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Excessive Debt or Excess Savings -- Transition Countries Sovereign Bond Spread Assessment

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Abstract

We study the sovereign yield spreads determinants in transition – Central and Eastern Europe (CEE) and Caucasus and Central Asia (CCA) -- countries and try to provide an answer to the key question: was the narrowing of the spreads and their compression a result of improvement of CEECCA countries sovereign's macroeconomic policy (implemented in early to mid 2000s), or was it due to global excess liquidity provision? If better domestic macroeconomic policy efforts and solid reforms implemented in this period have led to: i) improvement in sovereign debt management e.g., by increasing the average debt portfolio duration and reducing the stock of FOREX debt; ii) development of domestic financial markets with enlargement of the investor's base and enhancement of the risk management techniques; iii) continuing financial liberalization; iv) sustainable fiscal adjustment, reserve accumulation and price stability; and v) adoption of the most conductive to prosperity institutional structure, then it would be expected that any tighter monetary policy environment in the developed economies should have only a tiny effect on spreads.

The models are estimated on an individual basis -- country by country -- using a framework allowing for fractionally integrated variables (ARDL) as well as, by utilising panel data (cross-sectional-time-series) estimation whenever data availability allows.

We utilise daily data over the period 2006-2012 and quarterly data over the period 2002-2011. These are the periods for which meaningful comparable data are available for Bulgaria, Croatia, Hungary, Kazakhstan, Poland, Russia, Serbia, and Ukraine (in various combinations).

We are careful not to attempt to split the sample into (say two) potential segments for comparison of "normal" versus "crises" period estimates (as customary) as since 2002 / 2003 the transition economies have started to experience the powerful financial effect generated by the excess global liquidity, i.e., the entire period under consideration is constituted by two phases characterised by: i) excess liquidity (2002-2008); and, ii) the Great Depression Mark II (2008 – to present).

Keywords: sovereign yield spreads determinants, government debt and risk management, transition markets bond indices, credit default swaps, financial market volatility, transition economies

JEL Classification Numbers: C22, C23, E44, G12

1. Introduction

"Half-knowledge is more victorious than whole knowledge: it understands things as being more simple than they are and this renders its opinions more easily intelligible and more convincing."

Nietzsche: Human, All Too Human: A Book for Free Spirits

A range of academic studies have analysed the determinants of the difference between the sovereign's emerging market debt securities and US Treasury bonds and/or German bunds of similar maturities. Still, while there have been a number of papers dealing with yield spreads on Eurozone government bonds (e.g., Codogno, Favero and Missale (2003), Pagano and Von Thadden (2004), Mody (2009), and Klepsch and Wollmershauser, (2011)) there have not been many methodical studies on the price determination of sovereign bonds in emerging markets; particularly in the group of Central and Eastern Europe (CEE) and Caucasus and Central Asia (CCA) countries.

One early (partial exception) is the paper of Eichengreen and Mody (1998) examining launch spreads based on data for a mixed group of 55 emerging market countries over the period 1991 to 1996. They collect information on altogether 1,033 bonds split as follows: 670 from Latin America; 233 from East Asia; and 81 from Eastern Europe. Regressing spreads on various potential determinants they detect: "But the same explanatory variables have different effects in the principal debt issuing regions (Latin America, East Asia, and Eastern Europe)."

It is interesting to compare the coefficients of regression on the variables Debt/GNP and GDP growth between the combined group of Latin America and East Asia countries with the Eastern Europe bond issues. While for the former group the coefficient on Debt/GNP is relatively small, has positive sign (0.437) and is significant (t-stat 2.054), for Eastern Europe its value is big, negative (-1.255) and it is insignificant (t-stat -1.367). In the same vein the coefficient on GDP growth for Latin America and East Asia is positive sizable (2.253) though insignificant (t-stat 0.616) and the equivalent coefficient for Eastern Europe is negative, vast (-14.250) and significant (t-stat -1.954). Furthermore, the coefficient of mutual determination corrected for degrees of freedom for the Latin America and East Asia estimated model is close to 0.6, while it is only about 0.09 for Eastern Europe. These OLS results suggest that about 60 per cent of the variation in spreads is explained for Latin America and East Asia and just about 9 per cent for Eastern Europe, anticipating the authors' statement: "And when it comes to changes in spreads over time, we find that these are explained mainly by shifts in market sentiment rather than by shifts in fundamentals."

Hence, the established state of knowledge in this area is as yet by no means sufficient to resolve the question of what are the major determinants of sovereign bond spreads. Our research paper aims to help to reveal definite empirical regularities, plausible interconnections, and credible causalities in this area, providing an answer to the question -- was the general narrowing of the spreads and their compression a result of an improvement of CEECCA countries macroeconomic policy, implemented after 2002, or was it due to global excess liquidity provision.

2. Literature Review

The empirical research on the determinants of government bonds spreads in advanced economies is vast, whilst the existence of similar analytical papers dealing with the emerging markets economies is more restricted. Still, both have recently enlarged, in particular since the beginning of the financial and economic crisis -- the Great Depression Mark II -- from 2008.

The main focus is: macroeconomic fundamentals determining sovereign risk; external shocks related to global liquidity; risk aversion / appetite; state of development of domestic financial markets; and, quality of governance indicators.

Contributions about the influence of macroeconomic variables on sovereign spreads, include Min (1998), Eichengreen and Mody (1998), Kamin and von Kleist (1999), and Hilscher and Nosbusch (2010). In general, these studies find considerable association with macroeconomic fundamentals and evidence that sovereign spreads in the 1990s declined more than country fundamentals' changes could account for. Baek et al (2005), among others, offer a possible explanation: they "[p]ostulate that the market-assessed country risk premium is determined not only by economic fundamentals of a sovereign but also by non-country-specific factors, especially the market's attitude towards risk." In their analysis they find that the yield spreads, although reacting to alterations in economic aggregates, in principal are driven by changes in the market perception of risk. This finding is supported by the conclusions of the studies of various authors including: McGuire and Schrijvers (2003), Jaramillo and Weber (2013), Arora and Cerisola (2001), Ferrucci (2003), and Baldacci and Kumar (2010).

Arezki and Bruckner (2010), construct an individual international commodity price index per country that allows them to confine revenue windfalls from rising prices of exported commodities and in addition exploit two measures of political institutions. Their main findings are: i) "[p]ositive international commodity price shocks lead on average to a significant reduction in commodity exporting countries' spread on sovereign bonds."; ii) allowing for cross-country differences in political institutions entails that for democracies "[a] positive commodity price shock of size 1 standard deviation significantly reduced the spread on sovereign bonds by over 0.4 standard deviation. On the other hand [...] autocracies a shock of similar magnitude was associated with a significant increase in the spread on sovereign bonds by 0.3 standard deviations."; and, iii) "[i]n democracies [...] windfalls from international commodity price shocks were significantly positively associated with real per capita GDP growth, in autocracies they were associated with a significant decrease in real per capita GDP."

Hartelius, Kashiwase and Kodres (2008), and Gonzalez-Rozada and Levy-Yeyati (2008) find that macroeconomic fundamentals, global market liquidity and risk sensitivity mutually comprise the key causes of sovereign spread changes. Similar conclusions are established by Favero, Pagano and Von Thadden (2008), who analysed the sovereign spreads of European Union countries. Mody (2009) examines the interrelations linking sovereign bond spreads in the euro area countries and financial exposure and finds that financial exposure (calculated as a ratio of an equity index for the relevant country's financial sector to the equity index taken as a whole) is strongly correlated with spread changes.

Dell'Erba and Sola (2011) – estimate the effect of the monetary and fiscal policy stance on both long-term interest rates and sovereign spreads by constructing a semi-annual dataset of macroeconomic and fiscal forecasts for 17 OECD countries over the period 1989-2009. They find that more than 60% of the variance in the data can be accounted for by monetary and fiscal policy positions.

Kaminsky, Reinhart and Vegh (2005) examine the important question of procyclical versus countercyclical capital flows and monetary and fiscal policies depending on the country's level of economic development. Their major findings are: "While macroeconomic policies in OECD countries seem to be aimed mostly at stabilizing the business cycle (or, at the very least, remaining neutral), macroeconomic policies in developing countries seem mostly to reinforce the business cycle, turning sunny days into scorching infernos and rainy days into torrential downpours."

What's more, fiscal policies are incorporated as powerful forces of sovereign spread determination in European Union countries by Bernoth, Von Hagen and Schuknecht (2004); Afonso and Strauch (2004); and, Hallerberg and Wolff (2006). Hallerberg and Wolff (2006) after controlling for institutional changes, conclude that fiscal policy remains a significant determinant of the risk premium. According to them deficits and surpluses matter less for the risk premium in countries with better institutions. Apparently this reflects the market view that proper institutions will be able to deal with fiscal problems and make the monitoring of annual developments less important. The results are robust to controlling for country fixed effects and different estimation methodologies.

Maltriz (2012), embark upon the subject-matter with Bayesian Model Averaging (BMA). In his study the author applies BMA "[t]o identify the best models and assess the quality of potential regressors." They "[f]ind that the most important drivers of default risk in the Eurozone are government debt to GDP, budget balance to GDP and terms of trade. For economic growth, export growth, import growth and the US interest rate the likelihood is between 10 and 50%, whereas for some variables found to be significant in the literature, as interest rate costs, capital formation and inflation, this likelihood is below 10%."

Gibson, Hall, and Tavlas (2011), concentrate on a single country – Greece – and macroeconomic variables shaping spreads, providing evidence that "both undershooting and overshooting of spreads have occurred." This analysis is confirmed and extended additionally in space, time, and causality by De Grauwe and Ji (2012) who "[f]ind evidence that a significant part of the surge in the spreads of the PIGS countries (Portugal, Ireland, Greece and Spain) in the eurozone during 2010-11 was disconnected from underlying increases in the debt-to-GDP ratios and fiscal space variables, but rather was the result of negative self-fulfilling market sentiments [...]." They suppose that given the state of affairs: liquidity crisis, imposed austerity measures (presumably leading the country to recession), plus high interest rates on government securities could result in a solvency crisis. According to their model investors try to factor in the costs and benefits to the government from defaulting. "A major insight of the model is that the benefit of a default depends on whether this default is expected or not." If investors expect a default, a default would occur, if they do not, no such would take place. Furthermore, they consider that if a country is not a member of the Eurozone, "This makes it possible for the country to always avoid outright default because the central bank can be forced to provide all the liquidity that is necessary to avoid such an outcome."

While this argument may add up within its settings, one should not forget that investors may lose their confidence in the ability of the government of the "stand-alone country" to sustain its currency and take flight to safety by promptly exchanging the domestic currency denominated debt for cash – Euro or/and USD. Thus the self-fulfilling prophecy (or speculative crisis) may well become true – the country would rapidly lose foreign reserves; in time it would have no choice but to devaluate its currency; the level of the external debt would increase in local currency units; this would lead eventually to monetisation of the debt; this state of affairs brings forth new speculative attacks. Hence, just being a "stand-alone country" is not likely to be sufficient to insulate you from self-fulfilling expectations or speculative attacks.

Akitoby and Stratmann (2006) emphasises the importance of sustainable fiscal policy and high fiscal adjustment, where reduction in current expenditures proves to be more effective on spread reduction than tax increases. The shaping power of liberalisation of the capital account, the currency convertibility risk premium, and the rule of law are investigated by Bacha, Holland and Goncalves (2008) as determinants of the local interest rates of emerging economies. Whereas, Edwards (2005), by means of the bidirectional interrelation between interest rates and capital account liberalisation shows that the degree of convergence of domestic and international interest rates could be used to assess the real degree of openness of the capital account.

