertainty regarding the US monetary policy which is perceived to have a large impact on financial markets and on the process of financial assets allocation. This indicator is computed as the standard deviation of the difference between the implied yield on 3 month ahead 30 day fed fund futures and fed target rate using 90 days rolling window.

Dynamics of 3-months fed funds futures rate and of the volatility of its deviation from the FED funds rate (at monthly frequency) is presented in Figure 7..

Figure 6. Volatility index of S&P 500 (VIX)

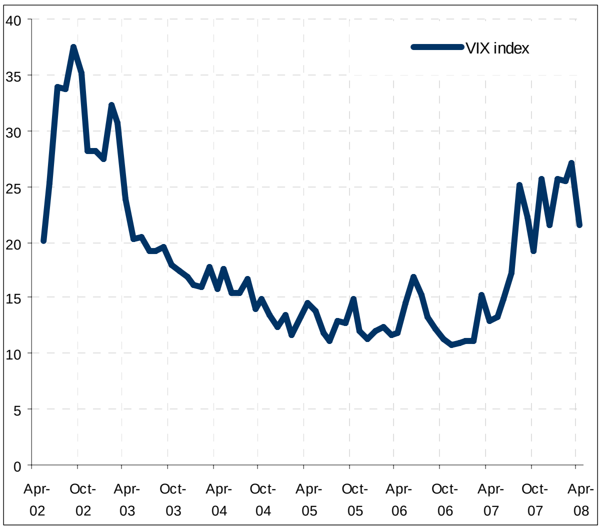
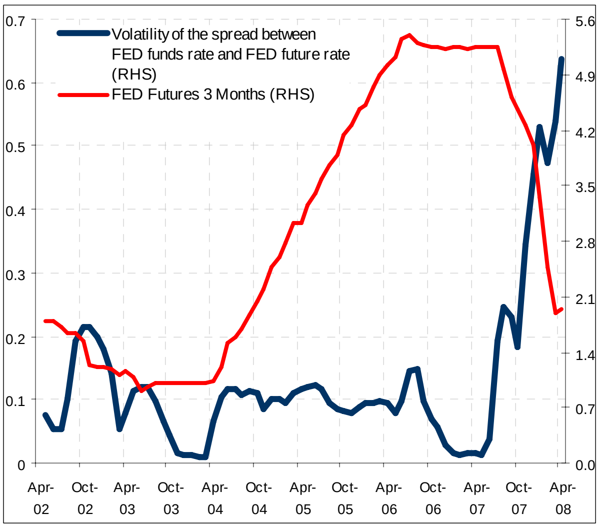


Figure 7. US monetary policy paramters



Unit root tests

Statistical properties of the data used in analysis are a key element in choosing the estimation techniques which would be used. For instance, many of the economic time series are non-stationary and this implies, for instance, the use of cointegration techniques. Accordingly, we start the analysis by testing the stationary of the time series for the variables that we intend to use in the estimations. We perform two categories of tests. Firstly we test the stationary of each data series individually using the Augmented Dickey Fuller and Phillips Perron tests. Then, because we will use panel estimation techniques, we will use also panel unit root tests. Recent literature suggests also that panel unit root tests have grater power than the unit root tests based on individual time series.

The results of the ADF and Phillips Perron unit root test are sumarized in the Appendix 3. The unit root tests for individual time series suggest that all series have unit roots. Although there is no theoretical reason to believe that the EMBIG series is not stationary in the long run, the unit root tests performed suggest the existence of a unit root. The logarithm of the VIX, but also the other two indicators of risk appetite on the international markets have also unit roots.

For testing the stationarity of the EMBIG time series (in logaritm) and of the filtered CROI time series (in logaritm) we use also the panel unit root tests available in Eviews. When performing the unit root test we allow for the presence of a intercept (and not for a trend) in the underlying equation of the test. The results of the panel unit root tests are presented in the Appendix 4. The panel unit root tests confirm also that there is a unit root in the EMBIG spreads (in logaritm) and also in HP filtered CROI series (in logaritm).

IV. II. Determinants of EMBIG spreads for Romania: estimation results

We want to explain the developments in the logarithm of EMBIG spreads of a country i $(s_{i,t})$ by its domestic fundamentals captured by the HP filtered Credit Rating Outlook Index $(hp\_croi_{i,t})$ and by a set of variables which measures the risk appetite of foreign investors: the logarithm of volatility index VIV $(\log\_vix)$, the 3-months futures on FED funds rate $(ff3m)$ and the volatility of the deviation of the 3-months futures on FED funds rate (v_ff1). The unit root tests showed that all of these variables are non-stationary, which means that the estimation of an OLS regression with the EMBIG spreads as a dependent variable and the other variables as the explanatory variables might not be preferable. Given that the the series are I(1) some of the cointegration techniques might be preferable.

The starting point of the analysis was to consider only the case of Romania and to try to explain the dynamics of Romanian EMBIG spreads $(\log\_embig\_ro)$ by the domestic fundamentals of the Romanian economy captured by the filtered CROI $(\log\_hp\_croi\_ro)$

and by external factors ($\log\_vix$, $ff3m$, v_ff1). Given that the variables are I(1), we tested for a cointegration relation between these variables using the Johansen cointegration procedure. But we don't succeed to find any long-run relationship between the spreads and any set of the external variables. Also, a regression of the EMBIG spreads on the three explanatory variables didn't perform better.

We extended then the analysis by considering also another 10 countries alongside Romania and move the attention to the panel estimation techniques. In fact, the panel estimation methods for the determinants of EMBIG spreads are largely used in the empirical analysis and it is supposed that they would result in better result than univariate methods.

We tested for the existence of a cointegration relationship between the variables using the panel unit root tests proposed by Pedroni (1999,2004) and by Kao (1999). The test showed that when considering the data for the 11 countries there is a cointegration relationship between $\log\_embig\_ro$, $\log\_hp\_croi\_ro$, and $\log\_vix$. The results of the panel cointegration tests are presented in the Appendix 5.

Given the cointegration between the tree variables we can use the panel regression estimations methods with more confidence. Two panel estimation methods are used:

1) A panel regression with fixed effects for the 11 countries with the logarithm of EMBIG spreads ($\log\_embig\_ro$) as the depended variable and the log of the CROI ($\log\_hp\_croi\_ro$) and the log of the VIX as explanatory variables ($\log\_vix$).
2) The pool mean group estimator due to Pesaran, Shin and Smith (1997) is used to find a long-run relationship between the log of EMBIG spreads ($\log\_embig\_ro$) and the other to variables ($\log\_hp\_croi\_ro$, $\log\_vix$).

Estimation results for the panel equation with fixed effects

We include 11 countries in the analysis: Romania, Bulgaria, Croatia, Poland, Slovakia, Hungary, Turkey, South Africa, Mexic, Brasil, and Venezuela. For each country there are available 72 observations with monthly frequency. We estimate the following equation with pooled data:

$$\log\_embig_{i,t} = \alpha + \beta_i + \delta \cdot \log\_croi_{it} + \gamma_i \cdot \log\_vix + \varepsilon_{i,t}$$

where $i = 1,2,...,11$ identifies the countries and $t = 1,2,...,72$ identifies the period of time.

The series of EMBIG spreads ($\log\_embig$) and of credit rating outlook indexes vary over the 11 cross section, while the series of VIX is the same for all the countries. The coefficient δ in the regression would be the same for all countries, which means that the EMBIG spreads reacts the same way at the changes in the domestic fundamentals in the case of any country. However, we assume that the EMBIG spreads reacts differently to changes in the risk appetite of investors. For instance, we expect the countries with weaker economic fundamentals or the countries which historically have been perceived as not implementing adequate macroeconomic policies to be penalized more when the sentiment on the international markets deteriorates. Changes in risk appetite of international investors are usually triggered by developments in the developed economy. In fact the developments in the US economy were at the root of changes in investor's sentiment on the international markets at the end of 2007 when the spreads increased. Usually, the emerging economies from the Latin America, the economy of South Africa and also the Turkish economy are more connected to the developments in the US economy. However, we will conduct a Wald coefficient test to see if the null hypothesis of identical coefficients for the VIX index across the countries might be or not accepted.

We allow at the same time for fixed effects. There might be other factors than the Credit Rating Outlook Index and the VIX index which are specific to each country in part (or to a group of two or more countries). For instance, the EU accession might have a specific impact on the new member countries. Due to an increase in the credibility of macroeconomic policies, the spreads of these countries might be lower than ones for the other emerging economies having the same value for the domestic fundamentals. Also, the degree in which a new EU member country has benefit from EU accession might be different as its policy credibility was different. As the credibility of macroeconomic policies is an unobservable variable it is difficult to be modeled separately. Also, there might other country specific factors that are not taken completely into account by the sovereign ratings. We think that as long as some heterogeneity exists in the data it is normal to assume the presence of fixed effects. We will test also for redundant fixed effects.

The estimation of the pooled equation is performed by imposing different weights to the observation. We estimate a feasible GLS specification in which we correct both for cross-section heteroskedasticity and contemporaneus correlation. The same specification is used when the standard errors of the coefficients are computed. There might be some cross-section correlation in the residuals as the spreads might react to other global factors which are not reflected in the evolution of the two explanatory variables.

The estimations results are presented in Table 2 and in Appendix 5. All coefficients are statistically significant and have the expected sign. An increase in the CROI index which reflects a deterioration of the country's fundamentals will trigger an increase in the EMBIG spreads. Also, an increase in the VIX index which reflects a decrease in the risk appetite on the international markets would result in higher EMBIG spreads. The higher coefficients for the VIX index was obtained in the case of two Latin American countries: 1.56 for Brasilia and for Venezuela. The lowest coefficients for the VIX index are obtained for Slovakia and (0.63) and South Africa. Coefficient for Bulgaria and Romania are also higher (1.24 and respectively 1.52). By estimation, the sum of crossed fixed effects is normalized to zero. But because we have a group of countries which covers a large part of the emerging markets economy, the value of the specific effect constant might be informative. For instance, this constant is negative for all new EU members excepting Slovakia (for Croatia and Hungary is close to zero), which means that these countries had lower borrowing costs than expected. On the other hand, Mexic, South Africa, Turkey and Slovakia had higher borrowing costs than expected. The coefficient of determination (R^2) is also very high 0.96, which means that the two variables explain a lot of the variance in the dependent variable. But there is also a problem with the estimation because there is autocorrelation in the residuals. This might due to the fact that not all determinants of the EMBIG spreads have been taken into account.

Table. 2 estimation result for the equation with pooled data and fixed effects

Dependent Variable: LOG_EMBIG_?
Method: Pooled EGLS (Cross-section SUR)
Included observations: 72
Cross-sections included: 11
Total pool (balanced) observations: 792
Linear estimation after one-step weighting matrix
Cross-section SUR (PCSE) standard errors & covariance (d.f. corrected)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-4.150236	0.196415	-21.12989	0.0000
LOG(HP_CROI_?)	2.544679	0.063320	40.18774	0.0000
BG--LOG(VIX)	1.245724	0.102703	12.12937	0.0000
BR--LOG(VIX)	1.561762	0.166648	9.371598	0.0000
CR--LOG(VIX)	0.953539	0.087435	10.90571	0.0000
HU--LOG(VIX)	0.880507	0.056080	15.70092	0.0000
MX--LOG(VIX)	0.823619	0.064358	12.79742	0.0000
PO--LOG(VIX)	1.176625	0.101113	11.63678	0.0000
RO--LOG(VIX)	1.159120	0.134922	8.591051	0.0000
SA--LOG(VIX)	0.785796	0.078891	9.960475	0.0000
SL--LOG(VIX)	0.637691	0.173224	3.681314	0.0002
TU--LOG(VIX)	0.877326	0.097576	8.991228	0.0000
VN--LOG(VIX)	1.318097	0.083019	15.87706	0.0000

Fixed Effects (Cross)	
BG--C	-0.729494
BR--C	-0.995616
CR--C	-0.032228
HU--C	-0.017427
MX--C	1.080493
PO--C	-0.617098
RO--C	-0.636171
SA--C	1.101355
SL--C	0.681114
TU--C	0.523515
VN--C	-0.358444

In order to test if the coefficient of the logarithm of the VIX is identical across countries, we use a Wald coefficient test. The results of the Wald tests reject this hypothesis, which means that we should allow for different coefficients for the VIX index across countries. The results of the test are presented in the following table.

