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The Determinants of Country Risk in Eastern European Countries. Evidence from Sovereign Bond Spreads.*

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Abstract

The paper studies country risk in two Central and Eastern European countries - Bulgaria and Poland. The long run relationship between the yield differential (spread) of Eastern European national bonds (denominated in US dollars) over a US Treasury bond on one the hand and the country's fundamentals as well as an US interest rate on the other hand, is examined. The cointegrated VAR model is used.

First, the yield differentials are analyzed on a country by country basis to extract stochastic trends which are common for all bonds in a given country. Thereafter, the risk is disentangled into country and higher level risk. This paper is among the first ones which use time series data to study the evidence from sovereign bond spreads in Eastern Europe.

Keywords: Sovereign bonds spreads, Brady bonds, Cointegration.

JEL codes: F34, C32.

1. Introduction

The purpose of this paper is to study country risk in two Central and Eastern European countries (CEEC) - Bulgaria and Poland. When considering country risk, one should distinguish between default risk and currency risk. The former is the risk which arises from the fact that the national authorities can default on their debt - either principal or interest payments. The latter comes from the fact that the national currency can be devalued, and if the debt is denominated in national currency, its real value will be reduced. This paper concentrates on the more important element of country risk - the default risk.

This study focuses on the long-run determinants of default risk. The long-run relationship between the yield differential (spread) of Eastern European national bonds (denominated in US Dollars) over a US Treasury bond, on the one hand, and a country's fundamentals as well as an US interest rate (to account for the world market conditions), on the other hand, will be examined. The cointegrated vector auto regression (VAR) model will be used.

The econometric analysis proceeds in two stages. In a *first stage* the yield differentials are analyzed on a country-by-country basis. In this stage, stochastic trends shared by all yield differentials in a given country are extracted as potential components of country risk. These trends, however, may also reflect regional (Central and Eastern European) and common emerging market risk. At a *second stage* an attempt is made to disentangle country and higher-level risk. The extracted country stochastic trends will be analyzed to check, first, how they are related to various macroeconomic fundamentals (country risk), and second, whether they have common components across countries (regional and market risk). The idea is to represent the I(1) component of risk as the sum of a linear combination of domestic variables, and a residual I(1) component which would represent the international risk. In the simplest case, in which domestic variables are not cointegrated across countries, the former linear combination can be interpreted as a part of country risk, while regional and market risk will be in the residual. If the chosen domestic variables exhaust all domestic sources of risk, the residual will contain only regional and market risk. Section 4.2 provides a more detailed discussion of this and other cases.

Some problems arise from the fact that most of the Eastern European countries started issuing debt certificates denominated in foreign currency only recently, during the past three to four years. Bulgaria and Poland are the exceptions. These countries issued bonds in early 1994 which were denominated in US dollars in terms of the framework of their debt and debt service reduction agreement under the Brady Plan[1]. Therefore, for reasons of data availability, the analysis will be applied to these two countries only.

The paper proceeds as follows. Section 2 summarizes the empirical literature on sovereign bond spreads and default risk. Section 3 provides an explanation of Brady Bonds as well as on the way of measuring of sovereign bond spreads. The

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1 See Section 3 for an explanation of Brady Bonds.

conceptual framework is explained in Section 4 and the results of the empirical analysis are reported in section 5 and in section 6. The paper concludes with a summary.

2. The literature

There have been some recent and systematic studies on the yield spreads on EMU government bonds (see, for example, Codogno, Favero and Missale 2003), however there have been only a few systematic studies on pricing of sovereign country bonds and the risk of default in emerging markets and this despite the explosive growth of emerging market debt.

The existing studies relate the pricing decision of emerging market debt to both national and international factors. To the national factors belong a series of different macroeconomic fundamentals such as per capita income, GDP (growth rate), the rate of inflation, fiscal and external balances, foreign debt, the real (effective) exchange rate. The default history as well as economic development are also considered. The latter factors, however, are difficult to quantify. Below is a brief explanation of the variables used in the economic literature on sovereign bond spreads (see, for example, Cantor and Packer (1996)).

Foreign debt: a higher debt burden corresponds to a higher risk of default.

Foreign reserves: the weight of the debt burden increases as a country's foreign currency debt rises relative to its foreign currency reserves (and earnings).

Fiscal balance: large budget deficits absorb private domestic savings and suggest that the government might be unwilling to tax its citizens in order to cover current expenses or to service its debt.

External balance (current account balance): a current account deficit is a sign of foreign indebtedness (which if large could become unsustainable over time). It indicates that the public and private sectors rely on foreign financing.

Real exchange rate: a real appreciation of the national currency might endanger the country's exports and worsen the current account balance of the country and hence the stream of foreign currency inflows.

Inflation: high inflation is a proxy for structural problems in the government's finance. It might be a sign that the government uses the inflation tax because it is unable or unwilling to collect taxes and/or to issue further debt. Furthermore, the fact that the population is dissatisfied with inflation might cause political instability.

Cantor and Packer (ibid.) analyze the determinants of spreads in 49 countries during the year 1995. They relate the yield differentials to per capita income, GDP growth, inflation, the fiscal balance, the external debt, to indicators of economic development and default history and to the average Moody's and Standard and Poor's country credit ratings. The authors, however, use credit ratings, which are discrete variables, in order to explain the risk of default, which is a continuous one. We believe that credit ratings are an imperfect post-substitute for macroeconomic variables. Therefore in the present study the use of the credit ratings is renounced.

Among the most commonly cited international factors which influence the pricing decisions of emerging market bonds and default risk are the interest rates of the advanced countries and particularly of the U.S. The yield differentials tend to move in the same direction as changes in the U.S. interest rates (see IMF (2000)). Periods of tighter liquidity conditions in the advanced countries have tended to be associated with widening sovereign spreads and lower capital flows into the emerging market countries. In 1994, for example, tightening in US monetary conditions was accompanied by widening in emerging market bond spreads. In 1998, the easing of the US monetary policy helped to reduce somewhat emerging market bond spreads. Eichengreen and Mody (1998a) analyze a dataset of new bond issue spreads and estimate a model to explain simultaneously both the probability of issuing emerging market bonds and the yield differentials. In their analysis they use the yields of the 10-year US Treasury bond rates as a proxy for U.S. monetary policy. The authors find that the decline in the 10-year US Treasury bond rate led to an increase in the issuance of bonds by emerging market countries. However, it should be mentioned that their analysis was based on data on newly issued bonds and not on bonds which were actually traded on the secondary markets.[2] Moreover, their analysis is based on two subperiods - 1991-93, a period when the market for sovereign debt was still not well established, and 1994-95, a period when access to the market was seriously restricted for lower quality issuers.

[2] See also Eichengreen and Mody (1998b) for the impact of U.S. interst rates on determining the volume of bond issuance.

Cline and Barnes (1997) find a positive effect of U.S. Treasury yields of sovereign spreads in selected emerging market countries during the mid-1990s (the authors use the 3-month US Treasury bill rate), a finding which was also confirmed by the results of Kamin and von Kleist (1999). The latter authors also use data on newly issued bonds and launch spreads and find that variations in industrial country short-term interest rates explain relatively little of the decline in the emerging market bond spreads. When launching spreads (instead of secondary market spreads) are used, however, the latter will be observed only when positive decisions to borrow and lend are made. Therefore, the present analysis concentrates on secondary market bond spreads only.

Arora and Cerisola (2001) also conducted a study on the influence of U.S. monetary policy on emerging market bond spreads. Their model tries to explain the fluctuations in bond spreads as a function of country specific macroeconomic variables, the level of the U.S. Federal Funds target rate and a proxy for volatility on the capital markets. The latter is intended to capture changes in investor sentiment that may be related to expected changes in U.S. monetary policy. The results suggest that the level of the U.S. Federal Funds target rate has a significant positive effect on emerging market spreads with an estimated coefficient ranging from 1/2 to 1. The econometric analysis also suggests that a significant proportion of the fluctuations in emerging market spreads is driven by country-specific fundamentals. In general, improved macroeconomic variables such as higher net foreign assets, lower fiscal deficits and lower ratios of debt service to exports and debt to GDP help to lower sovereign spreads across countries.

A second international factor cited in the economics literature derives from the possible link between stock prices in developed countries and gross private financing to emerging markets and hence on emerging market bond spreads. This judgement is based on the high correlation observed between emerging market bond spreads and stock returns in the Dow Jones indices. When stock market volatility rises and greater uncertainty motivates investors to reduce their overall risk exposure, they will first reduce holdings of the most liquid class of assets, among which is also emerging market debt (see IMF (2000)).

The vast majority of the above mentioned studies were performed using panel data models and data on emerging market bond spreads. Typically they relate the pricing of international bonds (the spreads between their yields and the yield of a benchmark fixed asset) to a vector of a country's characteristics. Some studies include interest rates in the advanced industrial countries and others credit ratings, but all of them suffer from one problem: not all potential issuers will be in the sample at every point in time. Therefore, the conditions for a consistent estimation between characteristics and spreads are unlikely to be met in practice.

Dungey, Martin and Pagan (2000) decomposed bond rate spreads into national and global factors. These factors are assumed to have GARCH-type presentations and serial dependence. The authors use weekly data from January 1991 till April

1999 for Australia, Japan, Germany, Canada and UK, issued in national currencies. They assume, however, that foreign exchange returns and spread return share the same factor model, which in our view is a strong assumption. Furthermore, they treat spreads as stationary variables. The results suggest that the world factor has the dominant influence on Australia and Canada, whereas Japan, Germany and UK exhibit individual country effects.

The main contribution of this paper is that it is among the first to link the pricing of sovereign bonds to economic fundamentals as well as the risk of default using time series data and cointegration analysis. Since the yield differential time series were found to be non-stationary (see also Barbone and Forni (1997)) standard asymptotic theory does not hold and OLS estimates can be misleading. The problem of spurious regression could arise if standard regression is used. The problem is faced using the concept of cointegration: a group of non-stationary times series is cointegrated if there is a linear combination of them which is stationary. Usually the cointegrating vector is interpreted as a long-run equilibrium relationship.