A connected subject matter that has received considerable attention is the relationship between sovereign spreads and default risk. Favero and Missale (2011) "[f]ind that default risk is the main driver of yield spreads, suggesting small gains from greater liquidity. Fiscal fundamentals matter in the pricing of default risk but only as they interact with other countries' yield spreads; that is, with the global risk that the market perceives. More importantly, the impact of this global risk variable is not constant over time, a clear sign of contagion driven by shifts in market."

Hilscher and Nosbusch (2010), investigate spread determinants by focusing on the volatility of fundamentals. They observe "[t]hat the volatility of the terms of trade is both statistically and economically significant in explaining spread variation. A one standard deviation increase in the volatility of terms of trade is associated with an increase of 164 basis points in spreads, which corresponds to around half of the standard deviation of observed spreads." The authors assert as well that the terms of trade volatility is a significant predictor of country default. However, an important restriction of their conclusions is the regional and economic divergence of the countries included in their sample (Latin America 12, Africa 5, Eastern Europe 6, and Middle East and Asia 9) for which (time-invariant factors) no controls are provided.

Another important area of research is the detection of short-term and long-term factors determining the sovereign bond spreads. Bellas, Papaioannou, and Petrova (2010) results indicate that in the long run, fundamentals are considerable determinants of emerging market sovereign bond spreads, while in the short run, financial volatility is rather the substantial determinant of spreads. Furthermore, researchers have also distinguished between the determinants of sovereign bond spreads during normal and crisis periods. Ebner (2009) highlights a noteworthy distinction in government bond spreads in Central and Eastern Europe throughout crisis and non-crisis periods. He provides evidence that market volatility, political instability and global causes gain in importance and predominantly explain the increase in spreads during crisis periods, while macroeconomic aggregates become less important.

Belhochine and Dell'Erba (2013), applying spread regression to a panel of 26 emerging economies (including 7 transition economies: Bulgaria, Hungary, Kazakhstan, Poland, Russia, Serbia, and Ukraine) and bringing in the difference between the debt stabilising primary balance and the factual primary balance as a measure of debt sustainability, they find "[t]hat debt sustainability is a major determinant of spreads with an elasticity of about 25 basis points for each 1 percentage point departure of the primary balance from its debt stabilizing level." Furthermore they claim "[t]hat the sensitivity of spreads to debt sustainability doubles as public debt increases above 45 percent of GDP."

In addition, another related approach in the literature deals with the interrelations between debt levels and their impact on economic growth (trough implicit transmission mechanisms) within the framework of a threshold model, where the behaviour of the variables is expected to change distinctly, when certain – threshold – levels are reached. The most influential paper in this respect has been (until very recently) the one published by Reinhart and Rogoff in 2010 (Growth in a Time of Debt). There the authors claim to have identified a key stylized fact: a burden of public debt larger than ninety percent of GDP notably and consistently reduces GDP growth. Examining public debt and GDP growth among twenty advanced economies in the period after the second world war, they determine that the average real GDP growth rate for countries having a public-debt-to-GDP ratio of over ninety per cent is, in fact, negative, amounting to -0.1 per cent.

However, Herndon Th., M. Ash and R. Pollin (2013) have replicated Reinhart and Rogoff (2010) and were able to establish that coding errors, biased exclusion of available data, and unconventional weighting of summary statistics have led to miscalculations that provide a misleading picture of the relationship between public debt and GDP growth. They reveal that when accurately calculated, the annual average real GDP growth for national economies with a public-debt-to-GDP ratio of over ninety per cent is actually 2.2 percent, not -0.1 percent as stated in Reinhart and Rogoff. That is to say, that average GDP growth, when public debt/GDP ratios are in excess of ninety per cent is not significantly different from the average GDP growth when debt/GDP ratios are lower.

Consequently, the conventional state of knowledge in this area is not adequate to resolve the question: was the general narrowing of the spreads and their compression in the CEECCA countries a result of these countries enhanced macroeconomic policies, implemented after 2002, or was it due to global excess liquidity provision (excess savings / underinvestment in real capital).

3. Methodology

In aiming to provide an answer (and illustrative evidence) to the above question we estimate various models: i) an individual basis model -- country by country -- using a framework allowing for fractionally integrated variables (ARDL); and, ii) a panel data model (cross-sectional-time-series) estimation.

We utilise daily data over the period 2006-2012 and quarterly data over the period 2002-2011. These are the periods for which meaningful comparable data -- for Bulgaria, Croatia, Hungary, Kazakhstan, Poland, Russia, Serbia, and Ukraine -- are available.

We start with the following equation with daily sampling frequencies:

eq. 1 $$\boldsymbol{SSEMBI_t = \alpha + \beta VIX_t + \gamma CDS_t + \varepsilon_t}$$

Where:

SSEMBI--Stripped Spread JPM EMBI GLOBAL

VIX – Volatility Index (proxy for global risk aversion)

CDS – Credit Default Swap (perceived individual country risk)

Initially we estimate the model on an individual country by country basis and then we move to panel data (cross-sectional-time-series) estimation. Our motivation for using a framework allowing for fractionally integrated variables (ARDL) is based on various important factors, including:

- The conventional (dichotomous) choice between unit root I(1) and level stationarity I(0) is overly restrictive many economic time series show signs of being neither I(0) nor I(1);
- Much more general and flexible apparatus than the traditional approach;
- Important for modelling a wide range of macroeconomic relationships;
- The standard practice of taking first differences may still lead to series with a component of long memory behaviour

Many researchers are accustomed to think in terms of the stationarity of any time series used in the construction of whichever econometric model is being developed. As the assumption of stationarity is an important one, non-stationary time series are commonly transformed to stationary ones by differencing. This would suggest that a model specified in differences of economic time series should be favoured for finding estimates of parameters. But one of the important notions in macroeconomics is the concept of the existence of a long-run equilibrium relationship. Theoretically in steady-state equilibrium economic variables remain unchanged, until the system is shocked. Therefore, if such an equilibrium relationship is specified in first differences, the steady-state differences would be zero and there is no solution.

Hence, in what follows we apply the (Autoregressive Distributed Lag) ARDL procedure developed by Pesaran and Shin (1995).

4. Data Availability and Data Integrity

Using data from transition economies necessitate careful discussion of its quality and consistency. These data may sometimes be characterised from pointless, through distorted, to completely inaccurate. Statistical and book-keeping standards under the socialist economic system have been very different from those commonly accepted in Western Europe. It has taken time to learn and understand it and to switch to the accepted international statistical standards.

Much of the necessary fundamental data are still to be composed and / or disclosed and made easily available to the public. We hope to provide an impetus to serious data collection and complete disclosure for all transition economies for enabling deep economic analysis and informing consistent policy-making. The situation on the statistical front is made even more complex by the supranational economic institutions (e.g., IMF and WB) practice not to distribute all the data they have (see Annex 1) and to avoid publishing the data they hand out in high frequencies1 (quarterly and monthly). Moreover, the data published in the International Financial Statistics

[1]The data frequency used may have potentially significant effects on empirical results. Of course there are pros and cons – if low frequency data is used it may not be able to grasp the dynamic changes/variability in the data generation process, whereas if daily or weekly data is analysed, it may lead to an incorrect association of bond spreads and CDS observations, particularly at a time when market activity is low and trades take place infrequently.

(IFS) and the World Economic Outlook (WEO) formats may and do differ, with access to the full database available only to internal IMF staff.

Tables 1 to 3 including (below) illustrate the data availability for the group of countries we examine.

Table 1. Macroeconomic aggregates, Quarterly – Data Availability

	(1) Gross Domestic Product, Nominal, National currency, millions	(2) Gross Domestic Product, Deflator, Index, 2010=100	(3) Consumer Prices, All items, Index, 2010=100	(4) Consumer Prices, All items; Percent Change over Corresponding Period of Previous Year	(5) General Govt. Public Sector Debt (PSDGG)	(6) Cash surplus/deficit, Cash (Budg. Cen. Govt.) National Currency Millions	(7) Saving, Gross, Nominal, National Currency, millions	(8) Import Prices, All Commodities, Index 2010=100	(9) Export Prices, All Commodities, Index 2010=100	(10) National Currency per U.S. Dollar, period average	(11) Gross capital outflows (portfolio investment assets+ direct investment assets+ other investment	(12) Official reserves
1 Albania	2009Q1-13Q3	--	2000Q1-14Q3	--	2009Q3-14Q2	--	--	--	--	2000Q1-14Q3	--	--
2 Armenia	2000Q1-14Q2	--	2000Q1-14Q3		--	--	--	--	--	2000Q1-08Q4	--	--
3 Azerbaijan	2001Q1-13Q4	--	--	2000Q1-14Q2	--	--	--	--	--	2000Q1-14Q3	--	--
4 Belarus	2000Q1-14Q1	--	--	2000Q1-14Q3	--	--	--	--	--	2000Q1-14Q3	--	--
5 BiH	--	--	2006Q1-13Q4	--	--	--	--	--	--	2000Q1-14Q3	--	--
6 Bulgaria	2000Q1-14Q2	2000Q1-13Q4	2000Q1-14Q3	2000Q1-14Q3	2008Q1-14Q1	--	2000Q1-13Q4	--	--	2000Q1-14Q3	--	--
7 Croatia	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	--	2000Q4-14Q2	--	--	--	--	2000Q1-14Q3	--	--
8 Czech Republic	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2000Q1-14Q1	--	2000Q1-14Q2	--	--	2000Q1-14Q3	--	--
9 Estonia	2000Q1-14Q1	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2000Q1-14Q1	--	2000Q1-14Q1	--	--	2000Q1-14Q3	--	--
10 Georgia	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2009Q3-14Q2	--	--	--	--	2000Q1-14Q3	--	--
11 Hungary	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q1	--	--	--	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	--	--
12 Kazakhstan	2000Q1-13Q3	--	2000Q1-14Q1	2000Q1-14Q1	--	--	--	--	--	2000Q1-14Q3	--	--
13 Kyrgyz Republic	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q2	--	--	--	--	--	2000Q1-14Q3	--	--
14 Latvia	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2000Q1-14Q3	--	2000Q1-14Q2	--	--	2000Q1-14Q3	--	--
15 Lithuania	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2005Q1-14Q2	--	--	--	--	2000Q1-14Q3	--	--
16 Macedonia	2003Q1-13Q4	2004Q1-13Q4	2000Q1-14Q3	2000Q1-14Q3	--	--	--	--	--	2000Q1-14Q3	--	--
17 Moldova	2000Q1-13Q4	--	2000Q1-14Q3	2000Q1-14Q3	2009Q3-14Q2	--	--	--	--	2000Q1-14Q3	--	--
18 Montenegro	--	--	2005Q1-14Q1	2006Q1-14Q1	--	--	--	--	--	2000Q1-14Q3	--	--
19 Poland	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2009Q4-14Q1	--	--	2000Q1-09Q4	2000Q1-09Q4	2000Q1-14Q3	--	--
20 Romania	2000Q1-14Q2	2000Q1-13Q4	2000Q1-14Q3	2000Q1-14Q3	2009Q2-14Q2	--	--	--	--	2000Q1-14Q3	--	--
21 Russian Federation	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2005Q2-08Q4	--	--	--	--	2000Q1-14Q3	--	--
22 Serbia	2001Q1-14Q1	2001Q1-14Q1	2000Q1-14Q3	2000Q1-14Q3	--	--	--	--	--	2000Q1-14Q3	--	--
23 Slovakia	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	2006Q4-14Q2	--	--	--	--	2000Q1-14Q3	--	--
24 Slovenia	2000Q1-14Q2	2000Q1-14Q2	2000Q1-14Q3	2000Q1-14Q3	--	--	2000Q1-14Q1	--	--	2000Q1-14Q3	--	--
25 Tajikistan	--	--	2000Q2-13Q4	2001Q1-13Q4	--	--	--	--	--	2000Q1-14Q3	--	--
26 Turkmenistan	--	--	--	--	--	--	--	--	--	--	--	--
27 Ukraine	2000Q1-14Q2	2001Q1-14Q1	2000Q1-14Q3	2000Q1-14Q3	--	--	--	--	--	2000Q1-14Q3	--	--
28 Uzbekistan	--	--	--	--	--	--	--	--	--	--	--	--