Wald Test:			
Test Statistic	Value	df	Probability
F-statistic	32.25910	(10, 769)	0.0000
Chi-square	322.5910	10	0.0000

Further, we test for the joint significance of the fixed effects. The test again suggests that we should allow for fixed effects in the estimation of the equation.

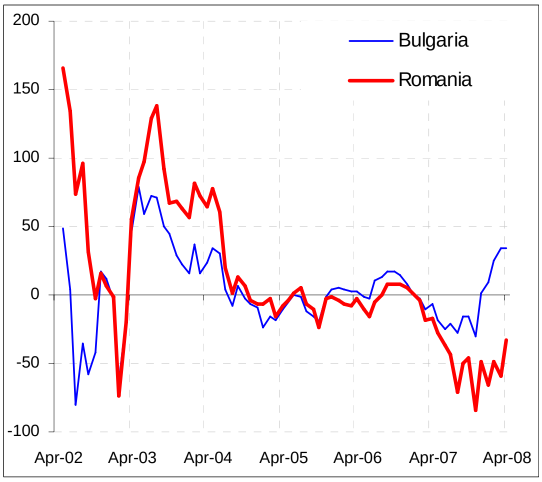
Redundant Fixed Effects Tests Test cross-section fixed effects			
Effects Test	Statistic	d.f.	Prob.
Cross-section F	45.238281	(10,769)	0.0000

The p-values associated to the F-statistic is 0, which provides strong evidence against the null hypothesis that the fixed effects are all equal to each other. This suggests that there is unobserved heterogeneity in the data and we should use a model with fixed effects.

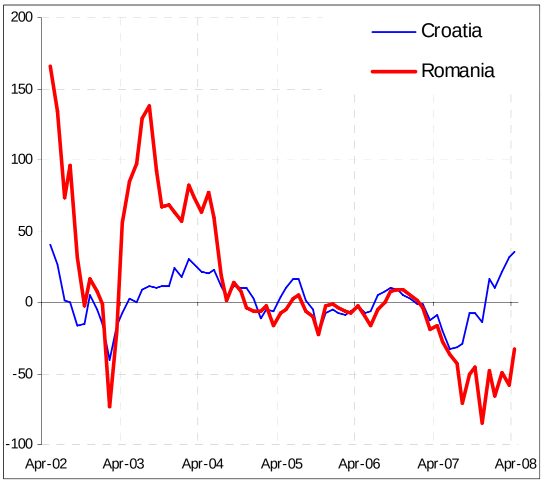
Given that the three variables are cointegrated, the errors from the pooled regression might be considered as a deviation from a long-run equilibrium relation. Deviations from the equilibrium level for Romania and the other 5 EU member countries included in analysis are presented in Figure 8 (A-E).

Figure 8. Residuals (deviations from the long run equilibrium) from panel estimation

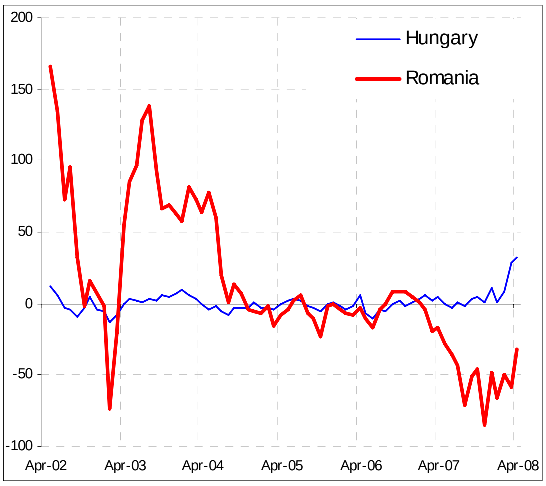
Romania vs. Bulgaria



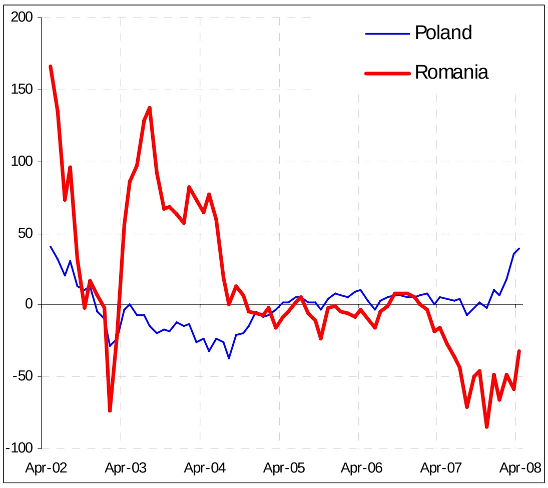
Romania vs. Croatia



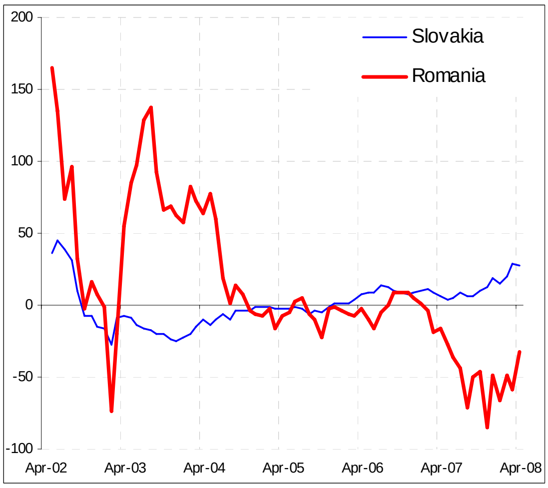
Romania vs. Hungary



Romania vs. Poland



Romania vs. Slovakia



We see that there is common pattern in the residuals from Romania on the one hand and Bulgaria and Croatia on the other hand. The three countries had the lowest sovereign ratings above the EU member countries and they made substantial economic progresses in the last years. The common pattern in the residuals might be explained by a common factor: for instance the investors could have perceived Romania and Bulgaria as being part of the second wave of the EU extension. The EMBIG spreads for Romania and Bulgaria started to increase in May 2003 after ten of the twelve candidate countries signed the Treaty of Accession to the European Union. Between May 2003 and June 2004, the spreads for the Romanian and Bulgarian bonds remained above their equilibrium level, while spreads for the new EU member countries (Poland and Slovakia) felt below the equilibrium level. On the other hand, the S&P rating agency continued to improve Romania and Bulgaria's sovereign ratings on the back of progresses in the economy. The EMBIG spreads for Romania and Bulgaria felt rapidly and moved to the equilibrium level in June 2004 when the EU accession moment o these two countries was confirmed for 2007. The EMBIG spreads for Romania and Bulgaria felt again, this time below the equilibrium level, in the first half of 2007 after the two countries have became full member of the European Union. EU accession had also a clear impact on the spreads of the Poland, Hungary and Slovakia which fell below the equilibrium level at the moment when their accession to the European Union became a certitude (in 2003).

The Figure 8 reveals that despite the fact the EMBIG spreads for Romania increased from 26 bp in May 2007 to 130 bp in April 2008, the level from April 2008 was below the equilibrium level (the one implied by the domestic fundamentals and the external conditions). S&P rating agency decreased the rating outlook from "positive" to "stable" in April 2007 and from „stable" to „negative" in November 2007, on the back on an increase in the domestic external disequilibria (especially the larger and larger current account deficit). The EMBIG spreads for Romania continue to decrease following the previous S&P move and they started to increase only when the risk appetite on the exterbal markets increased (in the second half of 2007). We think that the decision of the S&P to downgrade Romania's rating oulook was appropiate as it clearly reflects increasing risks in the Romanian economy. However, the developments in the markets spreads reveals that the foreign investors's perception was more important for the developments in the EMBIG spreads at that moment. Also, the fact that the EMBIG spreads are currently below their equilibrium level is proved by the recent issue of Romanian Eurobonds from June 2008 when the Government had to pay a premium of around 175 bp, above the level of the spreads in the market and close to the equilibrium level estimated from this model.

Figure 9. Deviations of spreads from equilibrium for Romania, Bulgaria and Croatia

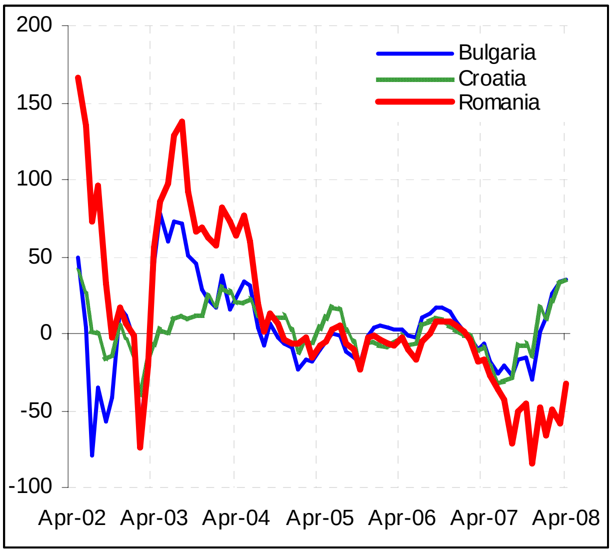
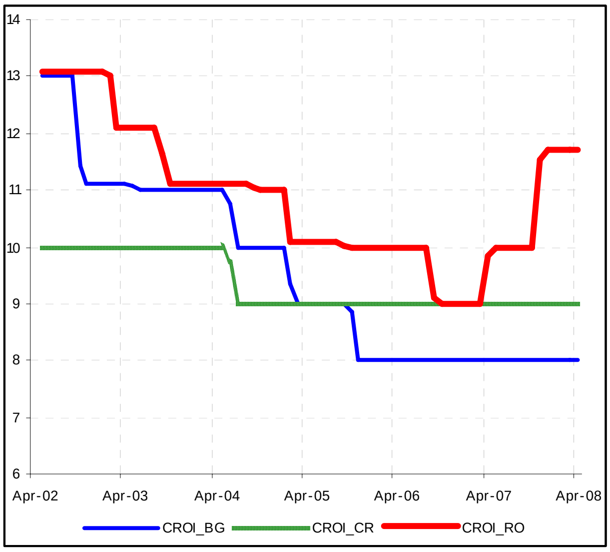


Figure 10. CROI index for Romania, Bulgaria and Croatia



Estimation results for the pool mean group estimator of Pesaran, Shin and Smith (1997)

The pool mean group (PMG) estimator introduced by Pesaran, Shin and Smith (1997) is applicable to panels with cross-section variation in the short run dynamics but lung-run communality in the equilibrium relationship. The PMG estimator constrains the long-run coefficients to be identical, but allows the short-run coefficients and error variances to differ across groups. An extension of the model allows only a subset of the long-run parameters to be the same across the groups while the others might be different. The PMG estimators lays between the extreme of fixed or random effects models that requires all slopes to be identical across groups and the very general model where the slopes are treated as completely unrelated (in this case separate regressions are performed for each group and a mean of the coefficients is computed (the mean group (MG) estimator).

Suppose that we have data on a number of time periods $t = 1,2,...,T$, and a number of groups, $i = 1,2,...,N$ and which to estimate an ARDL (p,q,q,…,q) model,

$$y_{i,t} = \sum_{j=1}^{p} \lambda_{i,j} y_{i,t-j} + \sum_{j=0}^{q} \delta_{i,j}^{'} X_{i,t-j} + \gamma_i^{'} d_t + \varepsilon_{it}$$

where $X_{i,t}$ $(k \times 1)$ and d_t $(s \times 1)$ are vectors of explanatory variables (repressors), the $X_{i,t}$ vary over both time periods and groups and the d_t only over time periods. T must be large enough than we can estimate the model for each group, but need not be the same for each group. It is also straightforward to allow for different lag orders on the different variables in

X_{it}. The coefficients of the lagged dependent variables $\lambda_{i,j}$ are scalars and $\delta_{i,j}$ and γ_i are $k \times 1$ and $s \times 1$ vectors of unknown parameters. The depend variables y_i and the explanatory variables in X_i might be non-stationary.