3. Yield Calculations and Spread to Measure Country Risk

This section describes the data used in the empirical research. Brady bonds were first issued by Bulgaria and Poland in 1994 and hence there are enough observations for an empirical (time series) research to be conducted. The bonds are denominated in foreign currency (the US dollar) thus allowing the research to concentrate on the estimation of default risk, without mixing it with currency risk. According to the Brady plan (Nicholas Brady was a former US Treasury Secretary) defaulted sovereign bank loans of some less developed countries and former communist countries were written down and converted into bonds (see Bauer 2002). The Brady agreement allowed the debtor countries to exchange foreign debt for tradable foreign currency denominated fixed income securities at lower interest rates. The agreements were implemented upon promises for structural reforms to be undertaken by those countries under the supervision of international financial institutions, the International Monetary Fund and the World Bank.

Both Bulgaria and Poland had defaulted on their debt and reached an agreement with their bank creditors in 1994. According to these agreements the bank creditors were given the option of choosing between debt or debt-service reduction. The first option was an exchange for *Discount bonds* with a reduction of the face value of the defaulted loans. According to these options, however, the bonds had to pay an above-market coupon rate. A second option was the issue of *Par bonds* which were characterized by the fact that they bore a below-market coupon. The Par bonds had no reduction of the face value. Bulgaria and Poland also issued a few types of Brady bonds which covered the past due interest. These were the Bulgarian FLIRBS (Front loaded interest reduction bonds), IABs (interest arrears bonds) and the Polish RSTA (Revolving Short Term Agreement) and PDI (Past due interest) bonds.

Whereas Par, Discount and RSTA bonds were collateralized with specially issued zero coupon US Treasury securities, the FLIRB, IAB, PDI bonds are left uncollaterized.

Bulgaria issued three kinds of Brady bonds: 3,700.02 million US Dollars in *Bulgaria Discount* bonds with a floating coupon and 30 years maturity. The face value of the bonds was below the face value of the outstanding loans, but the coupon rate was a Libor + 81.25 bps (for tranche A, 50% of all discount bonds) and Libor +1,3125% (for tranche B, 50% of all discount bonds). Bulgaria also issued a Front-loaded interest-reduction bond (FLIRB) with 18 years maturity and a step-coupon (the issued amount was USD 1,657.46 million) as well as a IAB bond with 17 years maturity and a coupon of 6 months Libor + 81.25 bps (the issued amount was USD 1,610.01 million). Poland issued 4 types of Brady bonds: Poland Par (USD 930 million), Poland RSTA (USD 900 million), Poland Discount (USD 3,000 million) and Poland PDI (USD 2,650 million). The maturities of all Polish bonds but the last one is 30 years. The Poland PDI bond has a maturity of 20 years. The Bulgarian and Polish Brady bonds are summarized in Table 1.

Wherever collateralized bonds were considered in the following econometric analysis, it was the stripped yield rather than the yield to maturity which was used in the calculations. The stripped yield is the one obtained when "stripping" the bond from the collateral backing given by the US securities. The spread (yield differen-

Table 1: Bulgarian and Polish Brady Bonds.

Bond	**Amount issued** in million USD	**Maturity** in years
Bulgaria Disc (A)	1,850.36	30
Bulgaria Flirb	1,657.46	18
Bulgaria IAB	1,610.01	17
Poland Par	930	30
Poland RSTA	900	30
Poland Disc	3,000	30
Poland PDI	2,650	20

tials) are calculated considering a theoretical yield for a US benchmark bond that has exactly the same life as the sovereign bonds.

4. Conceptual Framework

As was specified in the introduction, the empirical analysis is carried out in two stages. In the first stage, yield differentials are analyzed country by country, and risk measures are constructed. The second stage is in two steps. In the first step, the long-run relations between the risk measure and macroeconomic fundamentals are examined for each single country. In the second step, cross-country differences are studied.

4.1 Stage 1: Stochastic Trends in Country Spreads

Since the purpose of the present paper is to study the long run determinants of country risk, tools for extracting information on the unobservable risk variable from the observable yield differential series are discussed here. The framework of the cointegrated VAR (CVAR) model is adopted.

Consider the case in which all the yield differentials in a country data set are I(1). The following operational convention is maintained: the stochastic trend of the unobservable country-risk variable should be shared by all yield differentials for that country, and should not be shared by the yield differentials of other countries. Thus, stochastic trends specific to individual yields or to groups of yields, but not to all of them, only qualify for components of bond-specific risk, but not of country risk. On the other hand, stochastic trends common to yield differentials of different countries are attributable to regional rather than to single-country risk.

Analyzing the yield differentials of all countries simultaneously may pose a dimensionality problem. To avoid this, the study is conducted on a country-by-country basis. In the first stage of analysis, only the first part of the convention will be applied, according to which a country-risk trend is necessarily common to all yield differentials for that country. In this stage it will not be possible to distinguish country-specific risk from higher-level risk.

Let $Y_t = (y_{1t}, ..., y_{nt})'$, $t = 1, ..., T$, be the vector series of yield differentials for an arbitrary country, and let it be well described by a VEC model:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-1} + \varepsilon_t,$$

where the standard assumptions of Johansen (1995) hold, and the notation from the same book is used. The moving average (MA) representation of Y_t is given by

$$\begin{aligned} Y_t &= \tilde{\beta}_\perp \alpha'_\perp \sum_{i=1}^{t} \varepsilon_i + stationary\, term + initial\, values \\ &= \tilde{\beta}_\perp \alpha'_\perp \sum_{i=1}^{t} \varepsilon_i + I(0), \end{aligned}$$

with $\tilde{\beta}_\perp = \beta_\perp (\alpha'_\perp \Gamma \beta_\perp)^{-1}$, $\Gamma = I - \Gamma_1 - ... - \Gamma_k$. The $n-r$ stochastic trends, where r is the cointegrating rank, are defined as $s_t = \alpha'_\perp \sum_{i=1}^{t} \varepsilon_t$, and for practical purposes can be approximated by $\alpha'_\perp \Gamma Y_t$. The $\tilde{\beta}_\perp$ matrix contains the loadings of the stochastic trends. A column of $\tilde{\beta}_\perp$, all the elements of which are non-zero, corresponds to a trend which feeds into all yield differentials, i.e. which represents risks at country or higher level according to the accepted convention, and should therefore be extracted at the first stage of analysis[3].

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[3] There are two complications in this setup. The first one is that $\tilde{\beta}_\perp$ and $\alpha_\perp$ are not identified even if α and β are.

The second complication concerns inference on α and Γ_i, and indirectly, on $\alpha'_\perp \Gamma$. Although the unrestricted estimates $\hat{\alpha}$ and $\hat{\Gamma}_i$ delivered by Johansen's procedure are known to be consistent, in finite samples restricted estimates may be preferable for efficiency reasons. Anticipating the empirical results to come later, it can be pointed out that in the present study tests of restrictions on α and Γ_i may be difficult to conduct, as the following argument demonstrates. When there are common factors influencing the change in the individual yield differentials from period to period, which is quite likely to be the case, the components of ε_t will exhibit high contemporaneous correlation. If, for ease of exposition, one common factor is assumed to exist, ε_t may have the representation $\varepsilon_t = cf_t + \eta_t$, where c and η_t are respectively a constant and a stochastic $n-$vector, and f_t is a random variable. Then $Var(\varepsilon_t) = cc'\sigma_f^2 + \Sigma_\eta$, if f_t and η_t are independent. If the common factor dominates the idiosyncratic factors in the sense that it has a significantly higher contribution to $Var(\varepsilon_t)$ than η_t has, then $Var(\varepsilon_t)$ will be close to a reduced rank matrix, as it will be dominated by $cc'\sigma_f^2$, which is obviously of reduced rank. The asymptotic distribution of $\hat{\alpha}$ is given by

$$\sqrt{T}(\alpha - \hat{\alpha}) \xrightarrow{d} N\left(0, Var(\varepsilon_t) \otimes Var\left(\beta' X_t\right)\right),$$

(see Johansen (1995, p.184)), and a similar result holds for $\hat{\Gamma}_i$. When $Var(\varepsilon_t)$ is close to a reduced rank matrix, inference on α is problematic, since the estimates of α_{ij} are highly correlated. For example, it may happen that individual tests for weak exogeneity fail to reject for all the variables, while joint tests reject for all the pairs of variables, while in fact none of the variables is weakly

We concentrate on the case of $r = n-1$[4]. In this case there is a single stochastic trend driving all yield differentials. When $r = n - 1$, any weighted average of the components of Y_t contains the same stochastic trend as $\alpha'_{\perp}\Gamma Y_t$. Indeed, from the MA representation, it follows that $w'Y_t = as_t + I(0)$, where s_t is the stochastic trend, $a = \sum_i w_i \left(\tilde{\beta}_{\perp}\right)_i$ and $w = (w_1, ..., w_n)'$ is the vector of weights. The measure $w'Y_t$ can be constructed by using some economically meaningful weight vector[5] instead of $\hat{\Gamma}'\hat{\alpha}_{\perp}$.

4.2 Stage 2: Comparative Analysis of Countries

In this stage, for each country we analyze to what extent the permanent component of the risk variable $w'Y_t$, extracted at stage 1, is related to the permanent components of several macroeconomic indicators and what are the short-run adjustment mechanisms behind the long-run relations. Cointegrants such as government cash balance (as per cent of GDP) (Govbal), real effective exchange rate (Reer)[6], reserves of the national bank (Resbnb), monetary base, current account balance (CAB) (also as percent of GDP) and 10 year US (or German) Treasury Bond rate are considered. These

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exogenous. Therefore, a good restricted estimate of α, and similarly of Γ, may be hard to obtain.

[4] This is the case in which the two complications, explained earlier, can be overcome.