Source: International Financial Statistics, IMF eLibrary (accessed December 2014)

Table 2. JPM EMBI Global Stripped Spread[2], Daily -- Data Availability

		From	to	obs
1	Bulgaria	29/07/1994	05/11/2012	4767
2	Poland	24/10/1995	05/11/2012	4445
3	Croatia	30/08/1996	05/11/2012	4222
4	Russia	31/12/1997	05/11/2012	3874
5	Hungary	29/01/1999	05/11/2012	3592
6	Ukraine	31/05/2000	05/11/2012	3244
7	Serbia	01/07/2005	05/11/2012	1917
8	Kazakhstan	29/06/2007	05/11/2012	1398
9	Georgia	30/06/2008	05/11/2012	1136
10	Lithuania	31/11/2009	05/11/2012	766
11	Belarus	30/09/2010	05/11/2012	548
12	Romania	29/02/2012	05/11/2012	179

Source: DataStream (accessed November 2012)

While we have only been able to use data at the intersection of the table 2 and table 3 for daily frequencies empirical analysis and no more than the data, which overlap among all of the tables 1, 2, and 3 (for quarterly data estimates), we have been careful not to push our analysis beyond what both available and reliable data permits.

[2]*Note: Time until maturity -- Of the issues with at least a current face amount outstanding of US$500 million, only those instruments with at least 2½ years until maturity are considered for inclusion. Once added, an instrument may remain in the EMBI Global until 12 months before it matures. On the month-end preceding this anniversary, the instrument is removed from the EMBI Global (JP Morgan Securities Inc, Introducing the JP Morgan Emerging Markets Global (EMBI Global), 1999, New York).*

Table 3. Credit Default Swaps (CDS USD 5Y), Daily – Data Availability

		From	to	obs
1	Bulgaria	17/10/2005	19/11/2012	1405
2	Poland	25/01/2006	19/11/2012	1339
3	Croatia	02/12/2005	19/11/2012	1374
4	Russia	24/02/2006	19/11/2012	1576
5	Hungary	21/10/2005	19/11/2012	1324
6	Ukraine	27/02/2006	19/11/2012	1508
7	Serbia	04/05/2006	19/11/2012	1065
8	Kazakhstan	03/04/2007	10/10/2012	1243

Source: Bloomberg (accessed November 2012)

The dataset

We use daily data obtained directly from Bloomberg and ThompsonReuters. In general the data set for each country starts approximately mid-2006 and ends at mid-2012, comprising on average about 1600 observation per country. Technically the estimation is executed in Microfit 4.1 and EViews 6.

5. Sovereign Bond Spreads, Financial Markets Determinants – Spread Regressions by Country

A potential default is often mostly associated with an increase in yield spreads. To examine the determinants of sovereign bond spreads we estimate an equation for the sovereign bond spread (as dependent variable) determined by a range of exogenous variables.

Furthermore we assess the long-term determinants and short-run dynamics (error-correction model) of the sovereign bond spreads of Bulgaria, Croatia, Hungary, Kazakhstan, Poland, Russia, Serbia, and Ukraine – these are the relevant countries for which we have managed to obtain meaningful data, both statistically and economically. Likewise, we gain some additional understanding of the convergence process. Based on this specification we may be able to illustrate quantitatively the impact improved investors' confidence may have upon financing conditions as depicted by government bond spreads.

Emerging Markets Bond Indices

Figure 1 (below) depicts the developments in sovereign stripped spreads for selected CEE and Caucasus and Central Asia (CCA) countries over the period of 1994 to 2012. Over the period starting from the end of 2005 to around the first quarter of 2007, sovereign spreads clustered closely together, reaching their historically lowest point of below 200 basis points. Given that, undoubtedly, there were significant differences in the creditworthiness of the borrowers in the index -- this state of affairs at that time might suggest that investors did not differentiate adequately among borrowers. This situation was followed eventually by the Bear Sterns alarm in March 2008, which led to the increased discrimination in spreads across countries. Furthermore, the spreads widened extensively after September 2008, following the bankruptcy of the Lehman Brothers.

Hence, the key question is: was the narrowing of the spreads and their compression a result of an improvement of CEECCA country sovereigns' macroeconomic policy, implemented after 2002, or was it due to global excess liquidity provision?

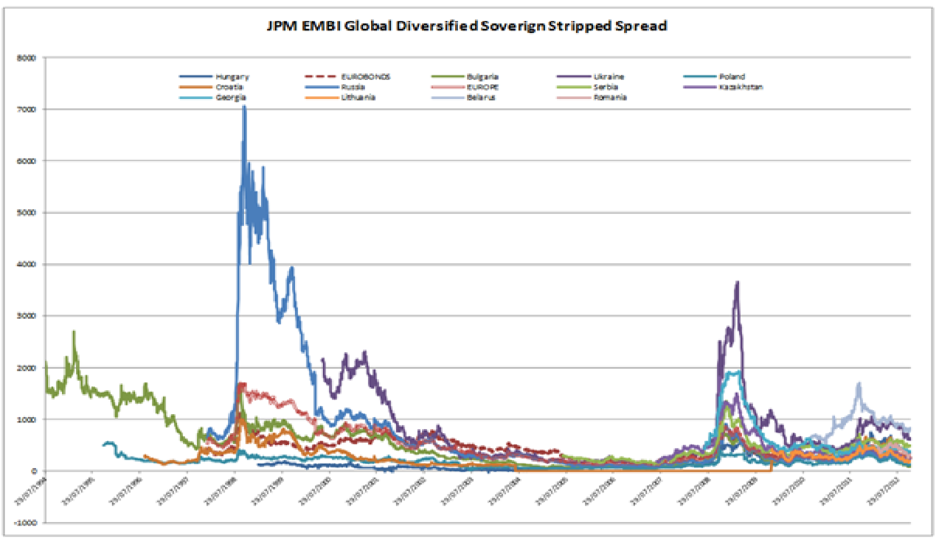


Figure 1. The Emerging Markets Bond Indices (EMBI) Sovereign Stripped Spread, Daily

Credit default swaps (CDS)

The spreads in Figure 2 (below) are for five-year contracts on CDSs with the spreads measured in basis points -- each basis point is equal to USD 1,000. Seemingly comparable to an insurance contract, purchasers of a CDS pay for insurance against a credit event on the public debt. Hence, they can be used as a convenient, standard risk measure on government debt quality. For illustration, the Ukraine five-year CDS, the insurance premium is the annual insurance payment relative to the amount of debt; in March 2009, these CDSs reached a spread of more than 3,800 basis points (with even more extreme values on a daily basis, as can be seen at chart 3, below), meaning that the buyer pays an insurance premium of about 38 percent per year of the value of the securities (i.e., USD 3,800,000 on $10,000,000 worth of debt). The credit default swap seller collects the premiums and pays out (the face value) if a credit event occurs. Thus CDS spreads can be interpreted as a measure of the perceived risk that a government will restructure or default on its debt. CDS spreads in April 2012 imply that the perceived probability of the Ukraine government defaulting is substantially higher than it was one year earlier, but lower than in 2009.

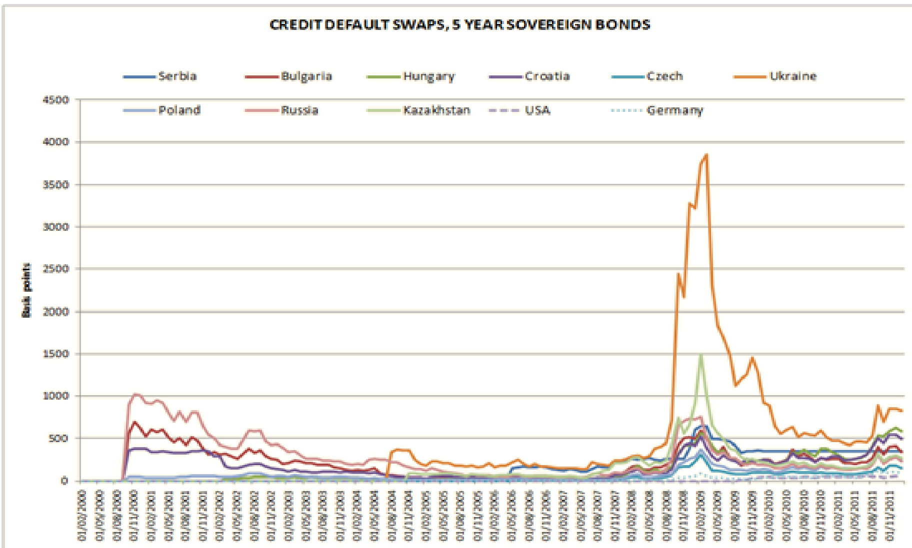


Figure 2. Emerging Markets Credit Default Swaps, Monthly

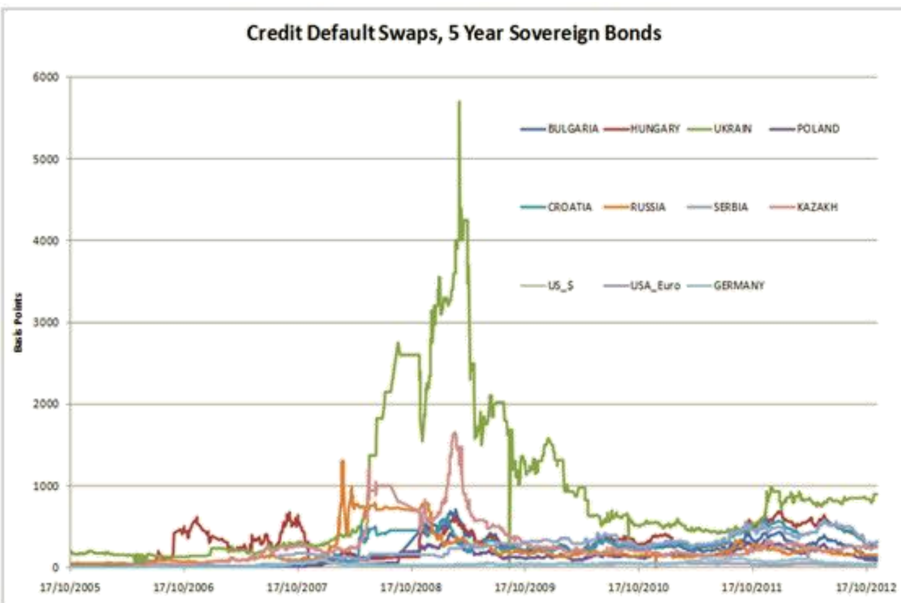


Figure 3. Emerging Markets Credit Default Swaps, Daily

Figures 2 and 3 (above) show the levels of spreads on credit default swaps (CDSs) for selected CEECCA countries sovereign debt plus the same indicators for two industrialized countries –-USA and Germany. Three countries are notable with high spreads at present (Nov 2012): Ukraine, Hungary, and Serbia (all above 300bp).