After appropriate transformations, the previous equation can be written in a error correction form:

$$\Delta y_{i,t} = \phi_i \left(y_{i,t-1} - \theta' \cdot X_{i,t} \right) + \sum_{j=1}^{p-1} \lambda_{i,j}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \Delta X_{i,t-j} \delta_{i,j}^{*} + d_t \gamma_i + \varepsilon_{i,t}$$

for $i = 1,..., N$ where $\xi_i(\theta) = y_{i,t-1} - X_{i,t} \cdot \theta$ is the error correction component. The constant and the deterministic trend are included in d_t, while ϕ_i is the speed of adjustment towards the equilibrium level. In the case when only a subset of the long-run parameters are constrained to be the same across the groups, the matrix of explanatory variables is partitioned as $X_i = \left(X_{1i,} X_{2i} \right)$ where X_{1i} corresponds to the variables in the long-run relationship which have the same coefficient across the groups. In this case the equation can be written in the error correction form:

$$\Delta y_{i,t} = \phi_i \left(y_{i,t-1} - X_{1i,t} \cdot \theta_1 - X_{2i,t} \cdot \theta_{2,i} \right) + \sum_{j=1}^{p-1} \lambda_{i,j}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \Delta X_{i,t-j} \delta_{i,j}^{*} + d_t \gamma_i + \varepsilon_{i,t}$$

where $\theta' = (\theta_1^{'}, \theta_2^{'})$ is the vector of coefficients in the long-run relationship.

In our situation, the dependent variable will be the EMBIG spreads $(\log\_embig\_ro)$ for the 11 countries. We will estimate an error correction model for the EMBIG spreads using as the explanatory variables in the long-run relation the CROI $(\log\_hp\_croi)$ and the VIX index $(\log\_vix)$. We will impose the restriction that the coefficient for the CROI index is the same across the countries but we will allow again the coefficient for the volatility index VIX to vary across countries, based on the same arguments as in the case of the previous panel estimation. The following equation in the error correction form is estimated:

$$\Delta \log\_embig_{i,t} = \phi_i \left(\log\_embig_{i,t-1} - \theta \cdot \log\_hp\_croi_{i,t} - \theta_i \log\_vix_t \right) + \sum_{j=1}^{p_i-1} \lambda_{i,j} \Delta \log\_embig_{i,t-j} +$$

$$+ \sum_{j=0}^{q_i-1} \lambda_{1i,j} \Delta \log\_hp\_croi_{i,t-j} + \sum_{j=0}^{r_i-1} \lambda_{2i,j} \Delta \log\_vix_{t-j} + c_i + \varepsilon_{it}$$

where $i = 1,2,...,11$ denotes the countries and $t = 1,2,...,72$ denotes the time periods.

The previous error correction model has some particular features: (1) in the lon-run relationship the coefficient for the CROI is the same across the countries (θ), while the coefficient of the volatility index VIX differs across the countries (θ_i); (2) the coefficient reflecting the speed of adjustment towards the equilibrium level is different across the countries (ϕ_i); (3) the coefficients for the lags of variables differ across the countries ($\lambda_{i,j}$, $\lambda_{1i,j}$, $\lambda_{2i,j}$); (4) the number of lags for the dependent variable might be different from the number of lags of the explanatory variables, which are also different from variable to variable. At the same time, the specification of the ARDL might different from a country to another country. In short, for a given country i will will estimate an ARDL model of the form $ARDL(p_i, q_i, r_i)$. All these features allow dealing better with the heterogeneity in the data.

To estimate the previous error correction model we use the GAUSS code of Pesaran, Shin and Smith (1997). The number of lags for the dependednt and explanatory variables is selected by minimization of Schwarz Information Criterion. The estimation results for the 11 countries are presented in the Appendix 7. In the case of Romania, an ARDL(2,0,1) model has been selected by the Schwarz Information Criterion. The results for the Romania are the folowing:

$$\begin{aligned}\Delta \log\_embig_{RO,t} &= \phi_{RO}\left(\log\_embig_{RO,t-1} - \theta \cdot \log\_hp\_croi_{RO,t} - \theta_{RO}\log\_vix_t\right) + \\ &+ \lambda_{RO,1}\Delta\log\_embig_{RO,t-1} + \lambda_{2RO,1}\Delta\log\_vix_t + c_{RO} + \varepsilon_{\mathrm{RO,t}}\end{aligned}$$

$$\begin{aligned}\Delta \log\_embig_{RO,t} &= \text{-}0.1185\left(\log\_embig_{RO,t-1} - 2.5270 \cdot \log\_hp\_croi_{RO,t} - 1.4888\log\_vix_t\right) + \\ &+ 0.2479 \cdot \Delta\log\_embig_{RO,t-1} + 0.3451\Delta\log\_vix_t \text{ - } 0.6844 + \varepsilon_{\mathrm{RO,t}}\end{aligned}$$

Variable	**Coefficient**	**Standard error**	**T-statistic**
ϕ_{RO}	-0.1185	0.0327	-3.6262
θ	2.5270	0.2249	11.2336
θ_{RO}	1.4888	0.3325	4.4782
$\lambda_{RO,1}$	0.2479	0.0949	2.6118
$\lambda_{2RO,1}$	0.3451	0.1063	3.2467
c_{RO}	-0.6844	0.2005	-3.4141

Summary statistics and diagnostics:

SIGMA	CH-SC	CH-FF	CH-NO	CH-HE	RBARSQ	LL	AIC	SC
0.11	0.89	3.53	0.17	0.35	0.44	61.00	55.00	48.25

Note: (a) SIGMA = standard deviation of the regression; (b) CH-SC = Chi-squared test of residual serial correlation; (c) CH-FF = Chi-squared test of functional form misspecification. (d) CH-NO = Chi-squared test of normality of residuals. (e) CH-HE = Chi-squared test of heteroskedsticity.

All regression coefficients are statistically significant. Also the low level low level of the statistics used to test residual of the regression for normality, heteroskedasticity, and serial correlation prove that the estimation equation is valid. Some comparisons with the results obtained for the other countries are also interesting. The coefficient reflecting the speed of adjustment towards the equilibrium level is negative $\phi_{RO} = -0.1185$ and is placed in the lower part of the range of coefficients for all set of the countries. The coefficient is again very close to the one estimated for Bulgaria. The coefficient for the VIX index is also very close to the value estimated for Bulgaria and Venezuela, it is lower that the values for Turkey and Brasil and it is higher than the values estimated for Hungary and Mexic. The estimation didn't produce satisfactory results for Poland and Slovakia.

Assuming that the intercept is also a part of the long-run relationship, we will have than in the long-term the equilibrium level for the EMBI spreads for Romania would be:

$$\log\_embig\_eq\__{RO,t-1} = 2.5270 \cdot \log\_hp\_croi_{RO,t} + 1.4888 \cdot \log\_vix_t - 5.7755$$

We remember that the pool estimation resulted in the following equation for Romania:

$$\log\_embig\_eq\__{RO,t-1} = 2.5447 \cdot \log\_hp\_croi_{RO,t} + 1.1591 \cdot \log\_vix_t - 4.6864$$

We observe that the coefficient of the Credit Rating Outlook has similar values in the two models, while there are some differences in the case of the coefficient for the VIX index. However, there are not large differences between the deviations of EMBIG spreads from their equilibrium level (Figure 11) in the two models. Also, the equilibrium level is close in the two models (Figure 12). When computing the equilibrium level of the Romanian EMBIG spreads we used a HP filter with $\lambda = 15$ (monthly frequency data) to smooth the VIX index.

Figure 11. Deviation from the equilibrium level in the case of the two models

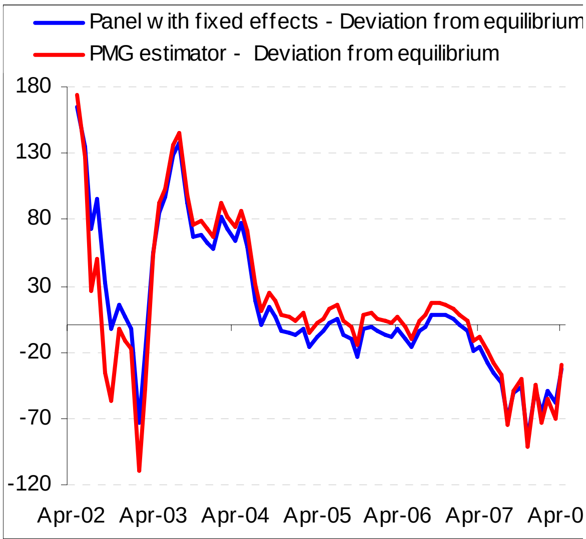
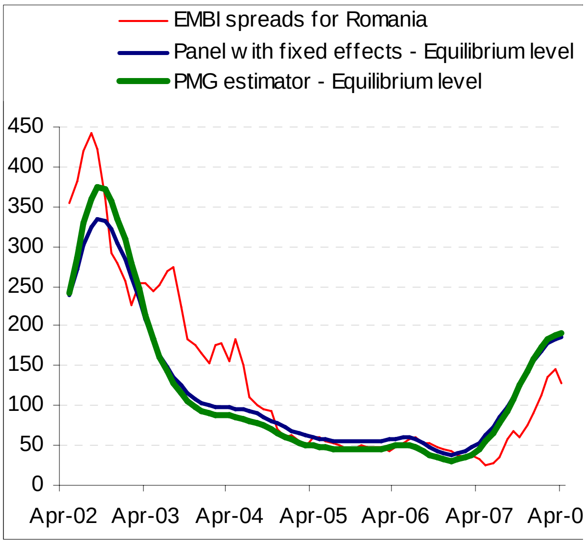


Figure 12. EMBIG spreads for Romania and their equilibrium level



The spreads for the Romanian sovereign bonds decreased by 225 bp between May 2002 and April 2008. The estimated model, based on the long-run equilibrium relationship, implies only a decrease of 51 bp. The higher decrease in the effective spreads is due to the fact that the Romanian bonds were undervaluated in 2002 (the spreads were above their equilibrium level) and they were overvaluated in April 2008 (the spreads were below their equilibrium level). The 51 bp decreased based on the equilibrium level is due exclusively to the fundamentals (as reflected by the decrease in the S&P sovereign rating), while the external factors had no impact during this interval. This is because following the crisis on the international markets the VIX index returned to the same level as in 2002, which means that the investors started to price appropriately the risk. There is also a practical implication from these observations: in the long run, a country cannot bet on the external factors to reduce its borrowing costs. Rather, it should implement appropriate domestic policies in order to improve domestic fundamentals.

IV. III. Co-movements and spillover effects in the daily returns of sovereign bonds of European Emerging Countries

In this section of the paper we are interested in testing for the existence of a co-movement in the prices of sovereign bonds and for the existence of spillover effects between the Emerging Countries from Europe. The estimations from the previous section showed that there was a common pattern in the deviation of equilibrium for Romania, Bulgaria and Croatia. Also there was a connection between the developments of EMBIG spreads for Poland, Hungary and Slovakia when their EU accession was validated.

In order to perform such an analysis, we consider this time the price index of EMBIG spreads for the six CEE countries: Poland, Hungary, Slovakia, Bulgaria, Romania, and Croatia. Based on these price indexes we have computed the daily returns. We have 1574 daily returns for each of the six series from May 2002 to the beginning of May 2008.

In order to find a comovement between EMBIG spreads volatility across countries included into analysis we employ a Component GARCH model (CGARCH) in the spirit of Engle and Lee (1993). The model decomposes conditional variance of the daily return series into a stochastic permanent or long run trend and a transitory or short run component. We decided to use a CGARCH model in order to have a much better image above the sources of the co-movements in the volatility of sovereign bond returns.

