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Marina Tkalec, Miroslav Verbič and Maruška Vizek

Long-run and Short-run Determinants of Original Sinners' Sovereign Spreads

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Marina Tkalec EIZG

Abstract

This paper builds an empirical model of sovereign spreads and its determinants, relying on recent theories on imperfect capital markets and balance sheet effects. We investigate nine European emerging economies that suffer from the "original sin", over the period 2001-2011, using dynamic panel error correction models proposed by Pesaran et al. (1999). This methodology improves estimation efficiency and model performance, but it also allows differentiation between long-run and short-run spread determinants. We find that in the long-run, sovereign spreads increase in response to a higher share of external debt in GDP, while they move in the opposite direction when the shares of current account and international reserves in GDP rise. In the short-run, sovereign spreads deviate from the long-run equilibrium, with half of the adjustment taking place in eight months. Our results suggest that in the short-run, higher external debt service caused by exchange rate depreciation, i.e. balance sheet effect, and market volatility tend to raise spreads, while higher tax revenues tend to decrease them. Moreover, we prove that the rise in sovereign spread is not due to external debt accumulation itself, but due to pure balance sheet effects.

Keywords: balance sheet effect, emerging Europe, euroization, original sin, sovereign spreads.
JEL Codes: F31, F34, G15, H63.

1. Introduction

Economic literature provides plenty of explanations for determinants of sovereign spreads – differences between interest rates that governments pay on their debts, and the interest rates that for example the United States or Germany pay on their debt. Sovereign spreads are a proxy for country risk premium, measure of the risk associated with a country's default on debt. This premium, or spread, is formed in order to compensate creditors for the risks of holding a risky asset until maturity. The whole idea of exploring spreads comes from the fact that sovereign spreads are higher for some countries than for the other. Emerging market economies have higher spreads than developed ones, arousing curiosity around spread determinants and channels of impact. Theory suggests that spreads depend on fundamental, macroeconomic conditions because in the long-run, spreads are affected by the size of the debt itself, total wealth, the current account and international reserves (Edwards, 1984). However, it is common that this long-run relationship breaks in the short-run, especially in turbulent times. For example, after Lehman Brothers collapsed in 2008, spreads on emerging market sovereign bonds raised swiftly, regardless the fact that their macroeconomic indicators stayed unchanged. This sort of behaviour suggests that something different is happening in the short-run, and that there are maybe some other determinants affecting spreads besides the usual suspects. This paper tries to detect why sovereign spreads deviate from the long-run equilibrium level, using market sentiment, monetary and fiscal policy as possible spread dynamics drivers.

Monetary policy and exchange rates are of special interest here, because there are opposing views on the impact they have on sovereign spreads. Although conventional open economy models suggest that real exchange rate depreciation is expansionary, recent theories on imperfect capital markets and balance sheet effects prove just the opposite (Aghion et al., 2004; Céspedes et al., 2000). For example, if a country is highly indebted in foreign currency, then debt servicing increases together with real exchange rate depreciation. Thereby causing deterioration in country's balance sheets, a fall in aggregate demand, and consequently, in economic activity too (Berganza et al., 2004). Contradicting theories can not itself decide on the importance and validity of these effects, so additional empirical work is needed in order to decide on the relevance of each theory. We therefore build a model that incorporates the newest theoretical and empirical findings, and empirically test the existence of a positive relationship between sovereign spreads and exchange rate depreciation, presented by the balance sheet effect. Among other things, we test if the increase in debt service caused by an unexpected real depreciation significantly raises sovereign spreads in the short-run. In this study, we find evidence of such positive balance sheet effects on sovereign spreads for European emerging countries, thus corroborating Berganza et al. (2004).