Credit default swaps pros and cons are debatable, to say the least, and the question are they instrument providing type of insurance or are they rather a device providing an unobstructed way to taking part in speculation is yet to be answered.

In May 2010 the German Federal Financial Supervisory Authority (BaFin) put into operation a complete ban on taking naked sovereign CDS positions.3 On March 14, 2012, the European Commission adopted a proposal for regulating short selling and certain aspects of credit default swaps, de facto permitting the use of CDS only for the purpose of hedging long positions already held by investors.4 As the Commission points out, there are resemblances between short selling stocks that one does not own and buying CDSs on assets that one does not have. These positions are such that speculators profit from adverse developments in the underlying security, and the positions could contribute to a decline in prices in the underlying assets, e.g., prices of government debt.

Economic theory is yet to provide an unambiguous answer to the long standing question about whether speculation in general and in derivative markets in particular is proving predominantly stabilizing or rather destabilizing to any given economic system. For example Portes (2010) concludes: “Banning naked CDS will require common action in the US and in the EU, but the political environment is right. We should not lose this opportunity.”At the same time, Duffie (2010) argues that “Regulations that severely restrict speculation in credit default swap markets could have the unintended consequences of reducing market liquidity, which raises trading execution costs for investors who are not speculating, and lowering the quality of information provided by credit default swap rates regarding the credit qualities of bond issuers. Regulations that severely restrict speculation in credit default swap markets could, as a result, increase sovereign borrowing costs somewhat.”

More obviously sovereign CDS spreads can have a potentially important functional role in the process of price discovery. Still, empirical results concerning who leads the price discovery – the sovereign CDS market or the government bond market are mixed and imprecise. These divergences may be partly related to the different time periods, sampling frequency, methodology and a choice of data. The empirical studies have revealed the following mixed conclusions so far: a number of papers provide support for the dominance of the government bond market, while others claim to have verified the primacy of CDS market. Gyntelberg et. al. (2013) find that CDS prices have a tendency to shift first in reaction to news followed by alteration in bond prices in the same direction and eventual convergence. Palladini and Portes (2011) conclude as well that CDS market spreads in general lead bond markets, but the adjustment towards equilibrium is sluggish. Fontana and Scheicher (2010) examine ten euro sovereigns (January 2006 – June 2010) and find that price discovery is uniformly divided between CDS and bond markets. O'Kane (2012) presents comparable results. Aktug et al. (2012) study thirty emerging markets and find that bond markets lead CDS markets largely, but not always. Support for the bond markets leading role is also found in Ammer and Cai (2011). They find a long-term relationship between CDS and bond markets for the majority of countries. Overall tentatively they conclude that the bond market leads the CDS market more often. Giannikos et al. (2013) inspect the links of price discovery via, daily CDS spreads; bond spreads and stock prices over the period 2005-2008 for ten US financial firms. They find that throughout

[3]General Decree of the Federal Financial Supervisory Authority (BaFin) on the prohibition of naked short-selling transactions in debt securities of Member States of the EU whose legal currency is the euro of 18 May 2010 (revoked with effect from 27 July 2010)

[4](14)Buying credit default swaps without having a long position in underlying sovereign debt or any assets, portfolio of assets, financial obligations or financial contracts the value of which is correlated to the value of the sovereign debt, can be, economically speaking, equivalent to taking a short position on the underlying debt instrument. The calculation of a net short position in relation to sovereign debt should therefore include credit default swaps relating to an obligation of a sovereign debt issuer. The credit default swap position should be taken into account both for the purposes of determining whether a natural or legal person has a significant net short position relating to sovereign debt that needs to be notified to a competent authority and where a competent authority suspends restrictions on uncovered credit default swap transactions for the purposes of determining the significant uncovered position in a credit default swap relating to a sovereign debt issuer that needs to be notified to the competent authority.

REGULATION (EU) No 236/2012 OF THE EUROPEAN PARLIAMENT AND OF THE COUNCIL

of 14 March 2012 on short selling and certain aspects of credit default swaps the sample period, CDS and bond spreads are evidently cointegrated -- the CDS market dominating in price discovery. Examining 18 industrial and emerging economies from January 2007 to March 2010, Coudert and Gex (2013) conclude that bonds appear to lead for “low-yield countries” (developed) European economies, while the derivative market tend to be the direction-finder for “high-yield” emerging economies.

Thus the evidence on price discovery presented above is, at any rate, adequate to challenge the conviction that the relatively small CDS market cannot influence bond spreads in sovereign debt markets as its net exposure is just a few per cent of the total government bond stock. Typically the proponents’ justification of this view may go like this: “Profitable manipulation through price impact is difficult. [...] [a]chieving a sizable price impact would require CDS manipulators to take positions that are large relative to the amount of debt outstanding. In the case of the financially weaker Eurozone sovereigns, the aggregate net CDS positions […] represent small fractions of their respective amounts of debt outstanding. With Greece, for example, the aggregate of the net CDS positions held in the entire market has remained well under 3% of the total amount of Greek debt outstanding. […] That is, even if all CDS protection buyers in the market were manipulators, and had conspired to drive up CDS rates, they would have had only a marginal impact on the total amount of sovereign credit risk borne by bond owners and sellers of protection. Supply and demand for the sovereign's credit would cross at a new price that is relatively close to the “fair-market” (unmanipulated) price (Duffie, 2010).”

A crisp competent answer – with which we completely concur -- is provided by Portes (2010), “We are told [...] that because net CDS exposures are only a few percent of the stock of outstanding government bonds, ‘the tail can’t wag the dog’, so the CDS market can’t be responsible for the rising spreads on the bonds. This of course contradicts the argument that the CDS market leads in price discovery because of its superior liquidity. More important, it is nonsense. Over a period of several days in September 1992, George Soros bet around $ 10 billion against sterling, and most observers believe that significantly affected the market – and the outcome. But daily foreign exchange trading in sterling then before serious speculation began was somewhat over $100 billion. The issue is how CDS prices affect market sentiment, whether they serve as a coordinating device for speculation.”

Furthermore, strong empirical support is provided from Shim and Zhu (2010). The authors analyse the time period of January 2003 to June 2009 and conclude: “[t]hat at the peak of the financial crisis the CDS market contributed to higher spreads in the bond market.”

Based on the evidence presented above we identify CDS as explanatory variable.

Chicago Board Options Exchange Volatility Index (VIX) -- Global Risk Aversion Proxy

VIX, was first initiated by the CBOE in 1993 (data series commencing in January 1986), as a weighted measure of the implicit volatility of eight S&P 100 at-the-money options (both put and call). In ten years time, it has been extended to exploit options based on the broader index (S&P 500), offering more precise scrutiny of investors' expectations on future market volatility. Thus VIX is a commonly used measure of market risk and is often referred to as the "investor fear gauge". VIX values bigger than 30 are normally associated with a large amount of volatility due to investor’s fear or insecurity, whereas values under 20 in general correspond to tranquil periods in the markets. When VIX reaches excessively high levels, this tends to imply that economic agents have bought puts as insurance against a falling market (the explanation is following on Investopedia.com, “VIX - CBOE volatility Index”).

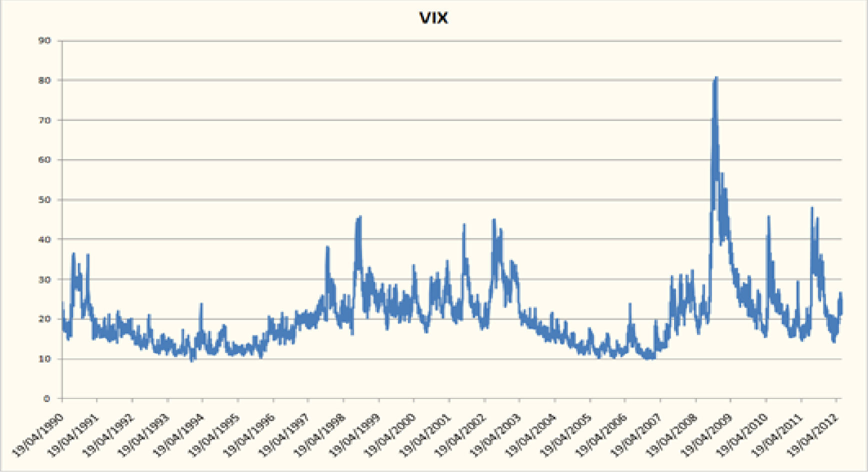


Figure 4. Chicago Board Options Exchange Volatility Index (VIX) – Global Risk Aversion Proxy, daily

We take VIX is an appropriate index to be used in our analysis due to its broad acceptance as representation of investor's expectations market volatility of S&P 500, plus its high frequency, long period time-series availability.

5.1 Sovereign Bond Yield Spreads, Financial Markets Determinants, June 2006 – June 2012, Daily, Estimated Equations and Results

Bulgaria

eq. 2 **SSEMBI = -77.4299 INPT + 0.38379 CDS + 9.9309 VIX - 0.044895 ECM (-1)**

(-6.2919) (7.9586) (13.6259) (-8.3508)

No. obs: 1667

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 27.2769[.000]

Likelihood Ratio Statistic CHSQ(3) = 27.5043[.000]

F Statistic **F(3,1614) = 9.0212[.000]**

R-bar-squared: **0.2137**

DW-statistics: **2.0085**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Bulgaria reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.2 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.2 above. All the explanatory variables are strongly significant (t-ratios shown in parenthesis) and with the expected sign. One point increase in the Bulgaria's risk (approximated by the CDS) would lead to increase of about 0.38 basis points in the dependent variable SSEMBI (Bulgaria's bond's spread) *ceteris paribus*. If the global risk aversion (proxied by VIX) goes up by one point an increase of about 9.9 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.045 implies just less than 15 working days half-life to equilibrium of the Bulgarian bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.2137 suggesting that about 21 per cent of variability in the dependent variable is explained.

Croatia

The joint test for zero restrictions on the coefficients of the lagged level variables does not reject the null hypothesis. Given that the unit root tests suggest that the underlying data series are non-stationary, they have to be modelled in an appropriate – cointegration -- econometric framework to avoid making inferences based on spurious regressions results. However, as the variables are not cointegrated such option is precluded.

Hungary

eq. 3 **SSEMBI = -24.369 INPT + 0.82123 CDS + 3.9391 VIX - 0.0416385 ECM (-1)**

(-1.3285) **(18.3975) (4.8582) (-6.3525)**

No. obs: 1662

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 158.6390[.000]

Likelihood Ratio Statistic CHSQ(3) = 166.8028[.000]

F Statistic **F(3,1608) = 57.0919[.000]**

R-bar-squared: **0.0675**

DW-statistics: **1.8846**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Hungary reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.3 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.3 above. All the explanatory variables are strongly significant (t-ratios shown in parenthesis) and with the expected sign. One point increase in the Hungary's risk (approximated by the CDS) would lead to increase of about 0.82 basis points in the dependent variable SSEMBI (Hungary's bond's spread) *ceteris paribus*. If the global risk aversion (proxied by

VIX) goes up by one point an increase of about 3.9 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.042 implies just less than 17 working days half-life to equilibrium of the Hungarian bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.0675 suggesting that just less than 1 per cent of variability in the dependent variable is explained.