Theoretical model

For each country we estimate (in Eviews) a CGARCH(1,1) model described by the following set o equations:

$$r_t = c \cdot \sigma_t^2 + \varepsilon_t, \text{ with } \varepsilon_t / I_{t-1} \sim N(0, \sigma_t^2) \tag{5}$$

$$\sigma_t^2 = q_t + a_1 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) + a_2 \cdot (\sigma_{t-1}^2 - q_{t-1}) + a_3 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) \cdot D_{t-1} \tag{6}$$

$$q_t = \omega + b_1 \cdot q_{t-1} + b_2 \cdot (\varepsilon_{t-1}^2 - \sigma_{t-1}^2) \tag{7}$$

Equation (5) is the mean equation, where r_t is the log difference of EMBIG index and hence the daily rate of return for the sovereign bond prices. ε_t reflects any unexpected change in EMBIG index and is assumed to be uncorrelated and conditionally normal distributed taking into account the information set I_{t-1} containing all information available at moment $t-1$. We choose to introduce an ARCH in mean term in return equation which reflects the fact that the expected return on EMBIG index is related to the expected risk. Taking into consideration that we deal with a market index we can interpret the coefficient c as a measure of the risk aversion degree of investors.

Equation (6) models conditional variance as a function of a time varying intercept, the lag in the squared realized residuals (ARCH term), the lagged conditional variance (GARCH term) and an asymmetric term that augments the ARCH term whenever a lagged residual is negative. We include the asymmetric term in variance equation through a dummy variable that takes the value 1 in the case of a negative shock ($\varepsilon_{t-1} < 0$) and 0 otherwise. We consider that there is an asymmetric movement in the bond prices in the sense that bad news (meaning negative shocks) has a greater impact on prices (and spreads) than positive news (meaning a positive shock).

By analogy with the classical GARCH(1,1) model:

$$\sigma_t^2 = \varpi + a_1 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) + a_2 \cdot (\sigma_{t-1}^2 - q_{t-1})$$

the component model also allows mean reversion. But in comparison with the classical model which shows mean reversion to a constant level, ϖ, the component model exhibits mean reversion to a time varying long run level q_t:

$$\sigma_t^2 - q_t = a_1 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) + a_2 \cdot (\sigma_{t-1}^2 - q_{t-1}) \tag{8}$$

Equation (7) is the distinctive feature of CGARCH and models the long run time varying component of conditional variance. The component depends on a time invariant permanent level ω, an AR term b_1 and a forecast error b_2 which is the difference between the lag of squared residuals and the forecast variance from the model on the basis of information available at time t-2. As the model shows, the long run permanent level is allowed to vary due to the forecast error, but on the long run it converges to the permanent level ω with power b_1 provided that $|b_1| < 1$.

Equation (8) describes the transitory or the short run component of conditional variance, $\sigma_t^2 - q_t$, which converge to zero with the power $a_1 + a_2$. The condition for this dynamics to hold is that $a_1 + a_2 < 1$. The reason for this inequality is the following. Taking into consideration the fact that in equation (7) the term $\varepsilon_{t-1}^2 - \sigma_{t-1}^2$ has zero expected value, accounting for all available information at moment t-1 the expected value of the long run volatility will be: $q_{t+n} = \dfrac{1 - b_1^n}{1 - b_1} \cdot \omega + b_1^n \cdot q_t$. Therefore, the transitory component at time t+n will have the form: $\sigma_{t+n}^2 - q_{t+n} = (a_1 + a_2)^n \cdot (\sigma_t^2 - q_t)$ which will converge to zero as n approaches infinite and the conditional variance will reach its trend in the long run.

Combining the above two conditions we get that if $a_1 + a_2 < b_1$ than the short run component will converge faster than the long run component which implies that over time the transitory component converges to zero and aggregate volatility converges to its long run trend. Also, if $b_1 = 1$, then the permanent component to which long term volatility forecasts mean revert is just a random walk.

In addition we also need to specify a set of conditions for ensuring positive values for out of sample variance forecasts: i) $0 < a_1 + a_2 < b_1 < 1$; ii) $0 < b_2 < a_2$; iii) $a_1 > 0$ and $\omega > 0$.

Estimation results

We implement the component GARCH model in a univariate manner, respectively we estimate for each of the six Eastern European Countries included in the analysis (Bulgaria, Croatioa, Romania, Poland, Hungary, and Slovakia) a model of the form described above:

$$r_t = c \cdot \sigma_t^2 + \varepsilon_t, \text{ with } \varepsilon_t / I_{t-1} \sim N(0, \sigma_t^2)$$

$$\sigma_t^2 = q_t + a_1 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) + a_2 \cdot (\sigma_{t-1}^2 - q_{t-1}) + a_3 \cdot (\varepsilon_{t-1}^2 - q_{t-1}) \cdot D_{t-1}$$

$$q_t = \omega + b_1 \cdot (q_{t-1} - \omega) + b_2 \cdot (\varepsilon_{t-1}^2 - \sigma_{t-1}^2)$$

We estimated the model with Eviews using the maximum likelihood estimation which has the advantage that generates an estimator which has all the properties of a maximum likelihood estimator. Therefore, the estimator is consistent, unbiased, asymptotical efficient.

We didn't obtain a satisfactory result in the case of Slovakia .We had also problems with Slovakia with the estimation of the pooled mean group estimator. The summary of the estimations results in the case of the other five countries are presented in Table 3. In order to validate our results we perform two test: the Ljung-Box Q statistic test in order to check for the existence of residuals autocorrelation and the ARCH – LM test in order to check for the existence of heteroskedastic effects. The Ljung-Box Q Statisctic is computed as: $Q_{LB} = T \cdot (T+2) \cdot \sum_{j=1}^{k} \frac{\tau_j^2}{T-J}$ under the null hypothesis that there is no autocorrelation up to the lag k. In all cases the results failed to reject the null hypothesis which means that there is no autocorrelation in residuals. The ARCH LM test under the null hypothesis that there is no ARCH effect up to the order q in residuals is computed by running a regression of the squared residuals on a constant and lagged squared residuals up to the lag q: $e_t^2 = b_0 + \sum_{k=1}^{g} b_k \cdot e_{t-k}^2 + v_t$. We report the result of the Obs*R-squared statistic which is asymptotically distributed as $\chi^2(q)$ in the Table 3. We performed the test for each model using various numbers of lags and in all cases the results failed to reject the null hypothesis (we obtained high values for the associated p-value of the statistic which means that there are no ARCH effects in the residuals).

Coefficients are generally highly significant (at 1% significance level) with few exceptions. In the long run component of volatility we found a positive and highly significant constant (ω) for all countries. The trend AR term of the permanent volatility (b_1) is also significant and it exhibits high levels in the vecinity of 0.99 so that q_t approaches ω very slowly. The coefficient of the forecast error (b_2) which shows how the permanent component of volatility is affected by shocks is positive and significant for all five countries.

Table 3: CGARCH Estimates Sample period: May 2002 – May 2008

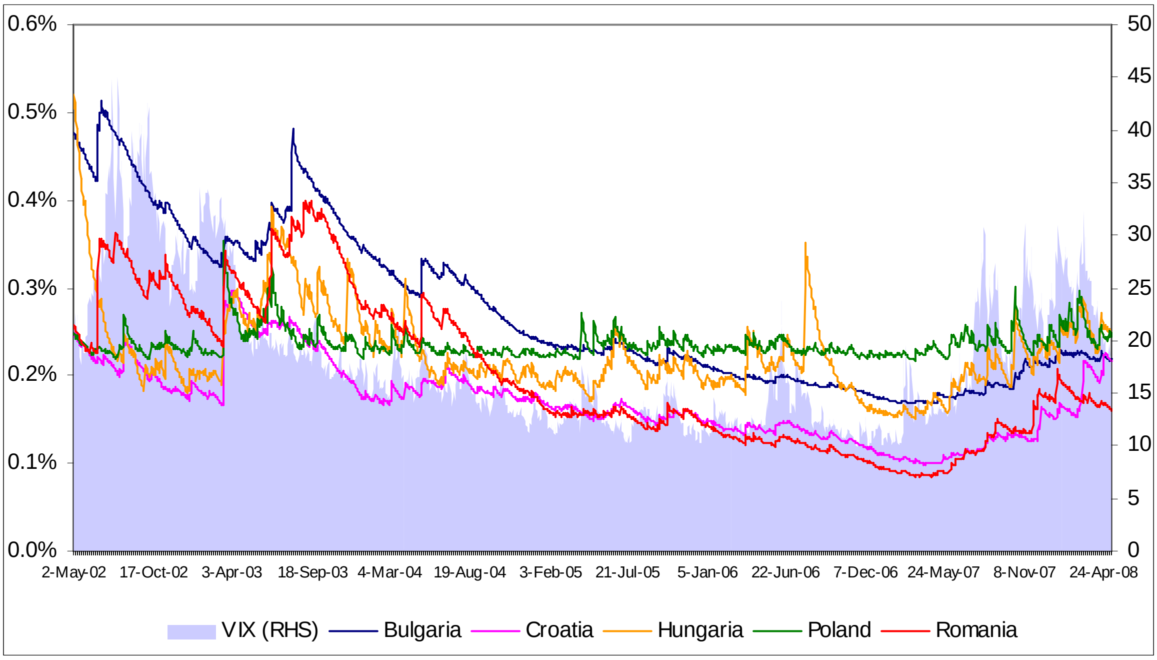
		Romania	Bulgaria	Croatia	Poland	Hungary
ARCH in Mean	c	70.198***	27.64***	86.1432***	36.4175***	29.3184***
ARCH Term	a_1	0.1938***	0.101***	0.1548**	0.0709**	0.0625***
GARCH Term	a_2	0.4147***	0.6193***	-0.0285	0.0671	0.2391*
Asymetric Term	a_3	-0.076*	n.a.	-0.1443**	-0.1098***	-0.0357
Trend Intercept	ω	0.000002***	0.000004***	0.000002***	0.000005***	0.000005***
Trend AR Term	b_1	0.9982***	0.9963***	0.9969***	0.8987***	0.9861***
Forecast error	b_2	0.0195***	0.0091**	0.0171***	0.0231***	0.0354***
LM Obs*R-squared		0.1035	1.6387	0.0329	0.2591	0.8766
$a_1 + a_2$		0.6085	0.7203	0.1263	0.138	0.3016

*, ** and *** indicate significance at the 1%, 5% and 10% confidence level, respectively

The combined coefficient for the short run component of volatility $(a_1 + a_2)$ is positive and smaller than the one for the long run volatility component (b_1) meaning that the persistence of the long run volatility is higher than for the short run component. In two cases the GARCH coefficients were not significant (for Croatia it was even negative), but even in these cases the sum of ARCH and GARCH term was still positive. This implies that the shocks to spreads price index were mostly of a long run nature. For several countries (except Bulgaria and Hungary) we found negative significant asymmetric terms.

The permanent and the transitorily component of the conditional variance for the daily returns of the five sovereign bonds price indexes are presented in the Appendix 8. As can be seen, the amplitude of the permanent component is much higher than the amplitude of the transitory component. Also, the persistence of the permanent component is very high which means that is related to the developments in underlying fundamentals variables. The Figure 13 presents the evolution of permanent components of the conditional standard deviation of daily return for the 5 European Emerging Economies. We put also the volatility index VIX. The figure shows a large degree of similarity in the permanent components of the conditional standard deviation for Bulgaria and Romania. The results are similar with the ones obtained in the pool regressions and they suggest that there was a common factor which moved the bond prices for these two countries. There is also a co-movement with the conditional standard deviation for Croatia. At the same time, we cannot say that the dynamics in the volatility index VIX display a co-movement with the estimated conditional standard deviation.