This paper uses multiple strands of literature to build a new empirical model of sovereign spread determinants. We combine three different strands of existing research to explain sovereign spread dynamics in countries that suffer from "original sin" - impossibility to issue debt in local currency (Eichengreen at el., 2003). We use the small open economy model by Céspedes et al. (2000) and Gertler et al. (2007) as our basis, to which we add two supplementary concepts. Firstly, we borrow the collateral value concept from Kiyotaki and Moore (1997), and then add the balance sheet effect empirical findings from Berganza et al. (2004). Onwards, we construct the model so that it differs between long-run and short-run, thereby allowing both differences between countries that occur in the short-run, and theoretical universalities that comply in the long-run. This is obtained by the panel version of the error-correction model, the pooled mean group (PMG) estimator developed by Pesaran et al. (1999). PMG provides a dynamic framework that allows a separation between the short-run and the long-run, enabling both short-run dynamics and equilibrium adjustment.

The main contribution of this paper is that it incorporates the balance sheet effect as a short-run sovereign spread determinant, just as observed in empirical data. Unlike previous studies, that either ignore differences between the short-run and the long-run, or the existence of balance sheet effects, our research allows for such an effect in the short-run. This is possible only because we construct a dynamic model that deviates from equilibrium in the short-run, and then gradually adjusts in the long-run. Moreover, we also assume that not all countries react the same to changes in fundamentals, and in that respect, we allow short-run heterogeneity between countries. Additionally, our data set includes the latest financial crisis data, thus taking into account sovereign spread volatility observed in the last few years. And finally, we add three countries, Croatia, Serbia, and Turkey, which were highly underrepresented in previous research.

The rest of the paper is organized as follows. Section 2 discusses previous work and sets the theoretical and empirical framework for our model. Section 3 describes the data set, while section 4 presents the empirical model and the estimation technique. Results are given in section 5, while the last section discusses possible implications, and concludes the paper.

2. Theoretical framework and empirical work

2.1. Theoretical framework

Small open economy models are a starting point for investigating emerging market borrowing, and related country risk premiums. The simplest framework is given in early works by Edwards (1984, 1986), in which the country risk premium is related to the probability of default on external liabilities. This probability is defined as a function of a number of macroeconomic and external variables, and has become the basis for studying government bond spreads. However, both Edwards (1984) and later Kim (1998) use a very constraining assumption of perfectly competitive financial markets and risk neutral lenders, assumptions becoming more and more relaxed nowadays.

A different approach is taken by Cantor and Packer (1996) who simply replace macroeconomic fundamentals with credit ratings, arguing that the inclusion of both would lead to multicollinearity. Another study by Kamin and Kleist (1999) ignores specific macroeconomic, as well as solvency and liquidity indicators, and uses solely credit ratings to explain sovereign yields. On the other hand, Eichengreen and Mody (1998a, 1998b) and Dell'Ariccia et al. (2002) claim that credit ratings are defined more broadly than macroeconomic variables, and in line with that there is no multicollinearity threats when using both. We follow both groups of authors, and build models with macroeconomic variables only, and models that include both credit ratings and macroeconomic fundamentals. Our results (see footnote 6) show that sovereign spreads do not have a statistically significant relationship with credit ratings.

As mentioned, sovereign spreads are a function of the probability of default, and related to that, a function of the probability of loss in case of default. The probability of default is represented by external debt sustainability, which is in turn measured by indicators of liquidity and solvency. The bottom line is that in reduced-form models, one uses macroeconomic variables to reflect liquidity and solvency, and accordingly, probability of default. This is the starting point made in Edwards (1984)[1], and is represented by a linear equation:

$$spread_t = \alpha + \sum_{j=1}^{J} \beta_{jt} x_{jt} + \varepsilon_t \quad (1)$$

where $spread_t$ is the sovereign spread at time t, α is an intercept term, J is the number of macroeconomic variables with x_j being a set of these variables, parameters β_j are slope coefficients, and ε_t is the error term. The specification in equation (1) develops with the inclusion of explanatory variables – different spread determinants. Since we explore emerging markets and their sovereign spreads, we will use a small open economy that is externally indebted. Economic theory asserts that small open economies borrow from abroad when their resources do not suffice their consumption potential, and that they repay the debts when they have extra resources. The main focus of foreign investors, in this setting, is whether a country will be able to repay its debts. First, does it have enough foreign exchange to service its obligations, and second, is the government able to collect enough resources to purchase foreign exchange to repay them. Ferrucci (2003) proposes a dynamic programming setting in which he defines a welfare function U_0, that depends on future discounted consumption. He assumes that the country maximizes its consumption C_t, or utility, by using its available resources and by issuing debt. The model is presented below:

$$\begin{aligned} &Max && U_0 = \sum_{t=0}^{\infty} \beta^t u(C_t) \\ &s.t. && G_t + rD_t \le T_t + D_{t+1} - D_t \\ & && Y_t = C_t + G_t \\ & && T_t = f(Y_t) \\ & && Y_t = (1+g)Y_{t-1} \end{aligned} \quad (2)$$

with β the discount factor. The maximizing problem is subject to two constraints – the government budget constraint and the accounting identity. The first constraint binds public spending G_t to total revenues less interest payments on existing external debt rD_t. Total revenues are defined as a sum of domestic tax revenues T_t, and newly issued external debt $D_{t+1} - D_t$. The accounting identity implies that total domestic output is composed of private and government consumption. Inserting the second constraint into the first, and rearranging we get:

$$D_{t+1} - D_t \ge Y_t - C_t - T_t + rD_t \quad (3).$$

[1] Work by Edwards (1984) is founded on some previous studies, such as Feder and Just (1977), Eaton and Gersovitz (1980), and Sachs (1981).

The left-hand side of expression (3) is the current account, or the external constraint, and it implies that newly issued external debt must be at least as large as the sum of private saving and interest payments on existing external debt. To close the model, tax revenues and domestic output need to be defined, given in last two expressions. Tax revenues are a function of output, which is exogenously defined by its lagged value and growth rate.
This dynamic model suggests that borrowers comply with the constraints in each time period, not only in the long-run. This leads us to the conclusion that governments should have enough liquid assets to repay their foreign investors in each period, and be solvent in the long-run. From specifications (2) and (3) it is evident that fiscal balance, external debt interest repayments, tax revenues, and current account are important factors in determining government solvency and liquidity. Although they are not directly included here, official foreign reserves are insurance that liabilities will be serviced even in cases when countries do not provide necessary funds to repay foreign investors.
Net present values of government and external constraints provide us with the external debt and fiscal sustainability conditions:

$$\left(1+r\right)D_t \le \sum_{i=0}^{\infty} \frac{PS_{t+i}}{\left(1+r\right)^i} \qquad (4)$$

$$\left(1+r\right)D_t \le \sum_{i=0}^{\infty} \frac{C_{t+i}+T_{t+i}-Y_{t+i}}{\left(1+r\right)^i} \qquad (5)$$

where $\mathrm{PS}_t = T_t - G_t$ is the primary fiscal balance, and $C_{t+i}+T_{t+i}-Y_{t+i}$ is future private saving. Equations (4) and (5) present sustainability conditions of fiscal policy and external debt, more specifically, suggest that external debt today, should not exceed the net present values of future primary fiscal surpluses or future private saving (both discounted by r, the capital cost). These two conditions are central for external debt sustainability, and consequently for country risk premium assessment. Different solvency indicators can serve as reliable determinants of external debt sustainability, such as tax revenues, level of public debt, current account, external debt level, official international reserves, international trade, etc. Trade, especially export, is important because it provides foreign currency necessary to repay the external debt.
The simple setting presented above ignores other significant factors, such as terms of trade, inflation, and exchange rates. Roughly speaking, terms of trade reflect how much foreign exchange is coming into the country from exports, relative to foreign exchange coming out of the country to pay for imports. High terms of trade are a signal for investors that their loans will be repaid. Min (1998) connects fiscal policy with inflation, and claims that high inflation rates reflect fiscal imbalance. In line with that, inflation affects the fiscal sustainability condition, and leads to higher sovereign spreads. On the other hand, McDonald (1982) links inflation to the balance of payments and finds that higher inflation leads to higher probabilities of balance of payments and default crises.
Another strand of literature builds on exchange rates and balance sheet effects. Open economies that borrow in foreign currency, rather than in local, suffer from the original sin (Eichengreen et al., 2003). High shares of foreign currency debt in total debt, or credit euroization[2], make a country vulnerable to exchange rate changes, if exports are not high enough to cover external liabilities. The reason is that exchange rate depreciation increases foreign currency debt in local currency terms. If a country has most of its assets in local currency, and liabilities in foreign currency, it suffers from a currency mismatch (Luca and Petrova, 2008), and is a potential victim of negative balance sheet