Poland

eq. 4 **SSEMBI = -35.8277 INPT + 0.007321 CDS + 8.3595 VIX - 0.0151 ECM (-1)**

(-1.1903) (0.3946) **(6.8532)** **(-4.9359)**

No. obs: 1664

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 13.5771[.004]

Likelihood Ratio Statistic CHSQ(3) = 13.6332[.003]

F Statistic **F(3,1611) = 4.4527[.004]**

R-bar-squared: **0.0329**

DW-statistics: **2.0014**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Poland reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.4 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.4 above. The explanatory variable VIX and the ECM term are strongly significant (t-ratios shown in parenthesis) and with the expected sign. However, the increase in the Polands's risk effect is too small and not statistically significantly different from zero. If the global risk aversion (proxied by VIX) goes up by one point an increase of about 8.4 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.015 implies about 45 working days half-life to equilibrium of the Poland bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.0329 suggesting that just less than 1 per cent of variability in the dependent variable is explained.

Russia

eq. 5 **SSEMBI = -3.3664 INPT + 0.6300 CDS + 6.0133 VIX - 0.0417 ECM (-1)**

(-0.1722) **(8.7187)** **(4.6800)** **(-6.5702)**

No. obs: 1642

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 31.6218[.000]

Likelihood Ratio Statistic CHSQ(3) = 31.9328[.000]

F Statistic **F(3,1589) = 10.4853[.000]**

R-bar-squared: **0.4946**

DW-statistics: **1.9889**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Russia reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.5 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.5 above. All the explanatory variables are strongly significant (t-ratios shown in parenthesis) and with the expected sign. One point increase in the Russia's risk (approximated by the CDS) would lead to increase of about 0.63 basis points in the dependent variable SSEMBI (Russia's bond's spread) ceteris paribus. If the global risk aversion (proxied by VIX) goes up by one point an increase of about 6.0 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.042 implies just about 17 working days half-life to equilibrium of the Russian bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.4946 suggesting that almost exactly 50 per cent of variability in the dependent variable is explained.

Ukraine

eq. 6 **SSEMBI = 8280.6 INPT + 8.0964 CDS - 604.8879 VIX - 0.0008373 ECM (-1)**

(0.15186) (0.16603) (-0.15024) (0.15269)

No. obs: 1641

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 21.5655[.000]

Likelihood Ratio Statistic CHSQ(3) = 21.7096[.000]

F Statistic **F(3,1588) = 7.1060[.000]**

R-bar-squared: **0.23489**

DW-statistics: **2.0134**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Ukraine reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.6 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.6 above. All the explanatory variables turn out to be statistically insignificant (t-ratios shown in parenthesis) and VIX is with the "wrong" sign. The error correction coefficient of about -0.0008 implies about 866 working days half-life to equilibrium of the Ukraine bond spread, but is statistically insignificant. The coefficient for mutual determination corrected for degrees of freedom equals 0.2348 suggesting that about 23 per cent of variability in the dependent variable is explained. All the explanatory variables being insignificant only in the specific case of Ukraine tend to suggest that the bond spread of the country is driven by other forces, possibly including low quality of governance, corruption and heavy speculation.

Serbia

eq. 7 **SSEMBI = -198.8189INPT + 0.47910CDS + 21.0931VIX - 0.020865ECM (-1)**

(-3.3529) (2.7677) (11.5718) (-5.8916)

No. obs: 1592

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 50.9643[.000]

Likelihood Ratio Statistic CHSQ(3) = 51.8049[.000]

F Statistic **F(3,1539) = 17.1100[.000]**

R-bar-squared: **0.1965**

DW-statistics: **2.0069**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Serbia reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.7 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.7 above. All the explanatory variables are strongly significant (t-ratios shown in parenthesis) and with the expected sign. One point increase in the Serbia's risk (approximated by the CDS) would lead to increase of about 0.48 basis points in the dependent variable SSEMBI (Hungary's bond's spread) ceteris paribus. If the global risk aversion (proxied by VIX) goes up by one point an increase of about 21 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.020 implies just about 34 working days half-life to equilibrium of the Serbian bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.1965 suggesting that just around 20 per cent of variability in the dependent variable is explained.

Kazakhstan

eq. 8 **SSEMBI = -173.21 INPT + 0.3261 CDS + 20.9384 VIX - 0.0417 ECM (-1)**

(-2.5738) (3.3923) (7.3869) (-5.1551)

No. obs: 1292

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(3) = 18.0163[.000]

Likelihood Ratio Statistic CHSQ(3) = 18.1444[.000]

F Statistic **F(3,1239) = 5.9007[.001]**

R-bar-squared: **0.2903**

DW-statistics: **1.9953**

The results of the F-statistic for the joint test of zero restrictions on the coefficients of additional variables for Kazakhstan reject the null hypothesis in favour of the existence of long-run relationship between SSEMBI, CDS and VIX. We estimate eq.8 and get the long-run coefficient; then we obtain the estimates of the error correction model associated with these long-run estimates and report the outcome as eq.8 above. All the explanatory variables are strongly significant (t-ratios shown in parenthesis) and with the expected sign. One point increase in the Kazakhstan's risk (approximated by the CDS) would lead to increase of about 0.33 basis points in the dependent variable SSEMBI (Kazakhstan's bond's spread) ceteris paribus. If the global risk aversion (proxied by VIX) goes up by one point an increase of about 21 basis points in SSEMBI would be induced everything else remaining the same. The error correction coefficient of about -0.042 implies just about 16 working days half-life to equilibrium of the Kazakhstan bond spread. The coefficient for mutual determination corrected for degrees of freedom equals 0.2903 suggesting that about 29 per cent of variability in the dependent variable is explained.

6. Sovereign Bond Spreads, Financial Markets Determinants: Cross Sectional Time Series Estimate – Pooled Least Squares

The cross-sectional-time-series (CSTS) data contains valuable information about both: i) changes between the subjects (cross-sectional information); and, ii) changes within the subjects (time-series information).

Turning to the panel data model, first we perform series of unit-root tests (checking both for individual and common unit root processes), on the basis of which, we are not able to reject the presence of unit roots (detailed results of the tests are presented in Annex 1) in the data.

Next we perform panel cointegration tests (see Annex 2), all of which reject the null hypothesis of no cointegration. Hence, given that our variables are cointegrated we proceed with estimating both fixed and random effects (cointegrated panels) models. In general, the fixed effects model assumes that each country differs in its intercept term, while the random effects model assumes that each country differs in its error term. We perform two test: i) Pedroni residual cointegration test; and Kao residual cointegrtion test.

All of the eleven statistics reported in the test of Pedroni reject the null hypothesis of no cointegration at a very high level of significance. The same strong result is obtained from the Kao test (Annex 2).

It should be noted that the literature on panel cointegration is still in a process of development and fine-tuning. In particular cointegration tests based on cross sectional dependence when improved further should replace / should be used together with the Pedroni and Kao (first-generation) tests which assume cross-sectional independence.

As a next step we proceed with estimating a fixed effect (FE) model. The results are shown at Table 5. below.

Table 4. Pooled Least Squares Fixed Effects Model, Estimation Results

Dependent Variable: SSEMBI
Total pool (balanced) observations: 9546

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-31.91908	4.139766	-7.710360	0.0000
CDS	0.424554	0.004007	105.9562	0.0000
VIX	10.13878	0.169247	59.90516	0.0000
Fixed Effects (Cross)				
BGN--C	-62.44306			
HUN--C	-51.47301			
POL--C	-118.8528			
RUS--C	-32.03785			
SER--C	99.14332			
UKR--C	165.6634			

R-squared	0.780282	Mean dependent var	342.4605
Adjusted R-squared	0.780121	S.D. dependent var	361.3367
S.E. of regression	169.4354	Akaike info criterion	13.10366
Sum squared resid	2.74E+08	Schwarz criterion	13.10966
Log likelihood	-62535.76	Hannan-Quinn criter.	13.10570
F-statistic	4838.884	Durbin-Watson stat	0.731852
Prob(F-statistic)	0.000000		

The fixed effects coefficients differ in sign and size. Consequently, we test for (unobserved) heterogeneity. The test applied is the standard (in EViews) Redundant Fixed Effects Tests, where the null hypothesis is that the fixed effects are all equal to each other.

Table 5. Redundant fixed effects test

Redundant Fixed Effects Tests
Pool: POOL01
Test cross-section fixed effects

Effects Test	Statistic	d.f.	Prob.
Cross-section F	559.253226	(5,9538)	**0.0000**
Cross-section Chi-square	2454.252098	5	**0.0000**

The p-values related to the F-statistic and the Chi-square statistics are both very small, (see Table 5, above) providing strong evidence against the null hypothesis and suggesting the existence of heterogeneity.

Next we plot and examine both the residual correlation and residual covariance matrices:

Table 6. Residual Correlation Matrix

	BGN	HUN	POL	RUS	SER	UKR
BGN	**1.000000**	0.616180	0.111101	0.833714	0.378806	0.010543
HUN	0.616180	**1.000000**	0.204054	0.742839	0.179843	-0.015597
POL	0.111101	0.204054	**1.000000**	0.146597	-0.060416	-0.075385
RUS	0.833714	0.742839	0.146597	**1.000000**	0.458929	0.097838
SER	0.378806	0.179843	-0.060416	0.458929	**1.000000**	0.761338
UKR	0.010543	-0.015597	-0.075385	0.097838	0.761338	**1.000000**

The correlation matrix indicates that there certainly is correlation observed among cross-sections. Interestingly, Ukraine displays negative correlations with Poland and Hungary, and such effect obtains between Serbia and Poland: an “anti-contagion” effect.

Table 7. Residual Covariance Matrix

	BGN	HUN	POL	RUS	SER	UKR
BGN	**2906.718**	3412.749	1702.920	2866.242	2700.845	134.8261
HUN	3412.749	**10553.35**	5959.585	4866.135	2443.262	-380.0592
POL	1702.920	5959.585	**80826.37**	2657.636	-2271.468	-5083.698
RUS	2866.242	4866.135	2657.636	**4066.196**	3870.088	1479.853
SER	2700.845	2443.262	-2271.468	3870.088	**17488.88**	23882.24
UKR	134.8261	-380.0592	-5083.698	1479.853	23882.24	**56264.30**

The diagonal demonstrates the variances of the residuals for each cross-section in bold; the remaining numbers of the matrix show the covariance of the residuals across cross-sectional units. Based on the results from tables 6 and 7, above, we explore the opportunity to obtain an efficient estimator (using EGLS with SUR weights) by utilising the correlations between the residuals. The results of the re-estimated model are presented below.

Table 8. Fixed effects model using estimated generalized least squares (EGLS) with seemingly unrelated regression (SUR) weights

Dependent Variable: SSEMBI
Total pool (balanced) observations: 9546

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	14.87775	2.999460	4.960143	0.0000
CDS	0.417642	0.002665	156.7225	0.0000
VIX	8.265180	0.123106	67.13875	0.0000
Fixed Effects (Cross)				
BGN--C	-63.15626			
HUN--C	-51.96075			
POL--C	-119.9469			
RUS--C	-32.85247			
SER--C	98.86306			
UKR--C	169.0533			

Weighted Statistics

R-squared	0.844508	Mean dependent var	1.607103
Adjusted R-squared	0.844394	S.D. dependent var	2.841986
S.E. of regression	0.983804	Sum squared resid	9231.543
F-statistic	7400.421	Durbin-Watson stat	0.343481
Prob(F-statistic)	0.000000		

Unweighted Statistics

R-squared	0.776572	Mean dependent var	342.4605
Sum squared resid	2.78E+08	Durbin-Watson stat	0.691820

The estimates of CDS and VIX are to some extent smaller, but as the heteroscedasticity EGLS is more efficient than OLS estimator the standard error of CDS and VIX are less significant.