Figure13. The permanent component of the conditional standard deviation



In order to quantify the degree of comovement in the long-run and short-run components of volatility, we compute the correlation coefficients and perform also a principal component analysis. Both the correlation coefficients and the principal components analysis revel that there is important co-movement in the permanent component of the conditional volatility, while there is only a little co-movement in the short term component of the volatility. The correlation coefficient between Romania and Bulgaria is very high in the case of permanent components (0.91) and it is also high in the case of transitory components of volatility (0.53). In the case of principal components analysis for the permanent components of volatility, the first principal component accounts for 66% of the total variance when including all the five countries, for 89% of the total variance when only Romania, Bulgaria, and Croatia are included, and for 95% of the total variance when only Romania and Bulgaria are included. It is clearly that there is a strong co-movement between Romania and Bulgaria and we think that this is related to the EU accession process.

Table 4. Correlation coefficients between conditional standard deviations

Permanent components

	Bulgaria	Croatia	Hungary	Poland
Romania	0.91	0.81	0.48	0.27
Bulgaria		0.77	0.56	0.14
Croatia			0.64	0.47
Hungary				0.45

Note: All coefficients are statistically significant

Transitory components

	Bulgaria	Croatia	Hungary	Poland
Romania	0.53	0.16	0.20	-0.11
Bulgaria		0.07	0.16	0%*
Croatia			0.31	0.22
Hungary				0.24

The coefficient for Poland and Bulgaria is not statistical significant

Table 5. Results of principal components analysis for the permanent conditional standard deviations

A. All five countries are included

Eigenvalues: (Sum = 5, Average = 1)

Number	Value	Difference	Proportion	Cumulative Value	Cumulative Proportion
1	3.290407	2.295343	0.6581	3.290407	0.6581
2	0.995064	0.520360	0.1990	4.285470	0.8571
3	0.474703	0.304614	0.0949	4.760174	0.9520
4	0.170090	0.100353	0.0340	4.930263	0.9861
5	0.069737	---	0.0139	5.000000	1.0000

B. Only Romania, Bulgaria and Croatia are included

Eigenvalues: (Sum = 3, Average = 1)

Number	Value	Difference	Proportion	Cumulative Value	Cumulative Proportion
1	2.657921	2.406227	0.8860	2.657921	0.8860
2	0.251694	0.161309	0.0839	2.909615	0.9699
3	0.090385	---	0.0301	3.000000	1.0000

C. Only Romania and Bulgaria are included

Eigenvalues: (Sum = 2, Average = 1)

Number	Value	Difference	Proportion	Cumulative Value	Cumulative Proportion
1	1.905270	1.810539	0.9526	1.905270	0.9526
2	0.094730	---	0.0474	2.000000	1.0000

In the final section of the paper we test for the spillover effects among the permanent component of the volatility. We reestimate the CGARCH models by allowing the long-run component of the volatility for a country to depend on the lagged value of the permanent component of volatility of any of the other four countries. For instance, in case of Romania we estimate four models with the following structure:

$$r_{RO,t} = c \cdot \sigma^2_{RO,t} + \varepsilon_{RO,t}, \text{ with } \varepsilon_{RO,t} / I_{t-1} \sim N(0, \sigma^2_{RO;t})$$

$$\sigma^2_{RO,t} = q_t + a_1 \cdot (\varepsilon^2_{RO,t-1} - q_{RO,t-1}) + a_2 \cdot (\sigma^2_{RO,t-1} - q_{RO,t-1}) + a_3 \cdot (\varepsilon^2_{RO,t-1} - q_{RO;t-1}) \cdot D_{RO,t-1}$$

$$q_{RO,t} = \omega + b_1 \cdot (q_{RO,t-1} - \omega) + b_2 \cdot (\varepsilon^2_{RO,t-1} - \sigma^2_{RO,t-1}) + b_3 \cdot q_{j,t-1}$$

where j denotes the other four countries (Bulgaria, Croatia, Poland, and Hungary). We have similar models for the other four countries. In most of the cases, when introducing the lagged term of the permanent component of the volatility for another country, the threshold coefficient has become statistically insignificant. We are interest in the sign, the amplitude and the statistical significance of the coefficient b_3. The results of the estimations are summarized in the Table 6 for the case in which we estimate the conditional volatility of returns for Romanian bonds using the lagged values for the long-term components of the volatilities of the returns for the other four countries, and in Table 7 when the permanent component of volatility for Romania is used in the equations of long-run component of volatility for the others countries. The estimations show that there are some spillover effects both from other countries to Romania but also from Romania. The most important are the spillover effects from Poland to Romania and from Romania to Bulgaria and Croatia. There is no spillover effect from Romania to Poland.

Table 6 Spillovers effects from the permanent component of volatility of country i to the permanent component of volatility for Romania

From country i to Romania	Coefficient b_3	Standard error	z-statistics	Prob.
Bulgaria	0.020721	0.008247	2.512418	0.0120
Croatia	0.018687	0.005076	3.681276	0.0002
Hungary	0.004638	0.001308	3.545817	0.0004
Poland	0.043294	0.00949	4.561988	0.0000

Table 7 Spillovers effects from the permanent component of volatility in Romania to the permanent component of volatility for the country i

From Romania to country i i	Coefficient b_3	Standard error	z-statistics	Prob.
Bulgaria	0.180597	0.061303	2.945963	0.0032
Croatia	0.107563	0.057154	1.881976	0.0598
Hungary	0.006334	0.003187	1.987443	0.0469
Poland	0.001284	0.00155	0.828561	0.4074

V. Conclusions

There was a rapid decrease in the spreads of sovereign bonds for emerging countries in last years. For most of the Emerging Economies, the spreads reached a record low level in the 2007 summer, slightly before the US subprime crisis hit the international financial markets. For instance, EMBIG spreads for Romania decreased from 355 bp in May 2002 to only 26 bp in May 2007.

We show that the developments in the domestic fundamentals and in the risk appetite of foreign investors on the international markets explain the developments in the spreads. Using data for EMBIG spreads for Romania and other ten Emerging Economies, we find a long-run relationship between the spreads on the one hand and a Credit Rating Outlook Index (CROI) and the volatility index VIX on the other hand. The CROI is a proxy for the developments in the domestic fundamentals, while the VIX is a proxy for the risk appetite of the international investors. To estimate the long-run relationship, we use both a pool equation with fixed effects and the pooled mean group (PMG) estimator of Pesaran, Shin, and Smith (1997). The increase in the CROI index reflects a deterioration of domestic fundamentals and results in higher spreads. Higher spreads result also from an increase in the VIX index which reflects a decrease of risk appetite of investors on the global markets.

The spreads for the Romanian sovereign bonds decreased by 225 bp between May 2002 and April 2008. The estimated model, based on the long-run equilibrium relationship, implies only a decrease of 51 bp. The higher decrease in the effective spreads is due to the fact that the Romanian bonds were undervaluated in 2002 (the spreads were above their equilibrium level) and they were overvaluated in April 2008 (the spreads were below their equilibrium level). The 51 bp decreased based on the equilibrium level is due exclusively to the fundamentals (as reflected by the decrease in the S&P sovereign rating), while the external factors had no impact on the cumulated change of equilibrium level of spreads between May 2002 and April 2008. This is because following the crisis on the international markets the VIX index returned to the same level as in 2002, which means that the investors started to price appropriately the risk. There is also a practical implication from these observations: in the long run, a country cannot bet on the external factors to reduce its borrowing costs. Rather, it should implement appropriate domestic policies in order to improve domestic fundamentals.

There is a large similitude between the deviations of spreads from the level implied by the long-run relationship in the case of Bulgaria and Romania, which we explain by the EU accession process of these two countries. For instance spreads increased above their equilibrium level in 2003 when two countries failed to be nominee for the EU accession in 2004. But they moved rapidly towards the equilibrium level in 2004 when their accession was confirmed for 2004, and they decrease even below the equilibrium level in 2007 after these two countries became full members of the European Union.

We find also a comovement in the volatility of daily returns of CEE sovereign bonds, with spillover effects especially between Bulgaria and Romania. The commovement is located at the level of the permanent component of the conditional volatility, which mean that is related to underling factors.

Although the results of the analysis are plausible from an economic point of view, we think that additional research is welcomed. For instance, modeling the impact of EU accession on the spreads of CEE sovereign bonds is challenging from an econometric point of view given that this is an unobservable variable. Also, alternative estimation methods might be used in order to check the robustness of the empirical results.

VI. References

Alexander, C. (2001), “Market Models: A Guide to Financial Data Analysis”, John Wiley & Sons Ltd

Chicago Board Options Exchange (2003), VIX CBOE Volatility Index, www.cboe.com

Edwards, S. (1983), “LCD’s Foreign Borrowing and Default Risk: An Empirical Investigation 1976 – 1980”, NBER Working Paper Series No. 1172

Ferrucci, G. (2003), “Empirical Determinants of Emerging Market Economies’ Sovereign Bond Spreads”, Bank of England, Working Paper No. 205

Fiess, N (2003), “Capital Flows, Country Risk and Contagion”, World Bank Policy Research Working Paper 2943

Gonzalez Hermosillo, B. (2008), “Investors’ Risk Appetite and Global Financial Market Conditions”, IMF Working Paper 08/85

Hartelius, K., K. Kashiwase and L.E. Kodres (2008), “Emerging Market Spread Compression: Is it Real or is it Liquidity”, IMF Working Paper 08/10

Hauner, D., J. Jonas, and M.S. Kumar (2007), “Policy Credibility and Sovereign Credit: The Case of New EU Member States”, IMF Working Paper 07/1

J.P. Morgan (1999), “Introducing the J.P. Morgan Emerging Market Bond Index Global (EMBI Global), www.morganmarkets.com

Luengnaruemitchai, P. and S. Schadler (2007), “Do Economists’ and Financial Markets’ Perspectives on the New Members of the EU Differ?” , IMF Working Paper 07/65

Maier, Philipp (2008), “Sound Policies or Good Fortune: What Drives the Compression in Emerging Market Spreads?”, Paper presented at the Canadian Economics Assosiation 42nd Annual Meetings at University of British Columbia, Vancouver

McGuire, P. and M.A. Schrijvers (2003), “Common Factors in Emerging Market Spreads”, BIS Quarterly Review

Ozatay, F. E. Ozmen and G. Sahinbeyoglu (2007), “Emerging Markets Sovereign Spreads, Global Financial Conditions and US Macroeconomic News”, Economic Research Center Working Papers in Economics 07/07

Pesaran, M.H., Y. Shin and R.P. Smith (1997), “Pool estimation of Long Run Relationship in Dynamic Heterogeneous Panels”, Birkbeck College London

Pramor, M. and N.T. Tamirisa (2006), “Common Volatility Trends in the Central and Eastern European Currencies and the Euro”, IMF Working Paper 06/206

VII. Appendixes

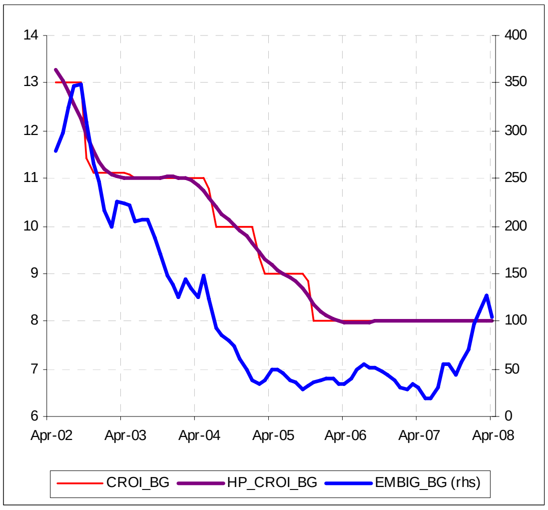
Appendix 1. The relation between the Crating Rating Outlook Index (CROI) and the S&P sovereign ratings

Category	S&P sovereign ratings	Outlook		
		Stable	Positive	Negative
Investment grade				
	AAA	1	0	2.7
	AA+	2	1	3.7
	AA	3	2	4.7
	AA-	4	3	5.7
	A+	5	4	6.7
	A	6	5	7.7
	A-	7	6	8.7
	BBB+	8	7	9.7
	BBB	9	8	10.7
	BBB-	10	9	11.7
Sub-investment grade, categoria I				
	BB+	11	10.1	12.7
	BB	12	11.1	13.7
	BB-	13	12.1	14.7
	B+	14	13.1	15.7
	B	15	14.1	16.7
	B-	16	15.1	17.7
	CCC+	17	16.1	18.7
Sub-investment grade, categoria II				
	CCC	18	18	18
	CCC-	19	19	19
	CC	20	20	20
	C	21	21	21
	SD	22	22	22