[2] Throughout the text, we will use the term euroization instead of dollarization. Although dollarization is a universal concept, not necessarily referring to dollars, we are exploring countries that are traditionally more connected to the euro, and consequently borrow in euro, or link their debt to the euro.

effects. Therefore, if a country's exchange rate is perceived to be overvalued and future depreciation is expected, its risk premium will increase accordingly.

A theoretical framework for including balance sheet effects into the small open economy model is motivated by Céspedes et al. (2000) and Gertler et al. (2007). In short, small net worth of a country (as defined below) implies a greater demand for external resources. Due to asymmetric information between domestic issuers and foreign creditors, this foreign borrowing increases agency costs. Exchange rate depreciations affect these agency costs in an adverse manner, which manifests itself in an increasing country risk premium. Another strand of literature (Kiyotaki and Moore, 1997) is based on the collateral value, as it argues that the cost of borrowing falls with the value of the collateral. In our case, the value of the collateral is comparable to real net worth of a country. Both of these views are combined in Berganza et al. (2004), a study that focuses on detecting whether there is an inverse relationship between government bond spreads and real net worth ω_t, presented by:

$$1 + spread_t = \Psi(\omega_t),\ \Psi' < 0 \quad (6)$$

where real net worth is assumed to be composed of tradable and nontradable goods. Under high credit euroization, and exchange rate depreciation, ω_t decreases in foreign exchange terms, leading to a rise in $spread_t$. This is shown by equation (7):

$$\omega_t = X_t - D_t R_t . \quad (7)$$

where X_t stands for a set of net worth determinants, and R_t for the real exchange rate (as before, D_t is external debt). Berganza et al. (2004) then take a linear approximation around the mean value of net worth, denoted by $\bar{\omega}$, and obtain the following:

$$\begin{aligned} 1 + spread_t &\approx \Psi(\bar{\omega}) + \Psi'(\omega_t - \bar{\omega}) \\ &\equiv \alpha - \beta\omega_t \\ &= \alpha - \beta X_t + \beta D_t R_t \end{aligned} \quad (8)$$

where $\alpha = \Psi(\bar{\omega}) - \Psi'(\bar{\omega})$ is the constant term, and $\beta = -\Psi'$ is the negative derivative of Ψ at ω_t. Inserting equation (7) into the second row of specification (8) leads us to the third row. The most important part is $\beta D_t R_t$ since it represents the balance sheet effect. An increase in the real exchange rate (depreciation), leads to a rise in the risk premium when the country has a high level of debt in foreign currency, and naturally, when β is significantly positive. Theoretically, there is no reason to believe that β is positive; it can be negative or zero. That is why it is necessary to empirically test the sign and the size of β. For that, however, we need further elaboration. By subtracting the expectation of the last expression in (8) from that expression without the expectation, conditional on information at $t-1$, assuming that D_t is predetermined, and after rearranging, we get:

$$spread_t = E_{t-1}spread_t + \beta D_t (R_t - E_{t-1}R_t) + \varepsilon_t \quad (9)$$

where $\varepsilon_t = \beta(X_t - E_{t-1}X_t)$ is the stochastic term of the explanatory variables. If we assume that this stochastic term is unobservable, we can estimate equation (9) in case ε_t is not correlated with