Next we experiment with estimating a random effects (RE) model (Table 9, below)

Table 9. Random Effects Model, Estimation Results

Dependent Variable: SSEMBI
Method: Pooled EGLS (Cross-section random effects)
Total pool (balanced) observations: 9546
Swamy and Arora estimator of component variances

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-31.80419	12.44861	-2.554839	0.0106
CDS?	0.426632	0.003996	106.7746	0.0000
VIX?	10.10665	0.169184	59.73777	0.0000
Random Effects (Cross)				
BGN--C	-60.89944			
HUN--C	-50.22998			
POL--C	-115.9926			
RUS--C	-31.11365			
SER--C	97.10691			
UKR--C	161.1287			
			S.D.	Rho
Cross-section random			28.75731	0.0280
Idiosyncratic random			169.4354	0.9720

Weighted Statistics

R-squared	0.702054	Mean dependent var	50.04312
Adjusted R-squared	0.701992	S.D. dependent var	311.4249
S.E. of regression	170.0072	Sum squared resid	2.76E+08
F-statistic	11243.17	Durbin-Watson stat	0.733375
Prob(F-statistic)	0.000000		

While the regression coefficients obtained are practically identical to those of the fixed effects model, the random effects model presumes that the random effects are uncorrelated with the explanatory variables – if not the estimators would be rendered inconsistent (endogeneity problem). We apply the Hausman test (Correlated Random Effects) to test this hypothesis.

Table 10. Correlated random effects – Hausman test

Correlated Random Effects - Hausman Test
Pool: POOL01
Test cross-section random effects

Test Summary	Chi-Sq. Statistic	Chi-Sq. d.f.	Prob.
Cross-section random	66.519114	2	0.0000

Cross-section random effects test comparisons:

Variable	Fixed	Random	Var(Diff.)	Prob.
CDS?	0.424554	0.426632	0.000000	0.0000
VIX?	10.138775	10.106654	0.000021	0.0000

The test (Table 11, above) rejects the null hypothesis at all conventional levels of confidence. Hence, the assumption that the random effects are uncorrelated to the explanatory variables is not acceptable, not allowing us to continue further with this approach.

7. Sovereign Bond Spreads, Macroeconomic Determinants – Spread Regressions by Country

In what follows we move to quarterly data frequency and try to assess the effect of the macroeconomic variables listed below as determinants of the sovereign bond spreads. We continue by applying the ARDL procedure.

SSEMBI **--** Stripped Spread JPM EMBI GLOBAL

VIX -- Volatility Index (proxy for global risk aversion)

PDGDP -- Government debt as per cent of GDP (Bulgaria, Croatia and Hungary)

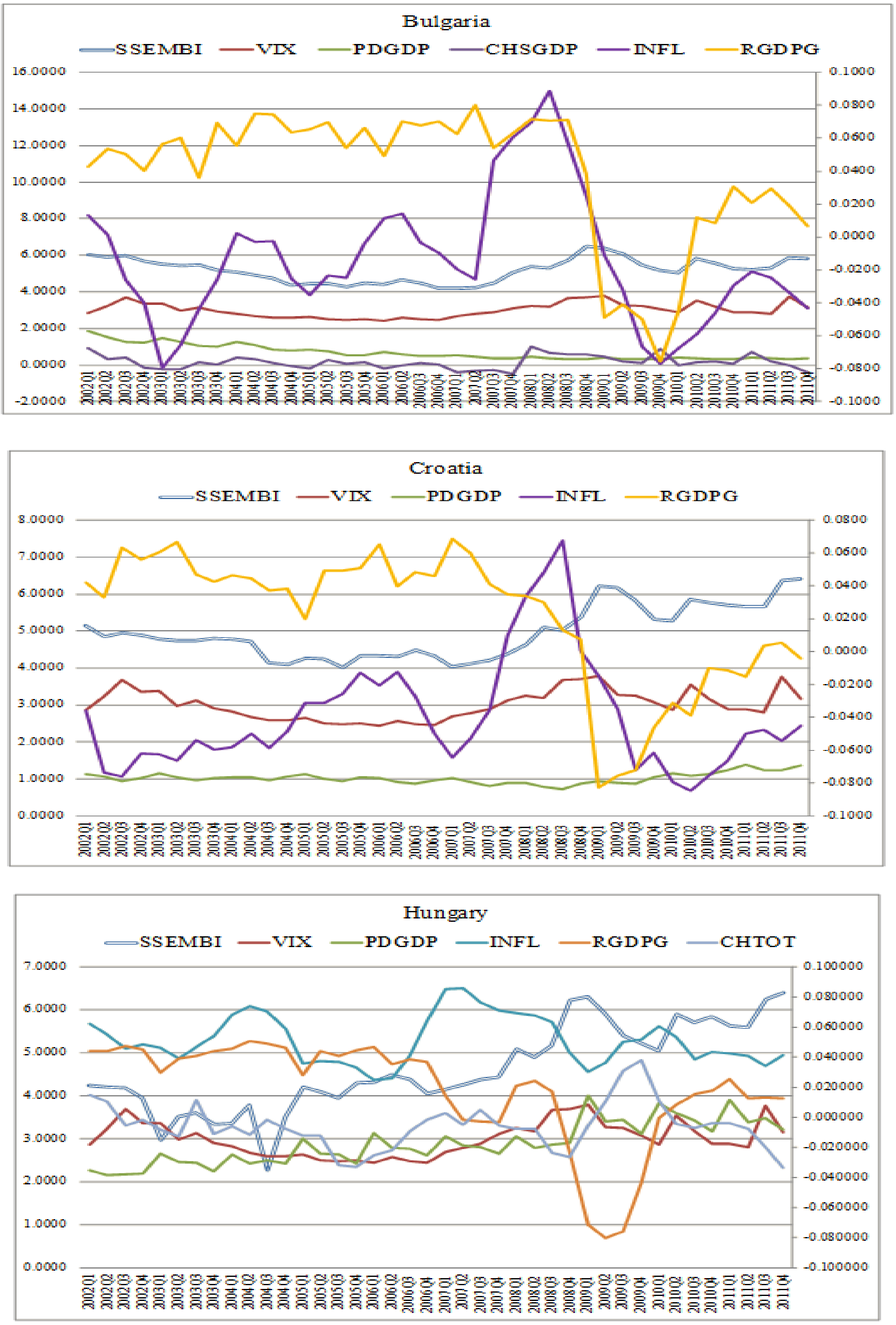
RGDPG -- Real GDP growth

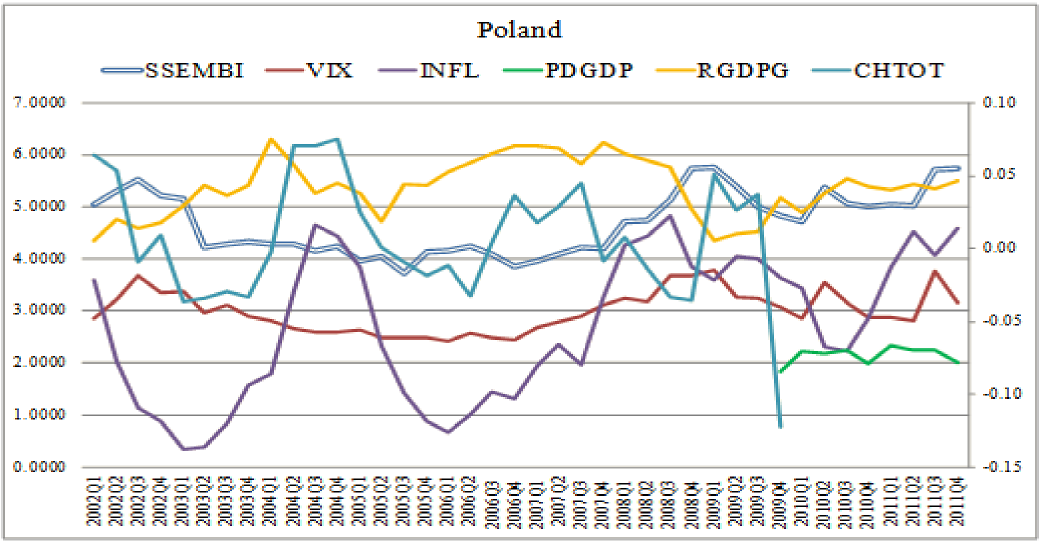
INFL -- Relative change in CPI

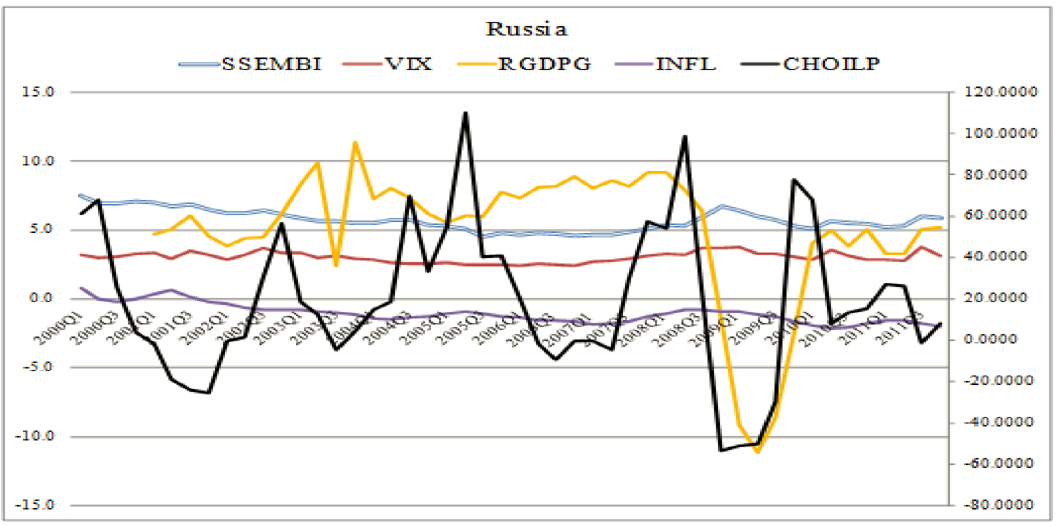
CHTOT – Change in the Terms of Trade (only for Hungary and Poland)

CHOILP – Change in Oil Prices (only for Russia)

Figure 5. Sovereign bond yield spreads, potential macroeconomic determinants: Bulgaria, Croatia, Hungary, Poland, and Russia







To test the existence of a long-run relationship between variables we estimate the error correction depiction of an underlying ARDL for five countries (Bulgaria, Croatia, Hungary, Poland, and Russia) for which there are data (to a degree) available over the period 2002Q2 to 2011Q4.

For Bulgaria and Croatia the ARDL model is:

$$DSSMBI_t = INPT + \sum_{i=1}^{4} \alpha DSSEMBI_{t-i} + \sum_{i=1}^{4} \beta DVIX_{t-i} + \sum_{i=1}^{4} \gamma DPDGDP_{t-i} + \sum_{i=1}^{4} \delta DRGDPG_{t-i}$$

$$+ \sum_{i=1}^{4} \theta DINFL_{t-i} + \pi_1 SSEMBI_{t-1} \;\; \pi_2 VIX_{t-1} + \pi_3 PDGDP_{t-1} + \pi_4 RGDPG_{t-1}$$

$$+ \pi_5 INFL_{t-1} + \varepsilon_t$$

We test the null hypothesis of the non-existence of a long-run relationship, i.e.,

$$H_0\colon \; \pi_1 = \pi_2 = \pi_3 = \pi_4 = \pi_5 = 0$$

versus

$$H_1\colon \pi_1 \neq 0, \; \pi_2 \; \neq 0, \pi_3 \; \neq 0, \pi_4 \; \neq 0, \; \pi_5 \; \neq 0$$

Bulgaria

Comparing the F-statistic (2.0662) obtained (below) with the critical value bounds determined by Pesaran, Shin and Smith (1996), the critical values at the 90 per cent level are specified as 2.425 to 3.574. Since the F-statistics is below the lower bound of the critical range, we cannot reject the null of no long-run relationship independent of the order of integration of the respective variables.