[5] It is essential, however, that the weight vector be *constant in time*. Otherwise $w'_tY_t = a_ts_t + I(0)$ with time-variable a_t (unless the components of $\tilde{\beta}_{\perp}$ are equal), and it is not clear whether a cointegrating relation with fixed coefficients between w'_tY_t and other variables will exist.

[6] The series is taken from the International Financial Statistics of IMF.

variables are supposed to measure sustainability and the solvency of the country and its standing with regard to the rest of the world.[7]

The choice of data frequency is conditioned by the short period since CEE bonds are in the market: no usable sample sizes can be achieved at lower-than-monthly frequencies. Admittedly, monthly yields can be expected to be moderately influenced by the long-term development of the economy, while some conjunctural factors may have stronger influence (Juselius (2006), forthcoming). The data analysis will be carried out with this restriction in mind[8].

Consider two countries, A and B, with risk measures $r_t^i = w_i' Y_t^i$, $i = A, B$. As discussed in section 4.1, they have the representation $r^i{}_t = a^i s_t^i + I(0)$, where s_t^i are stochastic trends capturing country and higher-level risk. An attempt will be made to distinguish between these two sources of risk. Let x_t^i and z_t^i be two cointe-

[7] Another variable used in the literature is the foreign debt to GDP ratio. Unfortunately for Bulgaria neither monthly nor quarterly reliable data on its foreign debt can be found and therefore we were forced to exclude it from the analysis.

The Government Cash Balance is defined as revenues minus expenditures and it is taken as percentage of GDP.

Finally, the quarterly GDP series were interpolated using the *interpol* procedure in the RATS Software assuming that the DGP is a random walk with drift.

[8] The CVAR approach assumes that the above-listed variables can be approximated by I(1) processes. Since economic logic prevents variables like CAB and Govbal (taken as a percentage of GDP) from having unbounded variances, the assumption needs further clarification. First, here the order of integration is understood as a sample-specific statistical property, and not as a property inherent to the data for all samples and periods. If a data series is statistically better described as non-stationary rather than as stationary, then a non-stationary model is to be preferred in order to ensure the validity of the inferential framework. Second, non-stationarity is understood as a local property, in the sense that although within the sample range the data behave as generated by a random walk, outside that range there could be thresholds above which reversion takes place (see Bec and Rahbek (2003)). For example, the CAB to GDP ratio may behave as a random walk between -5 and 5 percent, while reversion to this interval takes place when the variable leaves it. If the observed data is in fact between -5 and 5 percent, the random walk model is the preferable one, first, because it is a good approximation to the observed dynamics, and second, because there is little information in the data to estimate a model with reversion.

grants for each of the countries. In the actual analysis vectors of cointegrants will be used, but scalars will suffice for illustrative purposes. It is assumed that at least one cointegrating relation exists between r_t^i, x_t^i and z_t^i for both countries and that it can be normalized to $r(t,i)$.

Case 1 *Cointegration rank* $r_i = 1$, $i = A, B$. In this case two stochastic trends per country are present, say τ_{1t}^i and τ_{2t}^i.

Case 1.1 The trend τ_{1t} is common to both countries, while τ_{2t}^i are distinct. In this case τ_{2t}^i can be interpreted as domestic sources of risk, and τ_{1t} can be interpreted as international risk. The MA representation of the data is

$$\begin{pmatrix} r_t^i \\ x_t^i \\ z_t^i \end{pmatrix} = \begin{pmatrix} c_{11}^i & c_{12}^i \\ c_{21}^i & c_{22}^i \\ c_{31}^i & c_{32}^i \end{pmatrix} \begin{pmatrix} \tau_{1t} \\ \tau_{2t}^i \end{pmatrix} + I(0),$$

and in particular, $r_t^i = c_{11}^i \tau_{1t} + c_{12}^i \tau_{2t}^i + I(0)$ is decomposed into its domestic and international permanent components. The empirical analysis can proceed in the following way. First, cointegration analysis can be performed for each country in order to estimate the long-run relation between risk and the macro variables at country level. Two stochastic trends per country can be extracted. In general, these will be linear combinations of τ_{1t} and τ_{2t}^i, and can be expressed in terms of (r_t^i, x_t^i, z_t^i). Second, cointegration analysis among the extracted stochastic trends can be performed. One cointegration vector can be expected to be found, representable as $\left(r_t^A, x_t^A, z_t^A\right)\beta_A - \left(r_t^B, x_t^B, z_t^B\right)\beta_B = I(0)$.

From a common-trends prospective, up to proportionality, both $\left(r_t^A, x_t^A, z_t^A\right)\beta_A$ and $\left(r_t^B, x_t^B, z_t^B\right)\beta_B$ will have the international trend τ_{1t} as their permanent component.

The data analysis in this paper will be organized according to the two steps described above, although different conclusions on the cointegration ranks may be obtained. In the two-country six-variables example the analysis of the full system can be performed instead of several partial analyses, but this may not be feasible for higher-dimensional systems, and the two-step procedure is therefore relevant.

Case 1.2 Both trends τ_{1t} and τ_{2t} are common to both countries. In this case the macro variables are necessarily cointegrated across countries. The two steps of empirical analysis can be as in case 1.1. However, two cointegration vectors should obtain at the second step. One of them will be representable as $\left(x_t^A, z_t^A\right)\beta_A^1 - \left(x_t^B, z_t^B\right)\beta_B^2 = I\left(0\right)$, and the permanent component in either $\left(x_t^A, z_t^A\right)\beta_A^1$ or $\left(x_t^B, z_t^B\right)\beta_B^2$ could be called fundamental, in the sense that it is related to the macro fundamentals of the countries. It will also be a permanent component of r_t^i, and although for country i it will be representable in terms of the fundamentals of that country, it will not be strictly interpretable as domestic, since it will also be representable in terms of the fundamentals of the other country.

Case 1.3 Both trends τ^i_{1t} and τ^i_{2t} are country-specific. This means that there are two domestic sources of risk, and there is no shared risk between the two countries, i.e. r^i_t represents country risk alone. The first step of analysis is as in case 1.1, while at the second step no cointegration vectors should obtain.

Case 2 *Cointegration rank* $r_A = 2$. This implies that the macro variables for country A are cointegrated among themselves and for the purposes of risk analysis one of them can be eliminated. It would be best to find another set of macro variables so that the situation of case 1 is restored. If this is not done, the analysis will not be particularly conclusive, as discussed below. The reason is that the stochastic trend $a^A s^A_t$ in r^A_t cannot be decomposed.

Case 2.1 At least country B is in the situation of case 1, i.e. $r_B = 1$. Very tentative conclusions could be drawn. In this case $r^B_t = c^B_{11}\tau^B_{1t} + c^i_{12}\tau^B_{2t} + I(0)$, and τ^B_{1t} can be identified in such a way that it be a linear combination of x^B_t and z^B_t (up to an $I(0)$ process). If r^A_t happens to cointegrate with τ^B_{1t}, this could be attributed to unobserved cointegration between x^B_t and z^B_t as well as some fundamental variable y^A_t, which is not in the information set but drives r^A_t. Cointegration between r^A_t and a general linear combination of τ^B_{1t} and τ^B_{2t} may mean that, in the long run risk for country A is driven by regional and common emerging-market factors rather than by domestic fundamentals. However, no cointegration between r^A_t, τ^B_{1t} and τ^B_{2t} will obtain if there is a fundamental variable y^A_t, not cointegrated with the fundamentals of country B, and which contributes to the permanent component of r_t^A. The latter possibility is not an unlikely one.

Case 2.2 $r_B = 2$. Similarly to r_t^A, the permanent component of r_t^B is not decomposable. Now cointegration between r_t^A and r_t^B is possible in two, not very likely, occasions (and in mixtures of these). The first occasion is when both risk variables are only driven by a common market trend. On the second occasion they can be driven by country fundamentals, which are cointegrated across countries and whose combined effect on the risk variables is the same for the two countries.

Case 2 reiterates the importance of the information set for the outcome of the analysis. In general, when more than two macro variables are analyzed, several country-specific as well as several trends shared among countries can be present. The number of shared trends can be greater than one if and only if the macro variables are cointegrated across countries. Since one may expect variables with the same meaning to be cointegrated (GDP^A with GDP^B etc.), it might be useful to test for such cointegration separately.

5. The Bulgarian Case

5.1 Stage 1: Analysis of Yield Differentials

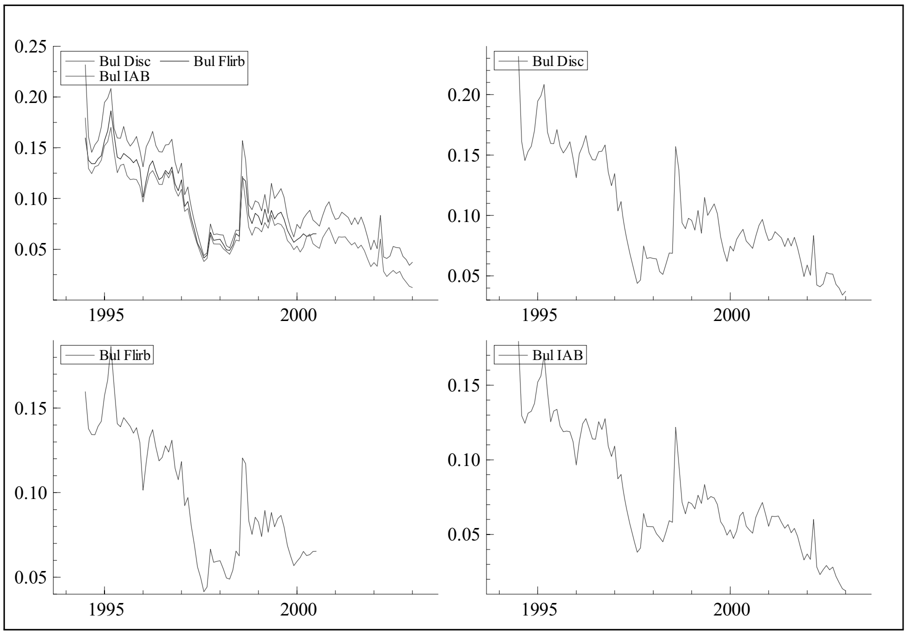
This section provides evidence in support of the conjecture that the Bulgarian yield differentials (see Figure 1) are well-modelled as I(1) variables with a single common stochastic trend. Namely, the stationarity hypothesis is rejected, within appropriate CVAR models[9], for all the three yield differentials (of Bulgarian Brady bonds over US Treasury bonds, called Bul Dics, Bul Flirb and Bul IAB),[10] and it is shown that only one common stochastic trend is present in the data. According to the discussion in Section 4.1, the common unit root process proxies the permanent component in the Bulgarian Brady bonds.