$D_t\left(R_t - E_{t-1}R_t\right)$. Having in mind that debt is predetermined, we only need to assume that ε_t is not correlated with $R_t - E_{t-1}R_t$. We reformulate equation (9) by replacing $E_{t-1}spread_t$ with $\gamma spread_{t-1}$ [3] and $E_{t-1}R_t$ with R_{t-1}. Although we loose some information, Berganza et al. (2004) observe that these losses are negligible, since spreads and exchange rates often behave like random walks. We also simplify a little bit more, by replacing $D_t\left(R_t - R_{t-1}\right)$ with S_t, and obtain:

$$spread_t = \beta S_t + \gamma spread_{t-1} + \varepsilon_t \ . \qquad (10)$$

S_t is defined as a product of real exchange rate depreciation and the value of external debt, or as the value of foreign currency denominated debt in local currency terms. Together with the balance sheet variable and its coefficient β, Berganza et al. (2004) use different net worth drivers, contained in X_t or in ε_t. This brings us to equation (1) and the model presented by Ferrucci (2003), who also uses different variables to describe the spread. Resulting from specification (10), we can now insert the balance sheet effect S_t into equation (1) and get a more complete risk premium model:

$$spread_t = \alpha + \beta S_t + \gamma spread_{t-1} + \sum_{j=1}^{J} \delta_{jt} x_{jt} + \varepsilon_t \qquad (11).$$

2.2. Empirical work

As surveyed here, economic theory deals with long-run spread determinants, regardless of the fact that spreads are rather volatile in the short-run. Although spreads typically follow the path predicted by theory, depending on external debt, international reserves, fiscal and current account balances, they also deviate in the short-run caused by some specific factors. Studies that use short-run variables for explaining spreads, and the ones that differentiate between long-run and short-run spread determinants are rather scarce and limited to recent literature. We discuss that specific strand of literature in more detail.

The usual methodological frameworks for analysing spreads are single-country and panel data studies. Both of these have pros and cons; single-country studies take care of country-specific characteristics, but suffer from difficulties with statistical inference that appear due to short data sets, a typical difficulty when dealing with emerging market countries. Panel data studies have larger data sets that lead to improvements in statistical inference, but they assume homogeneity of the slope coefficients, neglecting country-specific characteristics.

Single-country studies use time series methods for analyzing spreads. Usually debt, fiscal and current account balances are included as regressors, together with different control variables specific for the country explored. The most important time-series study for European transition economies is by Ebner (2009) who concludes that market variables, and not macroeconomic, are more important for explaining spreads in these countries. Panel data studies also combine theoretical spread drivers with some control variables, and come to similar conclusions. Dumičić and Ridzak (2011) use a panel of eight transition countries and differ between spread drivers before and after the crisis. They conclude that financial volatility factors are important spread drivers, and that countries with higher current account deficits had larger spread increases in the period after the crisis. On a similar note, Hagen et al. (2011) explore EU countries and find that the fiscal deficit coefficients are positive, implying higher spreads for less fiscal discipline, and that those coefficients are larger in the period after the crisis. Berganza et al. (2004) run a panel of 27 emerging countries, and among other results,

[3] In the original paper, Berganza et al. (2004), replace $E_{t-1}spread_t$ with the lagged GDP value because their estimation procedure does not allow for dynamics in the system. Ours does, so we leave the lagged dependent value in the equation.

find evidence that higher external debt service costs caused by exchange rate depreciations lead to higher sovereign spreads. They however, do not allow for any heterogeneity between countries nor do they differ between short-run and long-run. Malone (2009) follows the work by Berganza et al. (2004), explained here in detail, but extends their work on the endogeneity problem caused by the real exchange rate variable. Malone (2009) can not reject the hypothesis of no endogeneity and no simultaneity bias, thus corroborating the Berganza et al. (2004) model and findings.