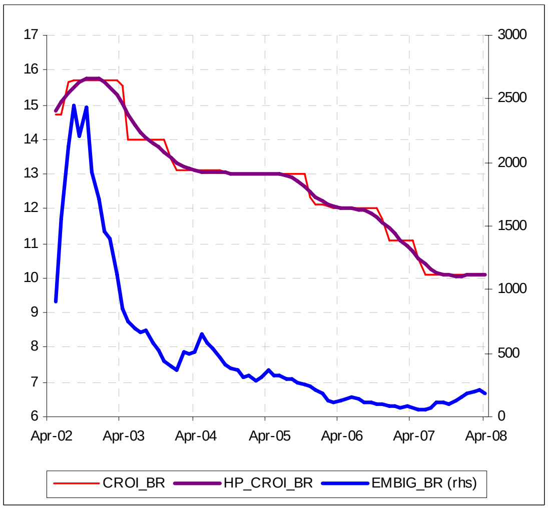
Source: Kashiwase si Kodres (2005)

Appendix 2. The relationship between the CROI index, the HP filtered CROI index and the EMBIG spreads for the countries included in analysis

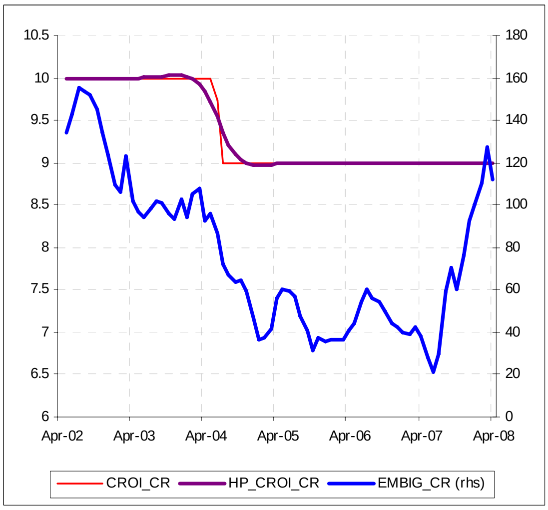
Bulgaria



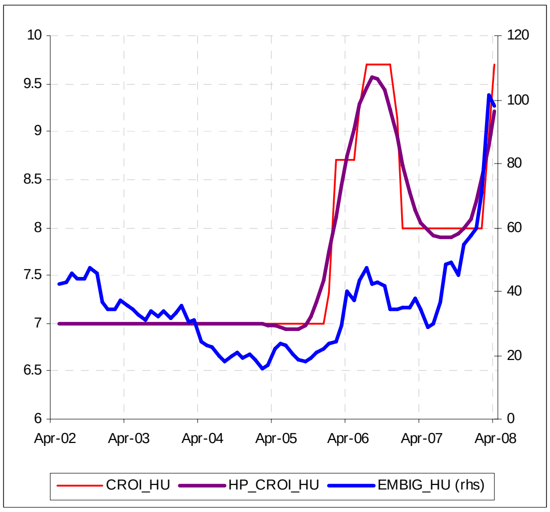
Brasil



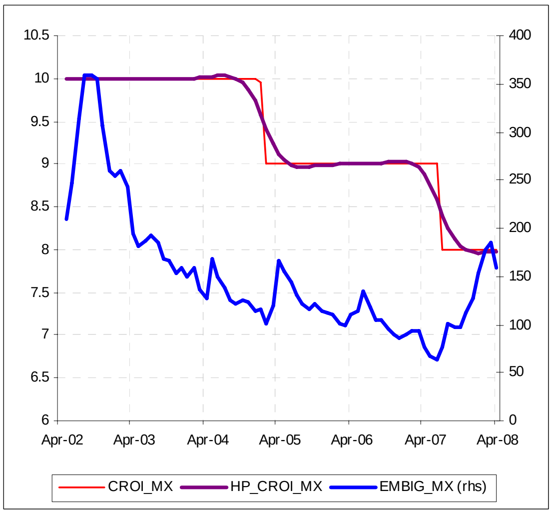
Croatia



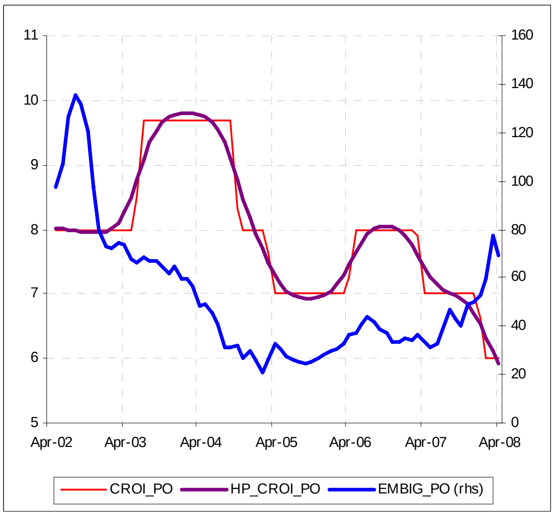
Hungary



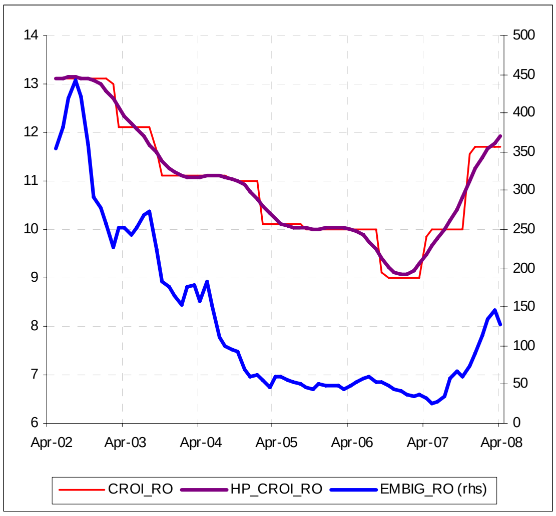
Mexic



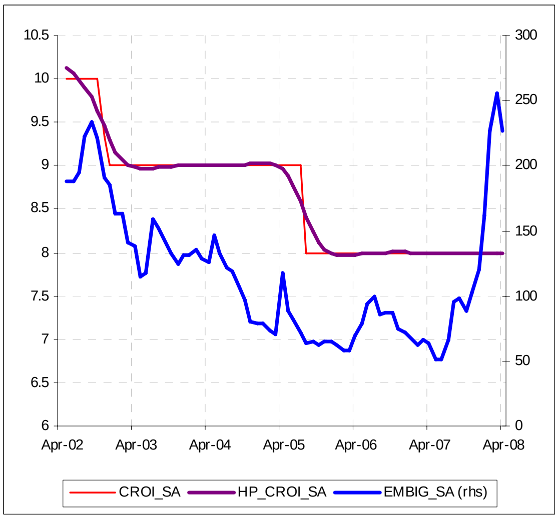
Poland



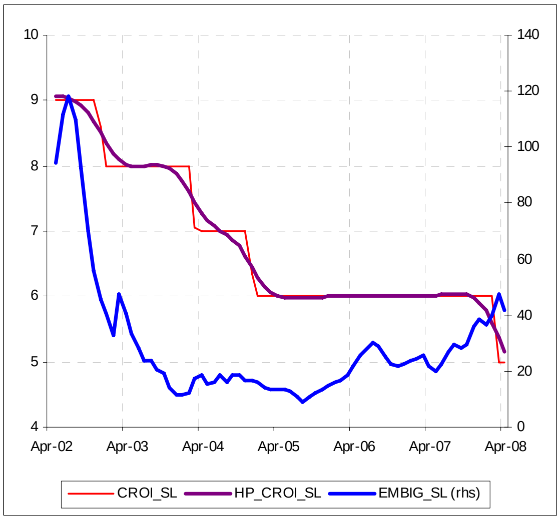
Romania



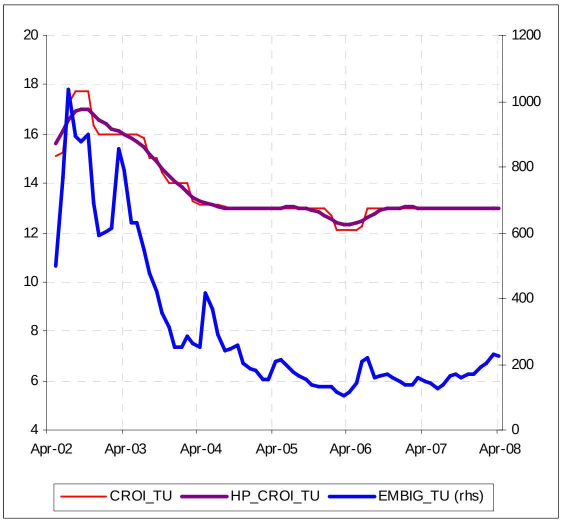
South Africa



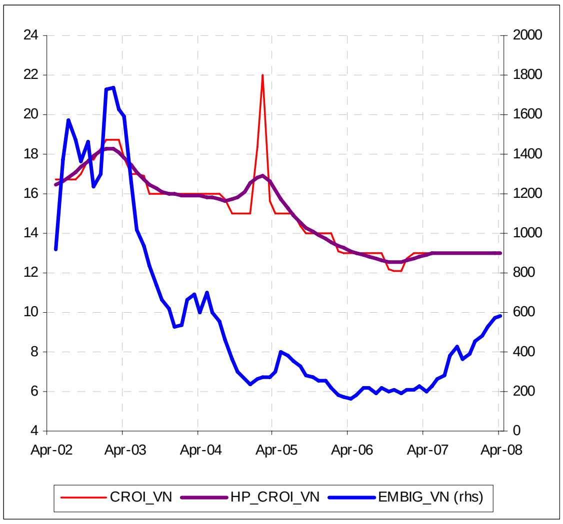
Slovakia



Turkey



Venezuela



Appendix 3. Unit root test for individuals time series

The Augmented Dickey Fuller test for a variable y_t implies the estimation of the following equation: $\Delta y_t = \phi y_{t-1} + \alpha_0 + \delta \cdot t + \sum_{i=k}^{p} \Delta y_{t-k} + \varepsilon_t$ (where $\Delta y_t = y_t - y_{t-1}$) and testing if $\phi = 0$ (which is the null hypothesis of the test and which is equivalent with unit root in the series) against the alternative of $\phi < 0$ (which implies that the series is stationary). Taking into account the patterns of the series (a decreasing trend during the whole period) we decided to use the model with the constant. The number of lags p was selected based on the Swartz Information criterion. In the case of the Phillips Perron test we also included a constant in the underliny equation $\Delta y_t = \phi y_{t-1} + \alpha_0 + \delta \cdot t + \varepsilon_t$.

The p-value in the table denote the probability associated to the null hypothesis that the series have a unit root.

Variable	Augmented Dickey-Fuller			Phillips-Perron		Result
	Lags	t-statistic	p-value	Adj. t-Stat	p-value	
log_embi_bg	2	-1.79	0.38	-1.58	0.49	I(1)
log_embi_br	0	-1.02	0.74	-1.13	0.70	I(1)
log_embi_cr	1	-2.20	0.21	-1.67	0.44	I(1)
log_embi_hu	0	0.25	0.97	-0.07	0.95	I(1)
log_embi_mx	0	-1.44	0.56	-1.57	0.49	I(1)
log_embi_po	0	-1.47	0.54	-1.53	0.51	I(1)
log_embi_ro	1	-1.84	0.36	-1.62	0.47	I(1)
log_embi_sa	0	-1.16	0.69	-1.39	0.58	I(1)
log_embi_sl	1	-2.81	0.06	-2.16	0.22	I(1)
log_embi_tu	0	-1.33	0.61	-1.37	0.59	I(1)
log_embi_vn	1	-1.83	0.36	-1.27	0.64	I(1)
log_vix	1	-1.64	0.45	-1.64	0.45	I(1)
vff1	4	-0.73	0.83	1.05	0.99	I(1)
ff3m	5	-2.70	0.08	-1.00	0.75	I(1)

Appendix 4. Panel unit root tests

There are six panel unit root tests available in Eviews. Tree of them assume that there is a common unit root process in the series (Levin, Lin and Chu test, Breitung test, and Hadri test), while the other test allow for individual unit root processes (Im, Pesaran and Shin test, Fisher-ADF test and PP test).