Due to data unavailability for bond Bul Flirb after August 2000, two VAR models are estimated. The first one covers the period from September 1994 to August 2000 and models the dynamics of all the three spreads. For this model it is tested that the cointegrating rank equals 2 and that none of the spreads is stationary. The second model covers the longer period from September 1994 to January 2003, and is a bivariate model for bonds Bul Disc and Bul IAB. Its cointegrating rank is one, and the cointegration relation between these two yield differentials appears to be the same as in the trivariate model over the shorter sample. Thus the analysis of both models favours the single driving trend conjecture.

[9] ADF tests were also performed, the results of which confirm the rejection of stationarity (see Table 16).

[10] The data is shifted backwards, so that the observation corresponding to the first day of a month, as provided by Datastream, is treated as end-of-month observation for the previous month.

Figure. 1: Bulgarian Yield Spreads over US Treasury Securities



The trivariate model is discussed first. A VAR with two lags, constant, and both a step and a blip dummy for August 1998, is estimated. The lag length is chosen to be the shortest one that makes the residuals serially uncorrelated. In the cointegration analysis the constant and the blip dummy are left unrestricted, while the step dummy is restricted to the cointegration space. This allows for a level shift in all directions in August 1998 . The dummy variables account for the Russian crisis in 1998. The coefficient of the blip dummy, for example, varies between 5.7 and 6.9 times the standard deviation of the respective equations, which is big relative to $\sqrt{\text{sample size}} = \sqrt{72} \approx 8.5$. Doornik et.al. (2002) argue that in such cases dummies should be included in the model specification since cointegration tests can be misleading otherwise.

Table 14 summarizes some misspecification tests which show that the model is a statistically good description of the data, and therefore, a valid inferential framework[11].

Table 2: Misspecification Tests for a Trivariate System of Bulgarian Brady Bonds.

	AR 1-5 test [p-value]	**AR 1-3 test**	**Jarque-Bera**	**ARCH 1-5**
Bul Flirb	0.46 [0.80] in $F(5{,}58)$	0.65 [0.59] in $F(3{,}60)$	0.09 [0.96] in $\chi^2(2)$	0.87 [0.51] in $F(5{,}53)$
Bul IAB	1.01 [0.42] in $F(5{,}58)$	1.67 [0.18] in $F(3{,}60)$	0.74 [0.69] in $\chi^2(2)$	0.38 [0.86] in $F(5{,}53)$
Bul Disc	0.68 [0.64] in $F(5{,}58)$	1.15 [0.33] in $F(3{,}60)$	0.86 [0.52] in $\chi^2(2)$	0.86 [0.52] in $F(5{,}53)$
System Test	0.97 [0.53] in $F(45{,}137)$	1.02 [0.45] in $F(27{,}152)$	10.22 [0.12] in $\chi^2(6)$	

Among the estimated roots of the companion matrix, only one seems to be close to unity (the three roots with largest real part are displayed in Table 3), which is consistent with cointegration rank equal to 2. The trace test provides the same conclusion. The hypothesis of stationarity, possibly around a changing mean, is rejected for all the three yield differentials (the test is conducted on the β-matrix, with tests statistics 15.11 [0.00]**, 15.32 [0.00]** and 14.75 [0.00]** respectively for Bul Flirb, Bul Disc and Bul IAB; the p-values are from the χ^2 (1) distribution). These tests confirm the results of the ADF tests of the univariate spreads series, as shawn in Table 16 (see appendix). Thus, the data complies with the setup from section 4.1: all yield differentials are I (1), and non-stationarity is introduced by a single

[11] These and the following calculations were performed with the software packages GiveWin and PcGive.

stochastic trend, interpretable as the permanent component of the unobservable risk variable.

Table 3: Trace Test Statistics for the Bulgarian Trivariate System

real part	**imaginary part**	**modulus**	$H_0 : r \leq$	Trace test	[P-value]
0.96	0	0.96	0	48.67	[0.00]**
0.47	0.12	0.48	1	19.05	[0.01]*
0.47	-0.12	0.48	2	1.68	[0.20]

Furthermore, pairwise cointegration between the yield differentials does not occur with (1, -1) vectors: (1, -1,*) between Bul Disc, Bul IAB and DS988 is rejected with test statistic of 15.156, [0.00]** in $\chi^2(1)$, and similarly, for Bul Flirb, Bul IAB and DS988 the test statistic is 7.76, [0.01]**. According to the concluding paragraph of section 4.2, the cointegration coefficients between a risk variable constructed with weighting vector w, and macro fundamentals, will depend on the choice of w.

An exactly identified $\hat{\beta}-$ matrix is provided in the leading columns of Table 4. It can be seen that pairwise cointegration occurs in terms of pseudo differences (t-statistics are given under the estimates). According to the argument in section 4.1, inference on the loadings is problematic, given that the residuals are highly contemporaneously correlated: the correlation matrix estimated from the unrestricted VAR is

$$\begin{pmatrix} 1 & & \\ 0.92 & 1 & \\ 0.94 & 0.92 & 1 \end{pmatrix}$$

Therefore, tests of correlated hypotheses lead to inconsistent conclusions. For example, weak exogeneity is not rejected for either Bul Disc and Bul IAB (respectively 0.516 [0.77] and 1.86 [0.39] in χ^2 (2)), while joint weak exogeneity is rejected.[12] Indeed, one variable at most can be weakly exogenous in the presence of a single unit root. To avoid the difficulties posed by similar inconsistencies, inference on α is not attempted. Mainly for presentational purposes, in the second panel of Table 4 estimates of β and α obtained under the restriction $\alpha_{11} = 0$ are shown. The estimates of the cointegration vectors are essentially unchanged, but the standard errors of the unrestricted elements of α are reduced due to the restriction (accepted with test statistic of 0.034, [0.85] in χ^2 (1)).

Table 4: Estimates of α and β for an Exactly Identified and for an Over-Identified Model of Bulgarian Yield Spreads.

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\alpha}_1$	$\hat{\alpha}_2$
Bul Disc	1		1			0.27 $_{0.67}$
Bul Flirb		-0.92 $_{-39.6}$		-0.92 $_{-39.5}$	0.37 $_{3.4}$	0.91 $_{2.8}$
Bul IAB	-1.32 $_{-36.4}$	1	-1.32 $_{-36.2}$	1	0.396 $_{4.3}$	0.29 $_{0.95}$
DS988	-0.012 $_{-4.7}$	0.005 $_{2.9}$	-0.012 $_{-4.7}$	0.005 $_{2.9}$		

A discussion of the bivariate model follows. Similarly to the trivariate one, it involves two lags, constant and the 1998 dummies, but possibly due to the smaller information set, more dummies are needed to improve the fit. A blip dummy in April

[12] Weak exogeneity is also rejected for Bul Flirb, 6.91 [0.03]* in χ^2 (2) .

1995, and a transitory +1,-1 dummy with 1 in March 2002 are added. Even then the specification is slightly worse, but still acceptable (compare results in Table 5).

Table 5: Misspesification test for a Bivariate Model of Bulgarian Brady Bonds.

	AR 1-6 test [p-value]	**AR 1-3 test**	**Jarque-Bera**	**ARCH 1-6**
Bul IAB	0.76 [0.60] in $F(6{,}86)$	1.38 [0.26] in $F(3{,}89)$	1.25 [0.54] in $\chi^2(2)$	0.69 [0.66] in $F(6{,}80)$
Bul Disc	0.41 [0.87] in $F(6{,}86)$	0.63 [0.60] in $F(3{,}89)$	2.39 [0.30] in $\chi^2(2)$	0.51 [0.80] in $F(6{,}80)$
System Test	1.47 [0.08] in $F(24{,}158)$	1.89 [0.04]$*$ in $F(12{,}170)$	8.06 [0.09] in $\chi^2(4)$	

For the bivariate system the results of the trace test suggests that the rank is 1 which is confirmed also by the roots of the companion matrix (see Table 6).

Table 6: Rank Determination for the Bivariate System of Bulgarian Brady Bonds.

real part	**imaginary part**	**modulus**	$H_0 : r \leq$	**Trace test**	**[Prob]**
0.96	0	0.96	0	16.29	[0.04]*
0.74	0	0.74	1	2.04	[0.15]
-0.03	0	0.03			

Stationarity analysis was also performed and the hypothesis of stationarity was rejected for both yield differentials (the test is conducted again on the β-matrix, with tests statistics 12.010 [0.0005]** and 12.196 [0.0005]** respectively for Bul Disc and Bul IAB; the p-values are from the χ^2 (1) distribution). Moreover, the cointegration relation between these two yield differentials appears to be the same as in the trivariate model over the shorter sample ($\hat{\beta}_1'$ is now (1,-1.27,-0.012) compared with (1,-1.32,-0.012) in the trivariate system).

In conclusion of this subsection it can be said that in the bivariate system the two yield differentials of bonds Bul Disc and Bul IAB are cointegrated and the cointegration rank is one. Therefore the analysis of both the trivariate and bivariate models favours the single stochastic trend interpretable as the permanent component of the unobservable risk variable.