There are only a few studies that allow for heterogeneity between countries in the short-run. Bellas et al. (2010) use the pooled mean group estimator for 14 emerging countries[4] and find that only financial stress indices and market volatility affect spreads in the short-run.[5] Alexopoulou et al. (2010) use a dynamic panel error-correction framework to model sovereign spreads in emerging Europe. They find that countries with low fiscal discipline suffer from more volatile spreads. Moreover, they argue that exchange rates have a positive effect on spreads in Hungary, Poland, and Slovakia. Our work differentiates from Alexopoulou et al. (2010) in three main points. Firstly, we construct a variable that takes into account the balance sheet effect that directly measures the relationship between the exchange rate and credit euroization. Additionally, our data spans through the years of the financial crisis, and we explore three countries not previously covered, Croatia, Serbia, and Turkey. Finally, Ferrucci (2003), using the same methodology, concludes that macroeconomic fundamentals and external liquidity conditions are most important spread drivers. Ferrucci (2003) however does not include a balance sheet effect variable, nor does he consider the influence of monetary policy.

To sum up, empirical research finds that macroeconomic variables and external debt have a significant influence on government spreads. Important part of literature suggests there are also balance sheet and market behaviour variables that are important spread drivers. Moreover, panel data studies seem to be more successful in acquiring more efficient estimates, but suffer from generality when making country-by-country conclusions. Additionally, it is observed that spreads have become more volatile, and that traditional determinants are not appropriate for explaining deviations in the short-run. This problem is usually solved by introducing new short-run variables, and by using a methodology that separates short-run from long-run variables and allows for differences between countries.

3. Data

As presented in equation 1, the dependent variable is sovereign spread, our chosen measure for country risk premium. A typical and widely used proxy for sovereign spread is the JP Morgan Euro Emerging Markets Bonds Indices (EMBI) Global. Euro EMBI Global is a spread by construction, as it is equal to the difference between returns on foreign currency bonds and corresponding US Treasury bonds. It has become a standard to use such secondary market spreads in order to represent country risk premiums. For example, Dell'Ariccia et al. (2002) use EMBI Global spreads for exploring international borrowing after the Russian crisis, and a number of further work relies on the same source as well (Ferrucci, 2003; Berganza et al., 2004; Bellas et al., 2010). Their advantage is that they overcome the bias that might arise out of a choice of the basket of bonds that are representative for country risk premium. However, EMBI series are rather short, as the longest data sets are available only since 1997. We use data in quarterly frequency because some of our regressors are available on quarterly basis only. We limit our sample to nine European countries (Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania, Serbia, Slovak Republic, and Turkey), as EMBI spreads are not available for other countries. Data for the biggest number of countries are available from the first quarter of 2001 until the fourth quarter of 2011, which sets our panel database to 396 observations altogether. However, for some countries and periods, observations are missing, so we had to work

[4] Their sample incorporates only three European countries: Bulgaria, Poland and Turkey.

[5] A more detailed study on the relationship of spreads and financial variables can be found in Mody (2009).

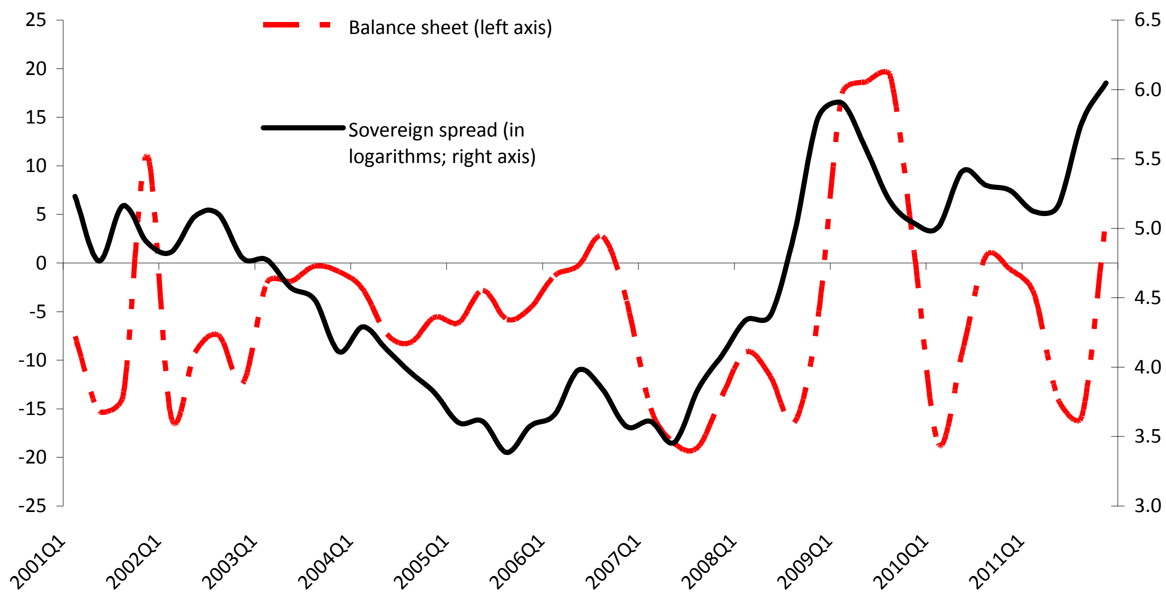
with an unbalanced panel. Table A1 of the Appendix presents descriptions, sources and expected signs for all the variables we use in the empirical examination.