If we consider the AR(1) process for panel data:

$$y_{i,t} = \rho_i t_{i,t-1} + X_{i,t}\delta_i + \varepsilon_{i,t}$$

where $i = 1,2,...,N$ are cross-section units or series that are observed over periods $t = 1,2,...,T$ and $X_{i,t}$ represent the exogenous variables in the model, including any fixed effects or individual trends, ρ_i are the autoregressive coefficients and the errors $\varepsilon_{i,t}$ are assumed to be mutually independent idiosyncratic disturbance. If $|\rho_i| < 1$, y_t is said to be weakly (trend-) stationary. On the other hand, if $|\rho_i| = 1$ then y_t contains a unit root. The first three panel unit root tests employ the assumption that $\rho_i = \rho$ for all i (common unit root tests), while the last three panel unit root tests allow ρ_i to vary freely across cross-sections.

The following table summarize the null hypothesis and the alternative hypothesis for each of the six panel unit root tests.

Test	**Null Hypotesis**	**Alternative hypotesis**
Levin, Lin and Chu	Unit root	No Unit Root
Breitung	Unit root	No Unit Root
Im, Pesaran and Shin W-stat Im, Pesaran and Shin W-stat	Unit root	Some crosssections Without unit roots
Fisher-ADF	Unit root	Some crosssections Without unit roots
Fisher-PP	Unit root	Some crosssections Without unit roots
Hadri	No Unit Root	Unit root

The results of the panel unit root tests for the EMBIG spreads (in logarithm) and for the HP filtered CROI index are summarized in the following tables.

Panel unit root test: Summary
Series: LOG_EMBIG
Sample: 2002M05 2008M04
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 2
Newey-West bandwidth selection using Bartlett kernel

Cross-

Method	Statistic	Prob.**	sections	Obs
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	-1.52752	0.0633	11	775
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	0.02552	0.5102	11	775
ADF - Fisher Chi-square	19.5146	0.6134	11	775
PP - Fisher Chi-square	14.2079	0.8939	11	781

** Probabilities for Fisher tests are computed using an asymptotic Chi
-square distribution. All other tests assume asymptotic normality.

Null Hypothesis: Stationarity
Series: LOG_EMBIG
Sample: 2002M05 2008M04
Exogenous variables: Individual effects
Newey-West bandwidth selection using Bartlett kernel
Total (balanced) observations: 792
Cross-sections included: 11

Method	Statistic	Prob.**
Hadri Z-stat	13.5215	0.0000
Heteroscedastic Consistent Z-stat	11.3013	0.0000

* Note: High autocorrelation leads to severe size distortion in Hadri test,
leading to over-rejection of the null.
** Probabilities are computed assuming asympotic normality

Panel unit root test: Summary
Series: LOG_HP_CROI
Sample: 2002M05 2008M04
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 4
Newey-West bandwidth selection using Bartlett kernel
Balanced observations for each test

Method	Statistic	Prob.**	Cross-sections	Obs
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	-0.55985	0.2878	11	737
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	1.39395	0.9183	11	737
ADF - Fisher Chi-square	15.9756	0.8171	11	737
PP - Fisher Chi-square	12.1553	0.9541	11	781

** Probabilities for Fisher tests are computed using an asymptotic Chi
-square distribution. All other tests assume asymptotic normality.

Null Hypothesis: Stationarity Series: LOG_HP_CROI Sample: 2002M05 2008M04 Exogenous variables: Individual effects Newey-West bandwidth selection using Bartlett kernel Total (balanced) observations: 792 Cross-sections included: 11

Method	Statistic	Prob.**
Hadri Z-stat	16.9923	0.0000
Heteroscedastic Consistent Z-stat	16.4550	0.0000

* Note: High autocorrelation leads to severe size distortion in Hadri test, leading to over-rejection of the null.
** Probabilities are computed assuming asympotic normality

Appendix 5 Results of panel cointegration test

The panel cointegration tests shows that there is cointegration relationship between the EMBIG spreads (in logarithm), the HP filtered CROI (in logarithm) and the VIX (in logarithm).

Pedroni Residual Cointegration Test Series: LOG_EMBI_? LOG_HP_CROI_? LOG_VIX Sample: 2002M05 2008M04 Included observations: 72 Cross-sections included: 11 Null Hypothesis: No cointegration Trend assumption: No deterministic trend Lag selection: Automatic SIC with a max lag of 11 Newey-West bandwidth selection with Bartlett kernel

Alternative hypothesis: common AR coefs. (within-dimension)

	Statistic	Prob.	Weighted Statistic	Prob.
Panel v-Statistic	3.182619	0.0025	3.177313	0.0026
Panel rho-Statistic	-2.115932	0.0425	-2.482330	0.0183
Panel PP-Statistic	-2.083885	0.0455	-2.307652	0.0278
Panel ADF-Statistic	-2.664280	0.0115	-2.684834	0.0109

Alternative hypothesis: individual AR coefs. (between-dimension)

	Statistic	Prob.
Group rho-Statistic	-1.279294	0.1760
Group PP-Statistic	-1.658859	0.1008
Group ADF-Statistic	-2.208617	0.0348

Kao Residual Cointegration Test
Series: LOG_EMBI_? LOG_HP_CROI_? LOG_VIX
Sample: 2002M05 2008M04
Included observations: 72
Null Hypothesis: No cointegration
Trend assumption: No deterministic trend
Lag selection: Automatic 1 lag by SIC with a max lag of 11
Newey-West bandwidth selection using Bartlett kernel

	t-Statistic	Prob.
ADF	-3.098928	0.0010
Residual variance	0.015426	
HAC variance	0.018871	

Appendix 6 Results of panel estimations with fixed effects

Dependent Variable: LOG_EMBI_?
Method: Pooled EGLS (Cross-section SUR)
Sample: 2002M05 2008M04
Included observations: 72
Cross-sections included: 11
Total pool (balanced) observations: 792
Linear estimation after one-step weighting matrix
Cross-section SUR (PCSE) standard errors & covariance (d.f. corrected)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-4.150236	0.196415	-21.12989	0.0000
LOG(HP_CROI_?)	2.544679	0.063320	40.18774	0.0000
BG--LOG(VIX)	1.245724	0.102703	12.12937	0.0000
BR--LOG(VIX)	1.561762	0.166648	9.371598	0.0000
CR--LOG(VIX)	0.953539	0.087435	10.90571	0.0000
HU--LOG(VIX)	0.880507	0.056080	15.70092	0.0000
MX--LOG(VIX)	0.823619	0.064358	12.79742	0.0000
PO--LOG(VIX)	1.176625	0.101113	11.63678	0.0000
RO--LOG(VIX)	1.159120	0.134922	8.591051	0.0000
SA--LOG(VIX)	0.785796	0.078891	9.960475	0.0000
SL--LOG(VIX)	0.637691	0.173224	3.681314	0.0002
TU--LOG(VIX)	0.877326	0.097576	8.991228	0.0000
VN--LOG(VIX)	1.318097	0.083019	15.87706	0.0000
Fixed Effects (Cross)				
BG--C	-0.729494			
BR--C	-0.995616			
CR--C	-0.032228			
HU--C	-0.017427			
MX--C	1.080493			
PO--C	-0.617098			
RO--C	-0.636171			
SA--C	1.101355			
SL--C	0.681114			
TU--C	0.523515			
VN--C	-0.358444			

Effects Specification

Cross-section fixed (dummy variables)

Weighted Statistics

R-squared	0.965400	Mean dependent var	13.10752
Adjusted R-squared	0.964410	S.D. dependent var	11.56280
S.E. of regression	1.007703	Sum squared resid	780.8934
F-statistic	975.3025	Durbin-Watson stat	0.585253
Prob(F-statistic)	0.000000		

Unweighted Statistics

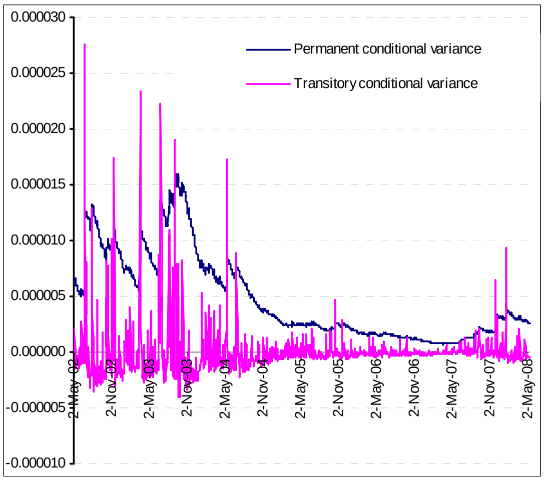
R-squared	0.918492	Mean dependent var	4.602713
Sum squared resid	76.31891	Durbin-Watson stat	0.221973

Appendix 7 Estimation results in the case of the pooled group mean (PGM) estimator

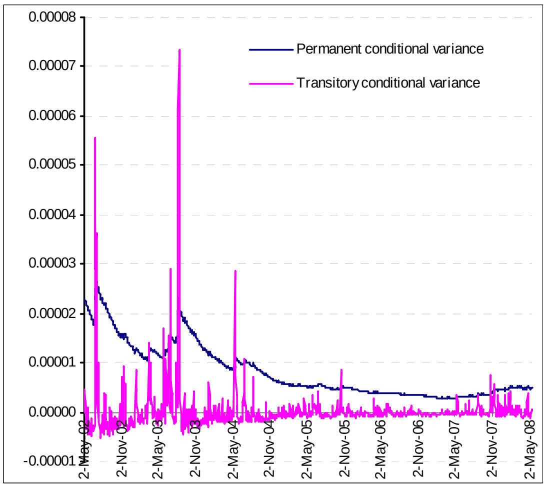
Country	ARDL Model		Phi	log_hp_croi	log_vix	SIGMA	CH-SC	CH-FF	CH-NO	CH-HE	RBARSQ	LL	AIC	SC
Bulgaria	2 0 1	Coefficient	-0.17	2.53	1.43	0.12	11.52	0.79	6.86	0.03	0.51	50.95	44.95	38.20
		Std. Error	0.05	0.22	0.27									
		t-statistic	-3.43	11.23	0.05									
Brasilia	1 1 1	Coefficient	-0.09	2.53	1.69	0.12	1.32	5.84	14.50	0.00	0.53	53.80	47.80	41.01
		Std. Error	0.03	0.22	0.51									
		t-statistic	-2.88	11.23	3.34									
Croatia	2 0 0	Coefficient	-0.29	2.53	1.08	0.12	3.87	4.30	6.24	2.86	0.40	51.72	46.72	41.10
		Std. Error	0.05	0.22	0.15									
		t-statistic	-5.97	11.23	7.05									
Hungary	2 0 0	Coefficient	-0.38	2.53	0.92	0.10	2.13	1.08	0.73	1.11	0.32	65.61	60.61	54.98
		Std. Error	0.07	0.22	0.09									
		t-statistic	-5.70	11.23	9.92									
Mexic	1 2 1	Coefficient	-0.17	2.53	0.89	0.09	0.89	0.11	0.79	0.02	0.47	76.67	69.67	61.80
		Std. Error	0.07	0.22	0.19									
		t-statistic	-2.51	11.23	4.62									
Poland	1 1 1	Coefficient	0.04	2.53	1.22	0.09	1.68	0.73	0.94	0.67	0.24	70.64	64.64	57.85
		Std. Error	0.04	0.22	0.76									
		t-statistic	1.00	11.23	1.60									
Romania	2 0 1	Coefficient	-0.12	2.53	1.49	0.11	0.89	3.53	0.17	0.35	0.44	61.00	55.00	48.25
		Std. Error	0.03	0.22	0.33									
		t-statistic	-3.63	11.23	4.48									
South Africa	1 1 0	Coefficient	-0.24	2.53	1.22	0.13	6.40	2.59	23.61	0.23	0.23	49.45	44.45	38.79
		Std. Error	0.06	0.22	0.22									
		t-statistic	-3.87	11.23	5.44									
Slovakia	1 0 0	Coefficient	0.00	2.53	12.08	0.14	6.62	0.99	2.95	6.09	-0.03	38.85	34.85	30.32
		Std. Error	0.04	0.22	86.80									
		t-statistic	0.13	11.23	0.14									
Turkey	1 2 1	Coefficient	-0.13	2.53	1.67	0.12	0.23	0.16	43.26	0.92	0.38	53.54	46.54	38.67
		Std. Error	0.05	0.22	0.51									
		t-statistic	-2.47	11.23	3.29									
Venezuela	2 0 0	Coefficient	-0.24	2.53	1.41	0.10	0.54	0.55	3.46	3.87	0.34	61.76	56.76	51.14
		Std. Error	0.04	0.22	0.16									
		t-statistic	-5.47	11.23	8.81									