5.2 Stage2: Analysis of Fundamentals

This section uses the Cointegrated VAR model to proceed with the analysis of risk extracted in the previous section, and relates it to macroeconomic variables as well as US (or German) interest rates.[13] The used variables are presented in Table 7 and are also plotted in Figure 2. The weights used in calculating the variable Spr from the yield differentials of the single Bulgarian Brady bonds are taken according to the issued amount for each of the two bonds for which data is available (Bul IAB 47% and Bul Disc 53%). Cointegrants such as governmental balance (Govbal), (log of) the real exchange rate (Lreer), (log of) the reserves of the national bank (Lresbnb), (log of) the monetary base (Lnm0), the current account balance (Cab) and the 10 year US Treasury Bond rate are considered. These variables measure the sustainability and

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[13] The following sections present the results of the econometric analysis when US interest rates were used. Since Bulgarian Currency Board used the Deutsche Mark (and since 1999 the Euro) as an ancor currency, the same analysis was performed also using the 10 years German Treasury Bond rate. All the results were confirmed and the same restrictions acceppted.

the solvency of the country and its standing with respect to the rest of the world. In the model weak exogeneity of the US (German) Treasury rate is assumed.

Table 7: Variables Used in the Stage 2 of the Econometric Analysis in the Case of Bulgaria.

Variable	**Explanation**
Spr	Weighted Spread of the yield differentials Bul Disc (53 %) and Bul IAB (47 %)
Govbal	Governmental Cash Balance (as percent of GDP)
Lreer	(Log of) the Real Effective Exchange Rate
Lresbnb	(Log of) the Reserves of the Bulgarian National Bank in USD
Lnm0	(Log of) Monetary Base (in BGN)
Cab	Current Account Balance (as percent of GDP)
Ustb10_1	(First Lag) of 10 years US Treasury Bond Rate
DS988	Step Dummy Variable (0 before August 98, 1 afterwards)
DT988	Blip (Impulse) Dummy Variable (1 in 8.98, 0 otherwise)

The period studied includes monthly observations from June 1997 until September 2002. During the hyperinflation period in 1996-1997 an explosive root is found in some of the data and therefore it was decided not to include these observations. A VAR with two lags of the endogenous variables, the first lag of the US interest rate, constant, centered seasonal dummies, and both a step and a blip dummy for August 1998, is estimated. The lagged rather than the current value of the US rate is used, so that the model can be regarded as a partial model with respect to a bigger one, in which an equation for the US rate is present. The dummies are the same as the ones used in the analysis of the yield differentials and account for the

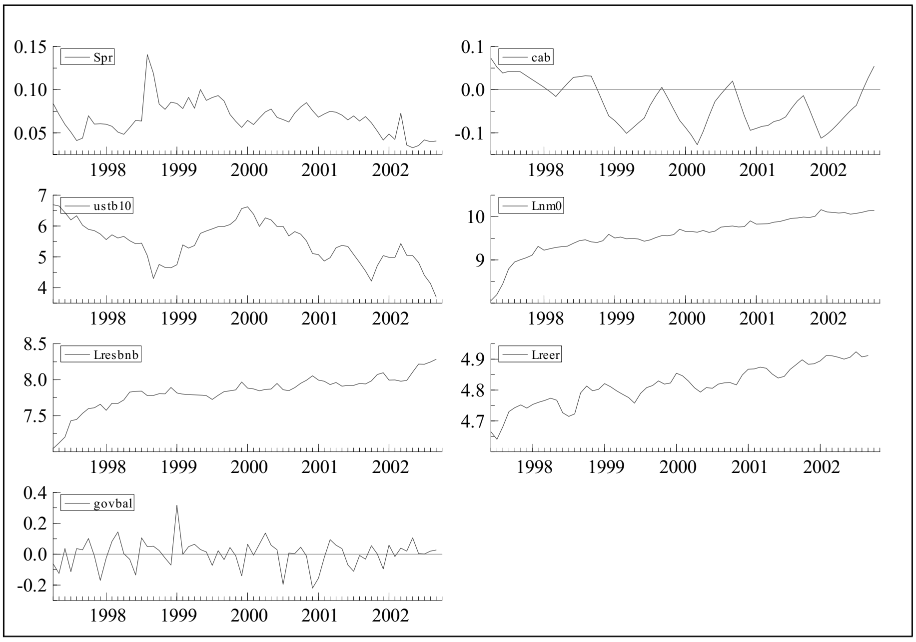

Figure 2: Variables used in Second Stage of the Econometric Analysis for the Case of Bulgaria.

Russian crisis in August 1998. The lag length again is chosen to be the shortest one that renders the residuals serially uncorrelated. In the cointegration analysis the constant and the blip dummy are left unrestricted, while the step dummy is restricted to the cointegration space. This again allows for a level shift in all directions in August 1998. The residuals seem well-behaved in terms of no serial autocorrelation (see Table 8), and also appear contemporaneously uncorrelated (with the single exception of the real effective exchange rate and the spread), as demonstrated by the estimated correlation matrix:

$$\begin{array}{c} Spr \\ Govbal \\ Lreer \\ Lres \\ Lnm0 \\ Cab \end{array} \begin{pmatrix} 1 & & & & & \\ -0.07 & 1 & & & & \\ -0.46 & -0.07 & 1 & & & \\ -0.09 & -0.05 & 0.03 & 1 & & \\ 0.08 & -0.24 & 0.13 & 0.25 & 1 & \\ -0.09 & -0.08 & 0.17 & 0.17 & 0.16 & 1 \end{pmatrix}.$$

The latter fact justifies labels such as ”shock to the current account balance”, ”shock to the government balance”, etc., to be attached to the single residuals.

The analysis proceeds in the following way: first, the cointegration rank of the system is determined and thereafter the cointegration space is identified. Overidentifying restrictions are imposed on the cointegration vectors and on the feedback coefficients. Finally, the common trends and their loadings are identified. After the completion of the technical tasks, conclusions are drawn from the obtained results.

Table 8: Univariate Diagnostics Tests for the VAR with 1 lag.

	AR 1-4, [p-value]	**ARCH 1-4**	**Jarque-Bera**
Spr	1.07 [0.39] in $F(4{,}33)$	0.11 [0.97] in $F(4{,}29)$	0.24 [0.89] in $\chi^2(2)$
Govbal	0.45 [0.77] in $F(4{,}33)$	0.13 [0.97] in $F(4{,}29)$	23.56 [0.00]** in $\chi^2(2)$
Lreer	0.27 [0.90] in $F(4{,}33)$	0.09 [0.99] in $F(4{,}29)$	26.40 [0.00]** in $\chi^2(2)$
Lresbnb	1.36 [0.26] in $F(4{,}33)$	0.73 [0.58] in $F(4{,}29)$	5.65 [0.06] in $\chi^2(2)$
Lnm0	0.49 [0.74] in $F(4{,}33)$	0.35 [0.85] in $F(4{,}29)$	3.36 [0.19] in $\chi^2(2)$
Cab	0.76 [0.56] in $F(4{,}33)$	0.44 [0.77] in $F(4{,}29)$	3.41 [0.18] in $\chi^2(2)$
System Test	1.13 [0.31] in $F(144{,}54)$	578.5 [0.36] in $\chi^2(567)$	56.03 [0.00]** in $\chi^2(12)$

Rank Determination

In this subsection the cointegration rank is estimated. The trace statistics suggest cointegration rank (r) is 4 (see Table 15). Note, however, that, as pointed out by Juselius (2006, forthcoming), in small samples the short-run dynamics cannot be assumed to have an insignificant effect on the distribution of the trace test. The author claims that when the sample size is below 100 (which is true also in the current analysis) the asymptotic tables should be used with caution. Therefore, because of the possible misestimation of the number of cointegration vectors further analysis is needed for a final judgement.

Additional information on the cointegration rank is provided by the estimated roots of the companion matrix. There is one unit root borne by the 10 year US Treasury Bond rate, assumed to be weakly exogenous. The unit roots in the remaining 6

Table 9: Trace Test Statistics for the Bulgarian Model.

H_0: r $\leq$	**Trace Test**	**P-value**
0	221.64	[0.000]**
1	133.42	[0.000]**
2	66.95	[0.000]**
3	32.88	[0.021]*
4	11.10	[0.209]
5	3.63	[0.057]

variables are counted below (see first column of Table 10). The companion matrix has as many unit roots as there are common stochastic trends, i.e. 6-r. The moduli of the two largest roots (whose real part is positive) are twice 0.91 indicating that all the roots are inside the unit circle. The estimated roots suggest that there exist 2 unit roots, leading to 4 cointegration vectors. If a rank is restricted to 5, one large unit root remains (visible in the second column in Table 10), suggesting again rank 4. Finally if rank 4 is imposed, fairly small roots remain, supporting this restriction (see third column of Table 10).

Table 10: Eigenvalues of the Companion Matrix for the Bulgarian Model.

Unrest. System			**Restr. System, r=5**			**Restr. System, r=4**		
real part	imag. part	mod.	real part	imag. part	mod.	real part	imag. part	mod.
0.91	0.06	0.92	1.00	0.00	1.00	1.00	0.00	1.00
0.91	-0.06	0.92	0.94	0.00	0.94	1.00	0.00	1.00
0.79	0.00	0.79	0.74	0.00	0.74	0.74	0.00	0.74
0.60	0.48	0.77	0.60	0.48	0.77	0.58	0.46	0.74
0.60	-0.48	0.77	0.60	-0.48	0.77	0.58	-0.46	0.74

For further evidence one should look at the significance of the coefficients of α. For a linear combination of the levels to qualify as stationary, there must be significant adjustment to it in at least one equation. Significance is tested with respect to Dickey-Fuller critical values for t-ratios, and the threshold of 3 (in absolute value) is used as suggested by Juselius (ibid.). The significance statistics of the coefficients of α (see the right hand panel of Table 11) suggests the presence of four cointegration relationships (for all four vectors the test statistics is greater than the critical value). If a fifth vector is admitted, no adjustment seems to occur.

Table 11: α and β Coefficients for a Just Identified Bulgarian Model.