Consistent with existing literature, we use several long-run variables found to be important determinants of spreads. These are external debt, current account, and international reserves. All are defined as percentages of GDP, and they were tested for stationarity using different panel unit root tests.

In accordance with equation 11, and taking into consideration that we differentiate between the long-run and the short-run, we include the balance sheet term S_t as a short-run spread determinant. S_t is an interaction term composed of external debt and real exchange rate changes, intended to account for the rise in the service of external debt in the aftermath of an exchange rate depreciation. In line with our model and Figure 1, we would expect a positive balance sheet effect. Put differently, we would expect that besides the fact that the real exchange rate depreciation increases external debt service, it also increases the country risk premium. Figure 1 shows average sovereign spreads together with the constructed balance sheet variables for the 2001-2011 period. We can see that the balance sheet variable follows the spread turning points, and that they move more-or-less in the same direction. Especially interesting is the strong positive co-movement observed in 2008 that coincides with the beginning of the financial crisis.

Following the literature, we include another short-run variable, one that measures market behaviour, more specifically, market volatility. We define that variable as a logarithm of CBOE (Chicago Board Options Exchange) VIX (Volatility IndeX). Finally, we consider the impact of short-run fiscal policy measures, such as tax revenues dynamics.

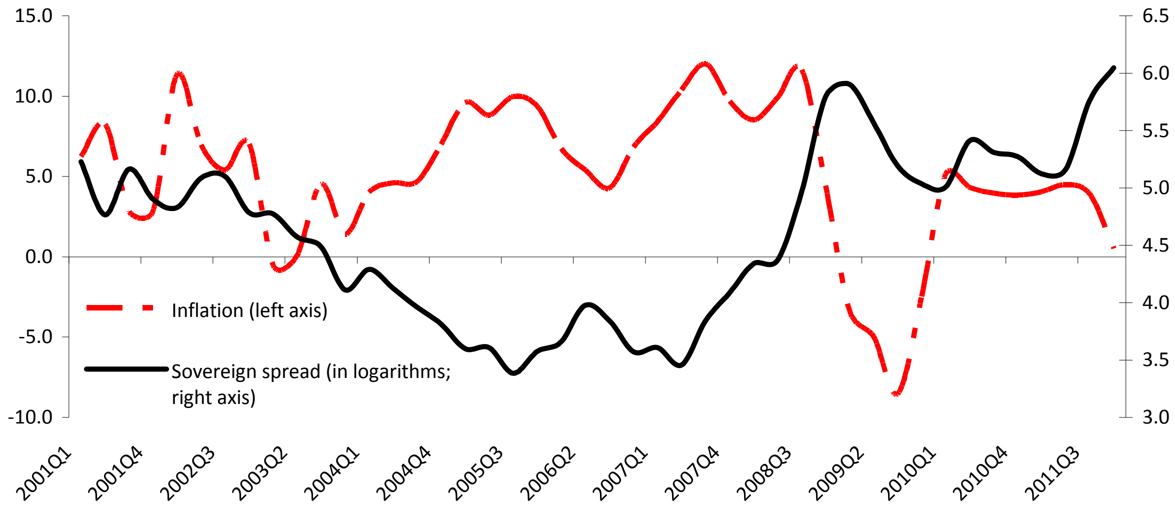
Figure 1. Balance sheet and sovereign spread movements in the period 2001-2011



Note: Variables presented here are averages of the nine countries.
Source: JP Morgan, national central banks, and own calculation.