Apendix 8. The permanent and the transitory components of the conditional variance

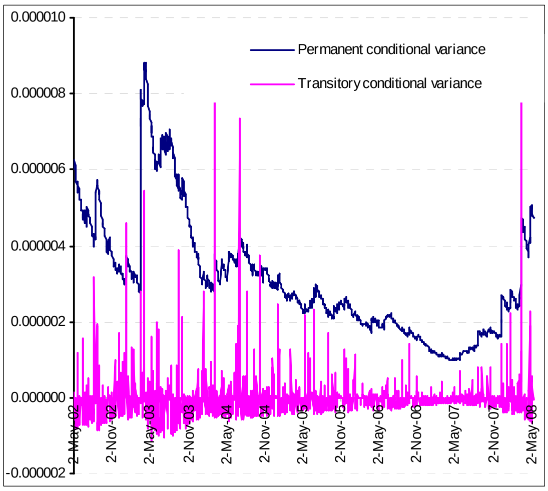
Romania



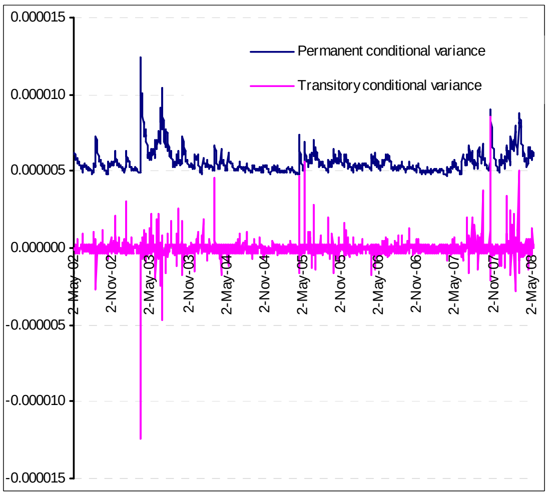
Bulgaria



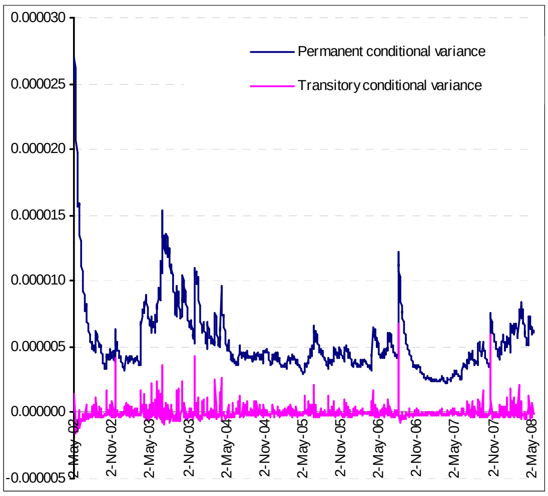
Croatia



Poland



Hungary



Appendix 9. Estimations results fot the CGARCH models

Dependent Variable: DLOG_EMBI_RO
Method: ML - ARCH (Marquardt) - Normal distribution
Sample (adjusted): 5/02/2002 5/13/2008
Included observations: 1574 after adjustments
Convergence achieved after 14 iterations
Presample variance: backcast (parameter = 0.7)
Q = C(2) + C(3)*(Q(-1) - C(2)) + C(4)*(RESID(-1)^2 - GARCH(-1))
GARCH = Q + (C(5) + C(6)*(RESID(-1)<0))*(RESID(-1)^2 - Q(-1)) + C(7)
*(GARCH(-1) - Q(-1))

	Coefficient	Std. Error	z-Statistic	Prob.
GARCH	70.19783	11.41794	6.148029	0.0000
	Variance Equation			
C(2)	2.22E-06	8.46E-07	2.624179	0.0087
C(3)	0.998218	0.000807	1236.752	0.0000
C(4)	0.019599	0.003522	5.565282	0.0000
C(5)	0.193867	0.038752	5.002799	0.0000
C(6)	-0.075663	0.042046	-1.799524	0.0719
C(7)	0.414703	0.096587	4.293546	0.0000

R-squared	-0.005687	Mean dependent var	0.000315
Adjusted R-squared	-0.009537	S.D. dependent var	0.002205
S.E. of regression	0.002215	Akaike info criterion	-9.802840
Sum squared resid	0.007691	Schwarz criterion	-9.778996
Log likelihood	7721.835	Hannan-Quinn criter.	-9.793978
Durbin-Watson stat	1.821409		

Dependent Variable: DLOG_EMBI_BG
Method: ML - ARCH (Marquardt) - Generalized error distribution (GED)
Sample (adjusted): 5/02/2002 5/13/2008
Included observations: 1574 after adjustments
Convergence achieved after 27 iterations
Presample variance: backcast (parameter = 0.7)
Q = C(2) + C(3)*(Q(-1) - C(2)) + C(4)*(RESID(-1)^2 - GARCH(-1))
GARCH = Q + C(5) * (RESID(-1)^2 - Q(-1)) + C(6)*(GARCH(-1) - Q(-1))

	Coefficient	Std. Error	z-Statistic	Prob.
GARCH	27.64016	6.954640	3.974348	0.0001
	Variance Equation			
C(2)	4.49E-06	1.17E-06	3.836925	0.0001
C(3)	0.996363	0.001264	788.4463	0.0000

C(4)	0.009154	0.004177	2.191422	0.0284
C(5)	0.101071	0.031480	3.210629	0.0013
C(6)	0.619328	0.127271	4.866204	0.0000
GED PARAMETER	1.080299	0.047544	22.72196	0.0000

R-squared	0.005347	Mean dependent var	0.000330
Adjusted R-squared	0.001539	S.D. dependent var	0.002937
S.E. of regression	0.002935	Akaike info criterion	-9.224668
Sum squared resid	0.013498	Schwarz criterion	-9.200825
Log likelihood	7266.814	Hannan-Quinn criter.	-9.215807
Durbin-Watson stat	2.049654		

Dependent Variable: DLOG_EMBI_CR
Method: ML - ARCH (Marquardt) - Generalized error distribution (GED)
Sample (adjusted): 5/02/2002 5/13/2008
Included observations: 1574 after adjustments
Convergence achieved after 15 iterations
Presample variance: backcast (parameter = 0.7)
Q = C(2) + C(3)*(Q(-1) - C(2)) + C(4)*(RESID(-1)^2 - GARCH(-1))
GARCH = Q + (C(5) + C(6)*(RESID(-1)<0))*(RESID(-1)^2 - Q(-1)) + C(7)
*(GARCH(-1) - Q(-1))

	Coefficient	Std. Error	z-Statistic	Prob.
GARCH	86.14324	12.73289	6.765414	0.0000
Variance Equation				
C(2)	2.51E-06	7.52E-07	3.334042	0.0009
C(3)	0.996935	0.001825	546.2710	0.0000
C(4)	0.017184	0.003658	4.697541	0.0000
C(5)	0.154821	0.066511	2.327746	0.0199
C(6)	-0.144328	0.070971	-2.033615	0.0420
C(7)	-0.028356	0.265797	-0.106684	0.9150
GED PARAMETER	1.233574	0.039660	31.10353	0.0000

R-squared	0.000203	Mean dependent var	0.000232
Adjusted R-squared	-0.004266	S.D. dependent var	0.001775
S.E. of regression	0.001779	Akaike info criterion	-10.02387
Sum squared resid	0.004956	Schwarz criterion	-9.996621
Log likelihood	7896.786	Hannan-Quinn criter.	-10.01374
Durbin-Watson stat	2.057967		

Dependent Variable: DLOG_EMBI_HU
Method: ML - ARCH (Marquardt) - Normal distribution
Sample (adjusted): 5/02/2002 5/13/2008
Included observations: 1574 after adjustments

Convergence achieved after 10 iterations
Presample variance: backcast (parameter = 0.7)
Q = C(2) + C(3)*(Q(-1) - C(2)) + C(4)*(RESID(-1)^2 - GARCH(-1))
GARCH = Q + (C(5) + C(6)*(RESID(-1)<0))*(RESID(-1)^2 - Q(-1)) + C(7)
*(GARCH(-1) - Q(-1))

	Coefficient	Std. Error	z-Statistic	Prob.
GARCH	29.31846	10.89303	2.691489	0.0071
	Variance Equation			
C(2)	5.13E-06	5.37E-07	9.562896	0.0000
C(3)	0.986029	0.004346	226.8733	0.0000
C(4)	0.035450	0.006268	5.655464	0.0000
C(5)	0.062520	0.018607	3.359945	0.0008
C(6)	-0.035680	0.026762	-1.333243	0.1825
C(7)	0.239013	0.344376	0.694047	0.4877

R-squared	-0.001838	Mean dependent var	0.000175
Adjusted R-squared	-0.005674	S.D. dependent var	0.002251
S.E. of regression	0.002258	Akaike info criterion	-9.404156
Sum squared resid	0.007987	Schwarz criterion	-9.380313
Log likelihood	7408.071	Hannan-Quinn criter.	-9.395295
Durbin-Watson stat	2.096016		

Dependent Variable: DLOG_EMBI_PO
Method: ML - ARCH (Marquardt) - Normal distribution
Sample (adjusted): 5/02/2002 5/13/2008
Included observations: 1574 after adjustments
Convergence achieved after 12 iterations
Presample variance: backcast (parameter = 0.7)
Q = C(2) + C(3)*(Q(-1) - C(2)) + C(4)*(RESID(-1)^2 - GARCH(-1))
GARCH = Q + (C(5) + C(6)*(RESID(-1)<0))*(RESID(-1)^2 - Q(-1)) + C(7)
*(GARCH(-1) - Q(-1))

	Coefficient	Std. Error	z-Statistic	Prob.
GARCH	36.41753	10.81025	3.368796	0.0008
	Variance Equation			
C(2)	5.59E-06	1.65E-07	33.83260	0.0000
C(3)	0.898771	0.028903	31.09653	0.0000
C(4)	0.023155	0.001934	11.97082	0.0000
C(5)	0.070944	0.031017	2.287239	0.0222
C(6)	-0.109856	0.035939	-3.056761	0.0022
C(7)	0.067186	0.385690	0.174198	0.8617

R-squared	-0.001175	Mean dependent var	0.000211
Adjusted R-squared	-0.005008	S.D. dependent var	0.002375
S.E. of regression	0.002381	Akaike info criterion	-9.265266
Sum squared resid	0.008886	Schwarz criterion	-9.241423

Log likelihood	7298.765	Hannan-Quinn criter.	-9.256405
Durbin-Watson stat	2.059645		