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\beta}_4$	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\alpha}_3$	$\hat{\alpha}_4$
Spr	1.00	1.00	1.00	1.00	-0.52	-0.1	-0.08	0.03
(t-value)					$(\mathbf{-3.20})$	(-0.67)	(-1.33)	(0.67)
Govbal			0.40		1.28	-0.91	-3.01	-0.53
(t-value)			(3.48)		(0.94)	(-0.76)	$(\mathbf{-6.09})$	(-1.66)
Lreer	0.33			2.25	0.03	-0.46	0.07	- 0.30
(t-value)	(10.04)			(10.59)	(0.05)	(-0.94)	(0.34)	(-2.25)
Lresbnb		0.11			0.84	-2.23	-0.04	-0.12
(t-value)		(4.84)			(1.00)	$(\mathbf{-3.03})$	(-0.12)	(-0.61)
Lnm0			0.09	-0.35	- 0.16	0.21	-0.12	0.39
(t-value)			(6.27)	(-7.47)	(-0.30)	(0.44)	(-0.61)	(3.05)
Cab	0.28	-0.20			-0.70	0.31	0.04	0.12
(t-value)	(4.10)	(-1.92)			(-4.43)	(2.27)	(0.612)	$(\mathbf{3.14})$
DS988	-0.02	-0.04	-0.05	-0.09				
(t-value)	(-2.93)	(-2.94)	(-4.36)	(-3.04)				
ustb10_1	0.008	0.004	0.015	0.014				
(t-value)	(3.05)	(0.73)	(2.73)	(0.95)				

Thus, all the evidence supports r=4, and all further analysis is based on four cointegration vectors and hence two common stochastic trends. Table 11 presents the coefficients of the just identified model.

Tests of Overidentifying Restrictions

Tests of long-run exclusion (weak exogeneity) reject at the 5 percent level for all the variables. Also stationarity with a possible break in the mean in August 98 is rejected at the 1 per cent level for all stochastic variables but the Govbal, for which the hypothesis is rejected at the 10 per cent level. Thus, the budget balance is a borderline case and in Figure 2 it is seen to fluctuate a lot, similarly to a stationary variable. Some of the fluctuations are due to the presence of strong seasonality, but it should nevertheless be admitted that the root in the budget balance is quite likely to be smaller than unity. For the purposes of this analysis, it is approximated it by a unit root. In order to limit the impact of this decision on the conclusions of the analysis, the cointegration space is identified in such a way that the budget balance enters in only one cointegration relation. Thus, the interpretation of the remaining relations is not affected by this decision. In the interpretation of the budget relation itself it should be borne in mind that what will be called a permanent change in the budget balance is likely to be a change of shorter life compared to permanent changes in the other variables studied.

Three overidentifying restrictions are imposed on the cointegration space. First, since the stochastic trend borne by the 10-year US Treasure bond rate can be expected to be shared only by the spread variable, and since all cointegration relations are normalized to the spread, the coefficient of the 10 year US rate is restricted to be the same in all relations. Thus, the variable $spr - c * ustb10\_1$ can be defined as a new risk variable, and will be referred to as such in the following analysis. The correction with $ustb10\_1$ may mean that the original spread is not calculated w.r.t. the best benchmark. Second, the current account balance is restricted not to enter the second relation, as its estimated coefficient seems not to be significant. Finally, the logged real effective exchange rate and the current account balance are restricted to enter the first relation with equal coefficients. The latter restriction is not crucial for the interpretation. None of the restrictions is rejected individually, and the joint test accepts too (3.41, p-value=[0.64] in χ^2 (5)).

Next, insignificant feedback coefficients are restricted to zero (18.03, p-value= [0.45] in χ^2 (18)) and the structure of Table 12 is obtained.

Table 12: α and β Coefficients for an Over-Identified Bulgarian Model.

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\beta}_4$	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\alpha}_3$	$\hat{\alpha}_4$
Spr (t-value)	1.00	1.00	1.00	1.00	-0.42 (-5.11)		-0.19 (-2.79)	
Govbal (t-value)			0.24 (7.20)		2.45 (2.20)		-4.93 (-6.39)	-1.13 (-2.48)
Lreer (t-value)	0.32 (11.13)			1.35 (10.74)				-0.44 (-4.17)
Lresbnb (t-value)		0.12 (4.92)				-1.63 (-3.64)		
Lnm0 (t-value)			0.07 (8.55)	-0.19 (-7.68)				0.61 (5.91)
Cab (t-value)	0.32				-0.81 (-4.54)	0.34 (2.44)		0.21 (3.33)
DS988 (t-value)	-0.01 (-3.14)	-0.02 (-1.66)	-0.04 (-5.42)	-0.06 (-4.00)				
ustb10_1 (t-value)	0.009 (4.37)	0.009	0.009	0.009				

The estimates of α presented therein are used in the construction of an estimate of $\alpha_{\perp}$, which provides the coefficients of the common trends (see Table 13). The first common trend is normalized to the money variable and is identified by the restriction that it contains no shocks to the spread, while the second trend is identified by the condition that it contains no monetary shocks and is normalized to the spread.

Table 13: Orthogonal Complements for the Bulgarian Model.

	Common Trends		Loadings of the Common Trends			
	$\alpha_{\perp,1}$	$\alpha_{\perp,2}$	$\tilde{\beta}_{\perp,1}$	$\tilde{\beta}_{\perp,2}$	$\tilde{\beta}_{\perp,1}$	$\tilde{\beta}_{\perp,2}$
Spr		1.00	-0.02	0.26		0.26
Govbal		-0.04	-0.39	-1.26	-0.45	-1.26
Lreer	1.39	-0.20	0.22	-0.12	0.21	-0.12
Lresbnb		-0.13	0.12	-2.20		-2.20
Lnm0	1.00		1.49	0.50	1.49	0.50
Cab		-0.63	-0.18	-0.69	-0.21	-0.69

The analysis of the common stochastic trends makes it possible to distinguish between the sources of unexpected permanent changes in the variables studied. In principle, in the CVAR the permanent component of a variable is a combination of deterministic and stochastic trends. The deterministic trend is the predictable permanent component. The stochastic trends, which have zero ex ante expectation, represent the unpredictable, or unexpected, permanent deviations from the deterministic trend. If a stochastic trend τ_t is given a structural interpretation, say as an international trend, then the loading $\tilde{b}_{\tau i}$ of that trend into variable i is a coefficient used to compute the permanent change in that variable attributable to unexpected international events in time t. Namely, this change is $\tilde{b}_{\tau i}\Delta\tau_t$. All loadings are collected in the matrix $\tilde{\beta}_{\perp} = C\alpha_{\perp}\left(\alpha_{\perp}'\alpha_{\perp}\right)^{-1}$, where $\tilde{\beta}_{\perp,ij}$ is the loading of trend j into variable i. From this matrix it can be seen how permanent changes are related across variables. The total unexpected permanent change in variable i due to events in period t is obtained by adding the changes attributed to the various stochastic trends. It should be distinguished from the shock to variable i in period t, or ε_{ti}, which is the instantaneous unexpected change. If the effect of unexpected time t events on variable i is to be plotted against time, the plot will start from ε_{ti} and will have the total unexpected permanent change as its horizontal asymptote. In this sense, the common trends analysis is an analytical impulse response analysis.

An attempt will now be made to give a structural interpretation to the two stochastic trends driving the already estimated CVAR. It turns out that only monetary shocks and shocks to the real exchange rate are cumulated in the first trend (see Table 13):

$$stochastic\, trend_{1t} = \sum_{i=0}^{t}\varepsilon_{lm0,i} + 1.39\sum_{i=0}^{t}\varepsilon_{lreer,i} = \ln m0_t + 1.39 lreer_t - \text{deterministic term.}$$

Since in the period studied the BGN/DEM exchange rate is fixed, the phenomenon of real appreciation of the national currency observed in this period is to a large extent due to the rising level of domestic prices. That is why the *nominal* trend in the price level is, at first sight paradoxically, present in the *real* exchange rate (see Figure 5).

This trend is also present in the level of base money, which is a nominal variable. Therefore, the first stochastic trend can be regarded as a *nominal trend*.

The loadings of the nominal trend (the first column of $\tilde{\beta}_{\perp}$) show, as could be expected, that the latter feeds into base money and the real exchange rate. It seems that *the nominal trend affects only slightly, if at all, the risk variable* ($\tilde{\beta}_{\perp,11} = -0.02$). It is interesting to test whether the effect is exactly zero ($\tilde{\beta}_{\perp,11} = 0$), both because of its own interpretation, and because it would facilitate considerably the interpretation of the cointegration relations. Since shocks to money do not cumulate in the second stochastic trend, $\tilde{\beta}_{\perp,11} = 0$ is equivalent to the statement that the long-run effect of shocks to money on risk is zero. This implies that the corresponding entry of the C matrix (the moving average impact matrix) is zero, or that the cumulated forecast error decomposition impulse response of risk to a shock in money is zero. The latter is plotted for the model with unrestricted α and β together with the 90 percent confidence intervals (see Figure 6), and zero is in the confidence band when the forecasting horizon increases. Thus, $\tilde{\beta}_{\perp,11} = 0$ is accepted.

This is an important finding, and one possible interpretation is that it proves the international credibility of the currency board regime: the price/monetary trend is not perceived as a danger to financial stability, because the latter is guaranteed by the currency board. A complementary explanation is that, at least in theory, the real convergence of the Bulgarian economy to European income levels can be expected, through the Balassa Samuelson effect, to be accompanied by price increases. Thus, if the bonds market interprets price increases as related to growth, they need not be a source of risk. Of course, such an explanation is consistent with $\tilde{\beta}_{\perp,11} = 0$, but is not implied by it. The analysis of other Central and Eastern European countries, which have different exchange rate regimes (e.g. Poland), may turn out to be helpful in resolving the interpretational ambiguity.