No. obs: 39

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(5) = 17.2694[.004]

Likelihood Ratio Statistic CHSQ(5) = 22.8087[.000]

F Statistic **F(5, 13) = 2.0662[.135]**

R-bar-squared: 0.6154

DW-statistic: 1.756

Still only for illustrative purposes we estimate the long-run coefficients and their levels of significance (t-statistics):

eq. 9 **SSEMBI** = **-.19174*INPT + 1.7028*VIX + .15018*PDGDP - .92055*RGDPG +.053177*INFL -.44284*ECM(-1)**

(-0.2468) (**2.8587**) (0.8537) (-0.7010) (1.7650) (**5.7035**)

Croatia

Following the same procedure we obtain:

No. obs: 39

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(5) = 14.7731[.011]

Likelihood Ratio Statistic CHSQ(5) = 18.5678[.002]

F Statistic **F(5, 28) = 3.4148[.016]**

R-bar-squared: 0.47318

DW-statistic: 1.8718

The value of the F-statistic (3.4148) obtained (above) falls within the critical value band (at the 90 per cent level) specified by 2.425 to 3.574. Hence, the results are inconclusive.

Again, only for illustrative purposes we estimate the long-run coefficients and their levels of significance (t-statistics):

eq. 10 **SSEMBI= -1.4081*INPT + 1.3656*VIX + 2.2858*PDGDP - .3836*RGDPG +.20194*INFL -.4093*ECM(-1)**

(-0.9671) **(4.8073)** **(3.1512)** **(-3.8066)** **(-2.5435)** **(-4.1130)**

Next we extend slightly the model to include the change in the terms of trade variable (CHTOT), below, and apply it for Hungary and Poland

$$DSSMBI_t = INPT + \sum_{i=1}^{4} \alpha DSSEMBI_{t-i} + \sum_{i=1}^{4} \beta DVIX_{t-i} + \sum_{i=1}^{4} \gamma DPDGDP_{t-i} + \sum_{i=1}^{4} \delta DRGDPG_{t-i}$$
$$+ \sum_{i=1}^{4} \theta DINFL_{t-i} + \sum_{i=1}^{4} \varphi DCHTOT_{t-i} + \pi_1 SSEMBI_{t-1} + \pi_2 VIX_{t-1} + \pi_3 PDGDP_{t-1}$$
$$+ \pi_4 RGDPG_{t-1} + \pi_5 INFL_{t-1} + \pi_6 CHTOT_{t-1} + \varepsilon_t$$

We test the null hypothesis of the non-existence of a long-run relationship, i.e.,

$$H_0: \pi_1 = \pi_2 = \pi_3 = \pi_4 = \pi_5 = \pi_6 = 0$$

versus

$$H_1: \pi_1 \neq 0,\ \pi_2 \neq 0, \pi_3 \neq 0, \pi_4 \neq 0,\ \pi_5 \neq 0,\ \pi_6 \neq 0$$

Hungary:

No. obs: 39

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(6) = 33.4248[.000]

Likelihood Ratio Statistic CHSQ(6) = 108.5336[.000]

F Statistic **F(6, 4) = 14.1462[.011]**

R-bar-squared: 0.3526

DW-statistic: 2.2134

We compare the F-statistic (14.1462) with the critical value bounds determined by Pesaran, Shin and Smith (1996). The critical values at the 99 per cent level are specified by 3.516 to 4.781. Since the F-statistics is above the upper bound of the critical value, we reject the null of no long-run relationship unconnected of the order of integration of the respective variables.

Then based on the Schwartz Bayesian information criteria (SBC) we select the ARDL(1,0,1,0,0,0) model specification and estimate the long-run coefficients; subsequently we estimate the error correction model related to these long-run coefficients and we get:

eq. 11 **SSEMBI = -3.4149*INPT + .93477*VIX + 1.7531*PDGDP - 3.1682*RGDPG + 2.2919*INFL - 24.6521*CHTOT - .59845*ECM(-1)**

(-2.1494) **(2.8587)** **(4.7103)**
(-0.6100) (-0.4263) **(-2.5381)** **(-4.2383)**

Not including RGDPG and INFL all other coefficients are statistically significant and with the expected sign. It is interesting to observe that for Hungary the CHTOT is exercising the most substantial effect on SSEMBI, i.e., one unit increase in the terms of trade would lead to an almost 25 basis points reduction in the spread (SSEMBI). The error correction coefficient is strongly significant, has the correct sign and implies a half-life to convergence of about 50 working days.

Poland

The value of the F-statistic (3.6998) attained (below) is just above the higher critical value bound (at the 90 per cent level) specified by 2.425 to 3.574. Hence, at this level, we can reject the null hypothesis of no long run relationship.

No. obs: 39

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(5) = 17.0314[.004]

Likelihood Ratio Statistic CHSQ(5) = 25.6654[.000]

F Statistic **F(5, 13) = 3.6998[.027]**

R-bar-squared: 0.67176

DW-statistic: 2.2495

Next, on the basis of the SBC we select the ARDL(1,2,1,1,0) model specification, then estimate the long-run coefficients and the error-correction model related to them.

eq. 12 **SSEMBI= -1.9125*INPT + .96280*VIX - 7.0988*RGDPG +.17453*INFL + 1.1700*CHTOT -.7010*ECM(-1)**

(3.1198) **(5.4058)** **(-2.5313)**
(2.6459) (0.8464) **(4.4196)**

With the exception of CHTOT all coefficients are statistically significant and with the expected sign. We observe that for Poland the RGDPG is having the most important effect on SSEMBI, i.e., one unit increase in the terms of trade would lead to about seven basis points reduction in the spread (SSEMBI). The error correction coefficient is strongly significant, has the correct sign and implies a half-life to convergence of about 38 working days.

Finally, we amend somewhat the model to include the change in oil prices variable (CHOILP), and remove the PDGDP (public debt as per cent of GDP – for which we do not have data) below, and apply it for Russia

$$DSSMBI_t = INPT + \sum_{i=1}^{4} \alpha DSSEMBI_{t-i} + \sum_{i=1}^{4} \beta DVIX_{t-i} + \sum_{i=1}^{4} \delta DRGDPG_{t-i}$$

$$+ \sum_{i=1}^{4} \theta DINFL_{t-i} + \sum_{i=1}^{4} \varphi DCHOILP_{t-i} + \pi_1 SSEMBI_{t-1} + \pi_2 VIX_{t-1} + \pi_3 PDGDP_{t-1}$$

$$+ \pi_4 RGDPG_{t-1} + \pi_5 INFL_{t-1} + \pi_6 CHOILP_{t-1} + \varepsilon_t$$

We test the null hypothesis of the non-existence of a long-run relationship, i.e.,

$$H_0: \pi_1 = \pi_2 = \pi_3 = \pi_4 = \pi_5 = \pi_6 = 0$$

versus

$$H_1: \pi_1 \neq 0,\ \pi_2 \neq 0, \pi_3 \neq 0, \pi_4 \neq 0,\ \pi_5 \neq 0,\ \pi_6 \neq 0$$

Russia

The value of the F-statistic (3.8821) attained (below) is above the upper critical value bound (at the 90 per cent level) specified by 2.425 to 3.574. Hence, at this level, we can reject the null hypothesis of no long run relationship.

No. obs: 41

Joint test of zero restrictions on the coefficients of additional variables:

Lagrange Multiplier Statistic CHSQ(5) = 17.9200[.003]

Likelihood Ratio Statistic CHSQ(5) = 23.5588[.000]

F Statistic F(5, 25) = 3.8821[.010]

R-bar-squared: 0.62109

DW-statistic: 2.0471

Next, on the basis of the SBC we select the ARDL(1,1,0,0,0) model specification, then estimate the long-run coefficients and the error-correction model related to them.

eq. 13 SSEMBI = -2.7018*INPT + 2.5998*VIX + .090291*RGDPG -.016650*INFL + 0.00552*CHOILP -.19191*ECM(-1)

(0.98488) (**2.7383**) (1.4145) (-0.27910) (0.8464) (**2.0286**)

With the exception of VIX all coefficients are not statistically significant and with the "wrong" sign. Interestingly, one of these coefficnts is CHOILP. The error correction coefficient is significant and has the correct sign. However, it implies quite a long half-life to convergence of about 215 working days.

8. Concluding Remarks and Policy Implications

First we analyse the financial markets (variables) explanatory power (using proxies for change in market sentiment (VIX) and for adjustment in country's risk (CDS)) over the emerging market bond index spread on a country by country basis.

Using the F-statistic test for joint significance of zero restrictions on the lagged levels of the additional variables (Pesaran, Shin and Smith, 1996) we cannot reject at conventional significance levels the null hypothesis that sovereign bond spreads are cointegrated with the VIX and the country specific CDS[5].

[5]With the single exception of Croatia.

On examination most of the explanatory variables are strongly significant (t-ratios are presented in parenthesis) and have the expected signs. The underlying ARDL equations also pass the diagnostic tests in the majority of cases.

Studying the range of the estimated values we observe that a one point increase in the country's risk (as measured by the CDS) would induce an increase in the region of about half a basis point (ranging from about 0.33 to 0.82) in the dependent variable SSEMBI (bond's spread), everything else remaining the same. If VIX (the proxy for global risk aversion) goes up by one point, this will induce on average about an 11 basis points increase (displaying values from about 3.9 to just above 21) in the country's spread.

The error correction coefficient estimates are within the cluster of –0.015 to -0.044 suggesting a reasonable speed of convergence to equilibrium, with a half-life reporting from fewer than 15 working days to about 45 working days. Hence, in just about two-thirds of a quarter the spread (SSEMBI) should return to its equilibrium. Interestingly, the error correction coefficients and hence the speed of convergence for most of the countries (Bulgaria, Hungary, Russia, and Kazakhstan) is almost one and the same (in the vicinity of –0.042 to -0.044). Therefore it is evident that hypothetically they would converge back to their respective equilibrium values for the SSEMBI more than three times as fast as Serbia and Poland.

The coefficients for mutual determination corrected for degrees of freedom are generally in-between 0.2 to 0.5 suggesting that about 20 to 50 per cent of the variability in the dependent variable (SSEMBI) has been explained. The exceptions are Hungary and Poland, where just about five per cent (on average) of the variability of the respective dependent variable is explained.

Furthermore, for Serbia the tests (for joint significance) suggest that the variables CDS and VIX can be treated as the long-run forcing variables for the dependent variable SSEMBI. Interestingly while this is valid for Serbia, for Poland, Russia and Ukraine our results suggest a bidirectional relationship between CDS (as potential dependent variable) and SSEMBI and VIX, and non-rejection of the null hypothesis that the lagged level variables CDS and SSEMBI do not enter significantly in the potential determination (potential equation) of VIX. In the case of Kazakhstan the null hypothesis that the lagged values of SSEMBI and VIX do not enter significantly in the determination of CDS cannot be rejected, but there is an apparent relationship between VIX and CDS and SSEMBI. Regarding Bulgaria and Hungary we observe complete bidirectional interrelations among all three variables.