An adjusted version of $\tilde{\beta}_{\perp}$, such that $\hat{\beta}'\tilde{\beta}_{\perp adj} = 0$ and $\tilde{\beta}_{\perp adj,11} = 0$, is easy to calculate (see the third column of Table 1.13). The assumption $\tilde{\beta}_{\perp,11} = 0$ implies $\tilde{\beta}_{\perp,41} = 0$, since otherwise the second column of $\hat{\beta}$ will fail to represent a stationary linear combination. Hence, there should be no nominal trend in the level of foreign reserves; these were taken to be measured in USD, and the implication is plausible. It receives confirmation from the impulse response analysis, as zero is in the 90 percent confidence interval for all plotted horizons (the fourth graph in Figure 6).

Another consequence of $\tilde{\beta}_{\perp,11} = 0$ is that the long-run feedback into the real exchange rate and into the current account balance sum up to zero, and the plotted confidence intervals seem not to contradict this conjecture. It means that an unexpected permanent real appreciation of one percent, if attributable to the nominal trend, is accompanied by a permanent decrease in the CAB-to-GDP ratio of one percentage point, which is consistent with economic logic. Nothing can be said about expected appreciation and about unexpected appreciation not due to domestic price dynamics.

Finally, there appears to be negative feedback from the nominal trend to the government budget balance. This feedback is significant, though marginally, if a 90 percent confidence-level test based on the impulse responses[14] is conducted. It needs to be significant, however, for the budget balance cointegration relation to be consistent with the evidence so far: for this relation to be stationary, the nominal trend should be shared by the budget balance or by the spread, and it has already been accepted that it is not shared by the spread. A possible explanation of the negative effect is as follows. In the Bulgarian currency board, foreign reserves cover the sum of base money and a government deposit with the Issue department of the Bulgarian National Bank. The government can thus conduct a kind of quasi monetary policy, since when it spends from the deposit, the level of base money increases, other things being equal. Say there is an increase in the demand for base money, which, however, is transitory. If the government cares to smooth the fluctuations of base money, it may accept a deterioration of its balance for some time, and finance it from the deposit.

With $\tilde{\beta}_{\perp,11} = 0$, the second trend is the only stochastic trend in the risk variable (the trend s_t in the terminology of section 1.4), and can be called the *risk trend.* Its composition shows that *positive shocks (ε_{ti}-s) to the current account, to the level of reserves, to the government budget balance, and to the real exchange rate, cumulate and reduce the long-run level of risk* (negative weights in $\alpha_{\perp}$, positive feedback into the spread). These shocks also cumulate in the latter four variables, which, however,

[14] The upper bounds of the confidence intervals are close to zero, but still negative.

have only two stochastic trends. Hence, they are cointegrated among themselves, meaning that risk can be characterized in terms of fewer variables.

Alternative characterizations are provided by the estimated cointegration vectors (see Table 12):

$$Spr - 0.009ustb10\_1 = -0.32(Cab + Lreer) + 0.01DS988 + I(0) \qquad (1.1)$$

Permanent increase in the current account balance, if not due to unexpected price-triggered real depreciation, is associated with a permanent decrease in the perceived level of risk (see the first column of β and eq. 1.1). Current account improvements caused by unexpected price-related real depreciation are risk neutral. Technically, the "nominal" trend in the current account and in the real exchange rate cancel, and the remaining trend is shared by the risk variable. Substantively, it must be the case that only those improvements/deteriorations in the current account, which result from competitiveness enhancing/weakening changes in the economic structure, as opposed to those resulting from price competitiveness, are perceived as improving/deteriorating the solvency prospects of the country in the long run, and hence, as decreasing/increasing the risk of default.

$$Spr - 0.009ustb10\_1 = -0.11Lresbnb + 0.01DS988 + I(0) \qquad (1.2)$$

Permanent increase in the level of foreign reserves (measured in USD), which is also an indicator of improved solvency perspectives, is associated with a permanent decrease in the risk variable (see eq. 1.2). Moreover, these two variables have the same stochastic trend (up to proportionality), meaning that in the long run reserves are a proxy for security (defined as minus risk).

$$Spr - 0.009ustb10\_1 = -0.24Govbal - 0.07Lnm0 + 0.04DS988 + I(0) \quad (1.3)$$

Permanent decrease of the government budget surplus, if not related to the nominal trend in base money, is associated with a permanent increase in the level of risk (see eq. (1.3). A decrease related to the nominal trend is risk neutral. Since a lasting deterioration of the government balance is a signal that the government may face difficulties in servicing its debt, a reverse relation can indeed be expected between the risk and the balance. However, consistent with the explanation proposed in the discussion of the stochastic trends, deteriorations of the balance which are readily financed from the government deposit with the BNB, and are thus ex ante fully backed with foreign reserves, present an exception to this reverse relation. The exception is even more justified if such deteriorations are indeed a tool of quasi monetary policy, and hence, are not attributable to deep problems in government finance.

$$Spr - 0.009ustb10\_1 = -1.35Lreer + 0.19Lnm0 + 0.06DS988 + I(0) \quad (1.4)$$

Permanent real appreciation, if not due to a price increase (or equivalently, if not related to the nominal trend), is associated with a permanent decrease in the level of risk (see eq. 1.4).

An inspection of $\hat{\alpha}$ shows that risk adjusts to the relation with the current account and to that with the government budget balance, i.e. permanent changes are propagated to risk by these two variables. This probably means that traders in the bonds market follow the dynamics of precisely these two variables, and much less of foreign reserves, for example. As was already mentioned, however, changes propagated by the budget balance are probably of shorter life, and calling them permanent is a compromise.

The fact that reserves adjust to the risk-reserves relation does not necessarily hide any causality. For example, if this relation is interpreted together with the first one, risk may stand as a proxy for the non-nominal trend in the current account, and thus reserves might be adjusting to the current account.[15] The same applies to the fourth relation. Since the purpose of this paper is to analyze the dynamics of risk,

[15] Indeed, the identification of β can be changed, so that the estimated second relation becomes $cab + Lreer - 0.32Lres$ =deterministic+stationary term, while the remaining relations remain identified as before, and the retrictions on α remain unchanged. The estimated adustment coefficients are 0.45 (3.26) and -0.11 (2.63) respectively for reserves and the current acount, and the model is accepted with a test statistic of 20.644, p-value= [0.3568] in $\chi^2\,(19)$.

and not the interrelations between the other variables, no further efforts are made in this direction.

Hence summarizing the results of the econometric analysis on Bulgaria one can conclude that two common stochastic trends were found. The first was nominal and it is risk neutral. The second is the risk trend. It is interesting to relate our findings to the theoretical and conceptual framework of Section 4.2. Although the analysis is not yet completed and one should wait for the results of the analysis of Poland in order to make a final judgement, at this stage it will still be possible to conclude that we are possibly in case 1.1 or case 1.3. In this situation Bulgaria and Poland would share either one common trend or will not have a common trend at all. This is because the second common stochastic trend in the Bulgarian data was found to be nominal and it is highly unlikely that the two countries will have a nominal common stochastic trend. In further research the Bulgarian and Polish common stochastic trends will be analysed in order to extract a possible international risk factor.

6. The Polish Case

As a first step of the analysis the order of integration of the three time series will be checked by performing unit root tests.

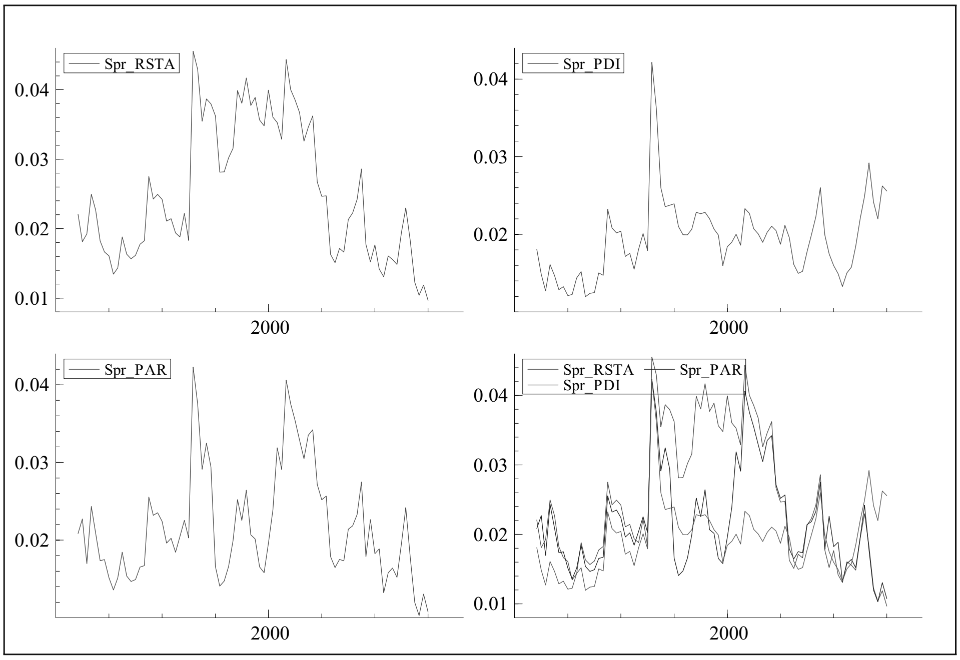

Figure 3: Polish Yield Differentials over an US Treasury Bond.

The applied ADF test[16] checked the hypothesis of existence of unit root against the alternative of stationarity of the DGP. The results (reported in Table 17) indicate quite surprisingly that two out of the three series were stationary and hence that there is no unexpected risk associated with them in the long run.

Therefore the further analysis proceeds in a somewhat different form than expected. A cointegration analysis is performed between the second Bulgarian common stochastic trend, which was found to represent the permanent risk trend, associated with Bulgarian Brady bonds and the non-stationary Polish yield differential.

A VAR with two lags, constant, and a step dummy is estimated. Tha data are plotted on fig. 4. The lag length is chosen again to be the shortest one that makes the residuals serially uncorrelated. In the cointegration analysis the constant and the dummy are left unrestricted.

Table 14 summarizes some misspecification tests which show that the model is a statistically good description of the data, and therefore, a valid inferential framework. The results of the cointegration tests are reported in table 15. Cointegration is cleraly rejected. Hence the Bulgarian risk factor cannot be explained by international risk.

[16] The choice of the appropriate lag lenght was made under the consideration that it should ensure serially uncorrelated residuals, taken into the account the results delivered by the information criteria. Serial autocorrelation was tested by the Portmanteau test.

Since a linear time trend in the levels of the time series can be observed for the series Spr PDI a linear trend and a constant (to account for no zero mean) were included.

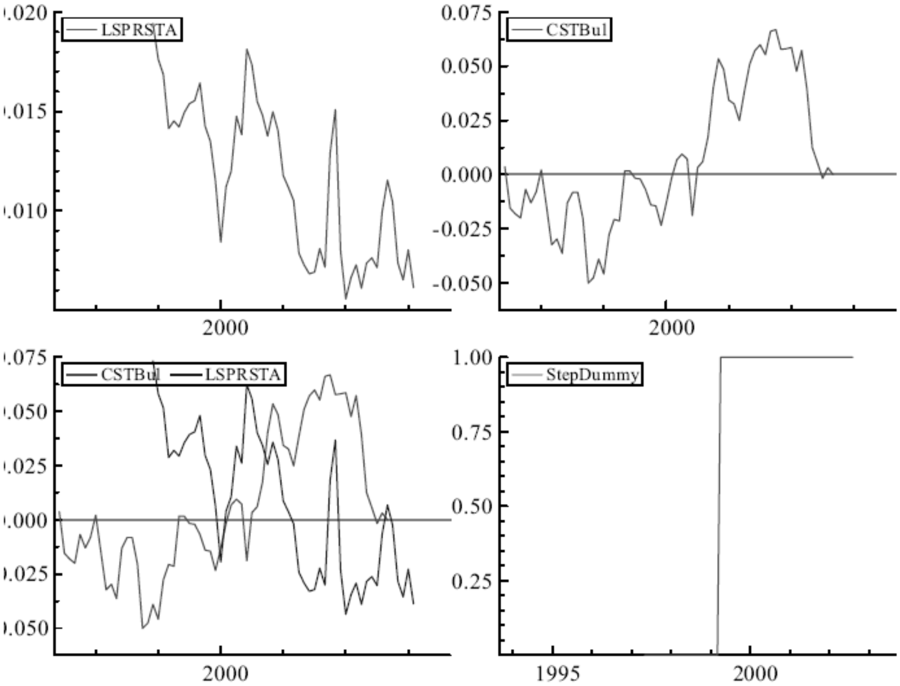

Figure 4: The Bulgarian Risk Trend and the Polish Bond Spread

Table 14: Misspecification Tests for a Trivariate System of Bulgarian Brady Bonds.

	AR 1-3 test	Jarque-Bera	ARCH 1-5
CSTBul	0.71 [0.55] in $F(3,34)$	0.51 [0.77] in $\chi^2(2)$	1.04 [0.39] in $F(3,31)$
LSPRSTA	1.07 [0.373] in $F(3,34)$	7.86 [0.02]* in $\chi^2(2)$	0.93 [0.43] in $F(3,31)$
System Test	0.66 [0.77] in $F(12,60)$	6.96 [0.14] in $\chi^2(4)$	

T a b l e 15: Trace Test Statistics for the Bulgarian Model.

H_0: r $\leq$	Trace Test	P-value
0	13.118	[0.111]
1	3.3151	[0.069]

7. Conclusions

The present paper is one of the first to use time series data and apply cointegration analysis to study default risk and pricing of emerging market sovereign bonds. From the analysis of the Bulgarian case the following conclusions can be drawn: the results confirm the importance of national factors when studying default risk. Moreover, it was found that the Bulgarian currency board is a credible regime and the price/monetary developments are not perceived as a danger to the country's financial stability. The results found, represent important findings. They suggest that market participants do differentiate between Eastern European countries: whereas Poland is a more advanced countries in its transformation process, Bulgaria is still lagging behind. Our findings support the view that it is national risk factors that play an important role in explaining default risk.

Appendix A: Tables and Figures

Table 16: ADF Tests for the Three Bulgarian Yield Spreads.

Variable	**Determ. term**	**num. of lags**	**Portmant. test** $\chi^2(12)$,[p-val]	**Test stat.**	**Critical values*** 10%	5%	1%
Bul Disc	constant, trend	0	2.55 [0.99]	-3.27	-3.13	-3.41	-3.96
ΔBul Disc	constant	0	6.47 [0.89]	-9.71	-2.57	-2.86	-3.43
Bul disc	constant	0	3.43 [0.99]	-2.91	-2.57	-2.86	-3.43
ΔBul Disc		0	6.46 [0.89]	-9.78	-1.62	-1.94	-2.56
Bul IAB	constant, trend		2.23 [0.99]	-3.16	-3.13	-3.41	-3.96
ΔBul IAB	constant		6.39 [0.90]	-9.74	-2.57	-2.86	-3.43
Bul IAB	constant	0	3.43 [0.99]	-2.23	-2.57	-2.86	-3.43
ΔBul IAB		0	6.34 [0.89]	-9.78	-1.62	-1.94	-2.56
Bul Flirb	constant, trend	0	3.46 [0.99]	-2.47	-3.13	-3.41	-3.96
ΔBul Flirb	constant	0	6.10 [0.91]	-6.42	-2.57	-2.86	-3.43
Bul Flirb	constant	0	3.54 [0.99]	-2.52	-2.57	-2.86	-3.43
ΔBul Flirb		0	6.10 [0.91]	-6.51	-1.62	-1.94	-2.56

Table 17: ADF Tests for the Three Polish Bond Spreads.

Variable	Determ. term	num. of lags	Portmant. test $\chi^2(12)$,[p-val]	Test stat.	Critical values* 10%	5%	1%
POL_PAR	constant	0	9.28 [0.67]	-3.02	-2.57	-2.86	-3.43
POL_PDI	constant	0	7.01 [0.85]	-3.55	-2.57	-2.86	-3.43
POL_PDI	constant, trend	0	8.07 [0.78]	-3.77	-3.13	-3.41	-3.96
POL_RSTA	constant	0	8.88 [0.71]	-2.10	-2.57	-2.86	-3.43
ΔPOL_RSTA		0	11.57 [0.48]	-10.29	-1.62	-1.94	-2.56

*Critical values from MacKinnon (1991)

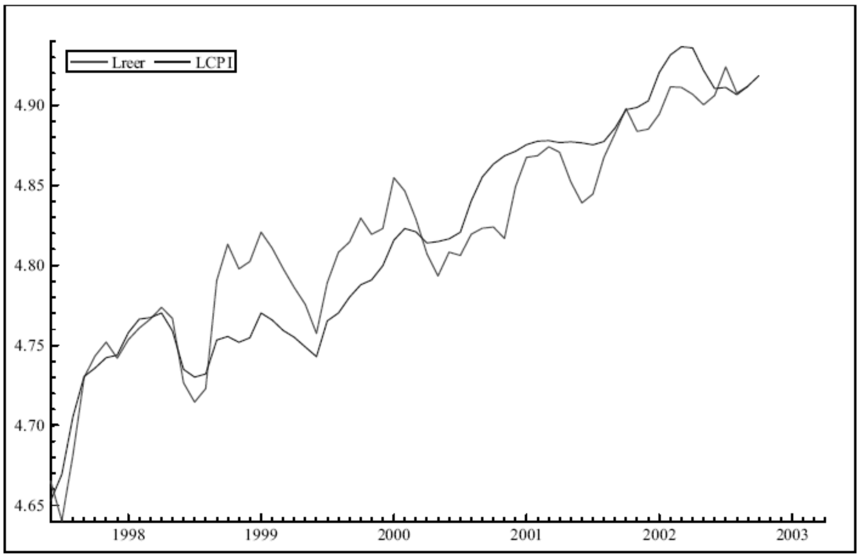

Figure 5: Real Effective Exchange Rate and Consumer Prices for Bulgaria

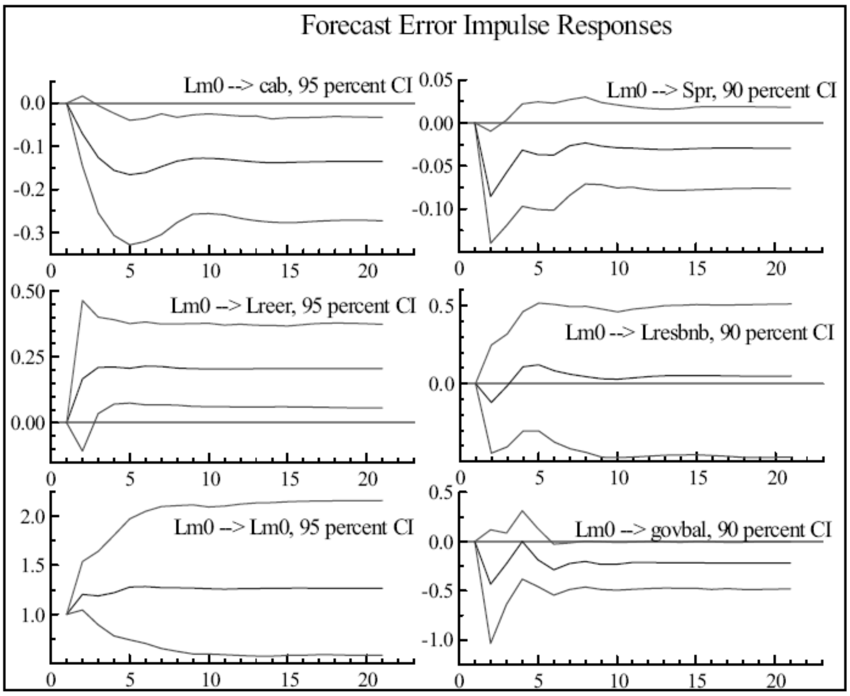

Figure 6: Forecast Error Impulse Responses

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