In our analysis we estimate separate equations / data generation processes for the various (former centrally planned) economies and find statistically significant and economically perceivable coefficients. The data shortage precluded any potential experimentation with different specifications or another dataset. Hence, if the coefficients tend to be homogenous, pooled panel estimation would be useful and suitable to be exploited.

For this reason we estimate cointegrated pooled panel models. The results from the fixed effects and random effects pooled panel data models are practically identical and are consistent with our previous findings from the individual equation estimates. Concretely, a one point increase in CDS (proxy for country risk) would add about 0.42 basis points to the variable SSEMBI, *ceteris paribus*; whereas a one unit increase of VIX (stand-in for global risk aversion) would bring about an 8.3 basis points increase in SSEMBI. The coefficient of mutual determination corrected for degrees of freedom is very high, suggesting that about 84 per cent of the variability of the dependent variable (sovereign bond spreads) is explained.

Next we examine the effect of a change in macroeconomic fundamentals on changes of spreads. A relatively noteworthy proportion of fluctuations in transition economies market spreads may be attributed to be driven by country-specific fundamentals. The results imply that improved macroeconomic fundamentals, such as lower ratios of debt to GDP, higher rates of real GDP growth, and low inflation help in reducing sovereign spreads.

For example, reduced indebtedness seems to contribute positively to sovereign spreads in Hungary; one may expect the same to be valid for Poland, but in the case of Poland, the model did not include any measure of indebtedness due to the lack of a time series from (at least) 2001Q1.

It is interesting that in the cases of Bulgaria and Russia we find four insignificant independent variables, whereas these are significant for some of the other countries. This seems to be a possible indication of institutional weakness, limiting the effect of the stance of the macroeconomic aggregates and making their impact trivial. This result is in agreement with Hallerberg and Wolff (2006).

Still, macroeconomic aggregates play a certain role in determining bond spreads, but mostly through the channel of global risk aversion / appetite corroborating Favero and Missale (2011) for our specific set of (CEE and CCA) countries.

Evidently, the only variable which appears in both financial market reaction and macroeconomic fundamentals equations and works strongly and consistently in the same direction is VIX. This suggest that the levels of spreads can be subject to significant alteration from the impact of financial market volatility (as measured by VIX) and could potentially be pushed up or down in ways that have little to do with their respective macroeconomic fundamentals.

The error correction coefficients suggest a return to equilibrium (with half-life) in the range of about 38 to 50 working days[6] – a very similar order of magnitude to that derived in the financial market high frequency data sample equations.

This sheds light and provides clear evidence on the critical factors that have a significant influence on the variation in spreads in the transition countries environment -- in reality worldwide factors are principally responsible for the changes in spreads. Hence, any kind of government intervention aiming to to bring down spreads may prove ineffective, unless strongly determined and unfalteringly pre-coordinated.

Now we may ask: has the transition ended? It is debatable, and an agreement on the appraisal of the results of transition is impractical as there are expectations, attitudes and beliefs involved. What would be the appropriate criteria? Obvious cases to look at for constructive suggestions would be Japan, South Korea and China. In their cases it seemed to be self-evident: supreme economic success guided by the respective government (developmental state). Considering transition economies; whatever their pros and cons; neither of them matches the remarkable economic growth achieved by the previous group. Why might that be? The answer is closely linked to the quality of governance, human capital development and corruption, and as a result the level of development of the social knowledge and its practical implementation, i.e., this generally is manifested by the stage of development of manufacturing.

Transition would then end when the transition economies find their place in the global production process and become equal partners with the industrialised world economies -- to become integrated into the international economic framework rather than to be subordinated to it. This would depend on their abilities in developing and exploiting knowledge in the contemporary exceptionally competitive world economy. If Government maintains strong incentives to provide public goods and retains motivation for wealth creation through the efficient use of capital and labour, as an outcome, the economy would remain connected to its comparative advantage, which (for a low-rent country) lies initially in labour-intensive manufactured goods. The brief initial dependence on primary product exports (of low-rent economies) encourages industrialization at a relatively low per capita income, which is therefore labour-intensive and competitive and triggers a beneficial economic advancement. Moreover, competitive diversification increases the capacity of the economy to cope with economic shocks and reinforces the resilience that arises from sustained high rates of investment.

There is a relationship between macroeconomic sustainable growth and the financial sector development. Adequate attention needs to be paid to institutional development and regulatory structure. The financial sector / banking sector features that are critical for successful intermediation and indispensible for growth include: i) transparency (e.g., independence of commercial bank governance from detrimental oligarchic “clients”; ii) sufficient central bank independence from government control; iii) macro-prudential policy needs to be oriented towards the resilience of the entire system and careful judgement (rather than just fixed rules) need to be exercised when applying macro-prudential instruments (in dealing with market failures, e.g., moral hazard, information frictions, risk illusion, herd behaviour, etc.); and, iv) enhanced efficiency of international cooperation in this area.

Potential major future research areas would include: dynamic interaction of local and international developments; absorbing capacity of transition economies; markets in transition economies; and, importance of modern manufacturing for transition economies.

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[6]With the exception of Russia where the half-life is about 215 working days

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Annex 1

Unit root tests

a) Null Hypothesis: Unit root (individual unit root process)
Series: SSBGN, SSHUN, SSPOL, SSRUS, SSSER, SSUKR
Date: 01/13/15 Time: 20:12
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 6
Total number of observations: 9533
Cross-sections included: 6

Method	Statistic	Prob.**
Im, Pesaran and Shin W-stat	- 0.23493	0.4071

** Probabilities are computed assuming asympotic normality

b) Null Hypothesis: Unit root (individual unit root process)
Series: SSBGN, SSHUN, SSPOL, SSRUS, SSSER, SSUKR
Date: 01/13/15 Time: 20:15
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 6
Total number of observations: 9533
Cross-sections included: 6

Method	Statistic	Prob.**
ADmnjjjF - Fisher Chi-square	9.11905	0.6927
ADF - Choi Z-stat	-0.11368	0.4547

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

c) Null Hypothesis: Unit root (common unit root process)
Series: SSBGN, SSHUN, SSPOL, SSRUS, SSSER, SSUKR
Date: 01/13/15 Time: 20:14
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 6
Newey-West bandwidth selection using Bartlett kernel
Total number of observations: 9533
Cross-sections included: 6

Method	Statistic	Prob.**
Levin, Lin & Chu t*	0.2360 1	0.5933

** Probabilities are computed assuming asymptotic normality

d) Null Hypothesis: Unit root (individual unit root process)

Series: CDSBGN, CDSHUN, CDSPOL, CDSRUS, CDSSER, CDSUKR
Date: 01/14/15 Time: 18:22
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 19
Total number of observations: 9501
Cross-sections included: 6

Method	Statistic	Prob.**
Im, Pesaran and Shin W-stat	-2.77327	0.0028

** Probabilities are computed assuming asymptotic normality

e) Null Hypothesis: Unit root (individual unit root process)

Series: CDSBGN, CDSHUN, CDSPOL, CDSRUS, CDSSER, CDSUKR
Date: 01/14/15 Time: 18:26
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 19
Total number of observations: 9501
Cross-sections included: 6

Method	Statistic	Prob.**
ADF - Fisher Chi-square	54.2077	0.0000
ADF - Choi Z-stat	-2.57448	0.0050

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

f) Null Hypothesis: Unit root (common unit root process)

Series: CDSBGN, CDSHUN, CDSPOL, CDSRUS, CDSSER, CDSUKR
Date: 01/14/15 Time: 18:24
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 0 to 19
Newey-West bandwidth selection using Bartlett kernel
Total number of observations: 9501
Cross-sections included: 6

Method	Statistic	Prob.**
Levin, Lin & Chu t*	0.66475	0.7469

** Probabilities are computed assuming asymptotic normality

g) Null Hypothesis: Unit root (individual unit root process)

Series: VIXBGN, VIXHUN, VIXPOL, VIXRUS, VIXSER, VIXUKR
Date: 01/14/15 Time: 18:33
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 2
Total number of observations: 9530
Cross-sections included: 6

Method	Statistic	Prob.**
Im, Pesaran and Shin W-stat	-4.86102	0.0000

** Probabilities are computed assuming asymptotic normality

h) Null Hypothesis: Unit root (individual unit root process)
Series: VIXBGN, VIXHUN, VIXPOL, VIXRUS, VIXSER, VIXUKR
Date: 01/14/15 Time: 18:34
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 2
Total number of observations: 9530
Cross-sections included: 6

Method	Statistic	Prob.**
ADF - Fisher Chi-square	48.1221	0.0000
ADF - Choi Z-stat	-5.12922	0.0000

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

i) Null Hypothesis: Unit root (common unit root process)
Series: VIXBGN, VIXHUN, VIXPOL, VIXRUS, VIXSER, VIXUKR
Date: 01/14/15 Time: 18:34
Sample: 5/04/2006 6/08/2012
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic selection of lags based on SIC: 2
Newey-West bandwidth selection using Bartlett kernel
Total number of observations: 9530
Cross-sections included: 6

Method	Statistic	Prob.**
Levin, Lin & Chu t*	-2.81335	0.0025

** Probabilities are computed assuming asymptotic normality

Annex 2:

Pedroni Residual Cointegration Test

Pedroni Residual Cointegration Test
Series: SS? CDS? VIX?
Date: 01/13/15 Time: 20:00
Sample: 5/04/2006 6/08/2012
Included observations: 1592
Cross-sections included: 6
Null Hypothesis: No cointegration
Trend assumption: No deterministic trend
Lag selection: fixed at 1
Newey-West bandwidth selection with Bartlett kernel

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

	Statistic	Prob.	Weighted Statistic	Prob.
Panel v-Statistic	28.64915	**0.0000**	11.19017	**0.0000**
Panel rho-Statistic	-28.67736	**0.0000**	-17.61534	**0.0000**
Panel PP-Statistic	-12.43802	**0.0000**	-8.880309	**0.0000**
Panel ADF-Statistic	-9.136756	**0.0000**	-7.728234	**0.0000**

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

	Statistic	Prob.
Group rho-Statistic	-26.71138	**0.0000**
Group PP-Statistic	-13.06744	**0.0000**
Group ADF-Statistic	-10.82616	**0.0000**

Kao Residual Cointegration Test

Kao Residual Cointegration Test
Series: SS? CDS? VIX?
hyDate: 01/13/15 Time: 20:06
Sample: 5/04/2006 6/08/2012
Included observations: 1592
Null Hypothesis: No cointegration
Trend assumption: No deterministic trend
Lag selection: fixed at 1
Newey-West bandwidth selection using Bartlett kernel

	t-Statistic	Prob.
ADF	-21.36444	0.0000
Residual variance	277.8210	
HAC variance	500.8972	

Augmented Dickey-Fuller Test Equation
Dependent Variable: D(RESID?)
Method: Panel Least Squares
Date: 01/13/15 Time: 20:06
Sample (adjusted): 5/08/2006 6/08/2012
Included observations: 1590 after adjustments
Cross-sections included: 6
Total pool (unbalanced) observations: 9537

Variable	Coefficient	Std. Error	t-Statistic	Prob.
RESID?(-1)	-0.181642	0.007573	-23.98541	0.0000
D(RESID?(-1))	-0.502852	0.008854	-56.79385	0.0000

R-squared	0.389329	Mean dependent var	0.070368
Adjusted R-squared	0.389265	S.D. dependent var	144.8971
S.E. of regression	113.2364	Akaike info criterion	12.29704
Sum squared resid	1.22E+08	Schwarz criterion	12.29854
Log likelihood	-58636.45	Hannan-Quinn criter.	12.29755
Durbin-Watson stat	2.171898		