We also include different control variables, such as exports. Besides the positive balance sheet effect, one would expect that the exchange rate depreciation spurs exports. Therefore, we add exports to the empirical model in order to control for such an effect, and to avoid possible omitted variable issues. For the same reason we add another short-run control variable, external debt, and an instrument variable for the balance sheet effect that is constructed as a product of external debt and inflation. The latter is necessary as it is sometimes argued that inflation actually causes spread rallies, and not the exchange rate (Min, 1998). However, already from Figure 2 we can expect that this is not the case here. Inflation and sovereign spreads sometimes move in tandem, but otherwise they seem to be diverging. Finally, all variables, except international reserves and the volatility index, are seasonally adjusted using ARIMA X12. The reason we did not apply seasonal adjustment to these two variables, is that they do not show any signs of seasonal activity. International reserves have a smooth increasing trend in all the countries we explore, while the volatility index resembles a random walk.

Figure 2. Inflation and sovereign spread movements in the period 2001-2011



Note: Variables presented here are averages of the nine countries.
Source: JP Morgan, national central banks, and own calculation.

Table 1 presents descriptive statistics for the variables used in our baseline model, while Table 2 shows correlation coefficients between spreads and its determinants, country-by-country. Although these are only correlations, they still suggest that the regressors are highly correlated with sovereign spreads in all countries we explore. As expected, the long-run variables show high correlation, especially external debt. However, short-run drivers are also significantly correlated with spreads. The variable of our interest, balance sheet, is significantly correlated with spreads in five out of nine countries. We expect however, that pooling data will give us more insight and preferably more consistent results.

Table 1. Descriptive statistics

	Mean	Median	Min	Max	St dev
Spread	4.6	4.8	3.4	6.0	0.8
External debt	265.6	230.7	156.8	447.0	100.2
Current account	-6.6	-5.8	-15.1	-2.1	3.4
International reserves	85.8	88.2	47.3	131.6	23.5
Balance sheet	-5.2	-5.7	-18.9	19.3	9.4
Volatility	3.0	3.1	2.4	4.1	0.4
Tax revenues	22.3	22.5	16.9	27.4	3.4

Note: Values presented here are for averages of the nine countries.

Table 2. Correlation coefficients

Spread	Bulgaria	Croatia	Czech Republic	Hungary	Poland	Romania	Serbia	Slovak Republic	Turkey
External debt	0.429*** [0.007]	0.875*** [0.000]	0.860*** [0.000]	0.849*** [0.000]	0.597*** [0.000]	0.397** [0.010]	0.587*** [0.001]	-0.122 [0.506]	0.354* [0.055]
Current account	0.387** [0.016]	0.450** [0.013]	-0.216 [0.390]	0.762*** [0.000]	-0.138 [0.372]	0.404*** [0.009]	0.141 [0.475]	0.068 [0.712]	0.272 [0.146]
International reserves	0.195 [0.240]	0.730*** [0.000]	-0.262 [0.293]	0.896*** [0.000]	0.625*** [0.000]	0.141 [0.378]	0.344* [0.074]	-0.808*** [0.000]	-0.032 [0.869]
Balance sheet	0.111 [0.507]	0.352* [0.057]	-0.549** [0.018]	0.273* [0.076]	0.416*** [0.005]	0.606*** [0.000]	0.234 [0.231]	0.248 [0.171]	0.225 [0.232]
Volatility	0.867*** [0.000]	0.869*** [0.000]	0.933*** [0.000]	0.627*** [0.000]	0.820*** [0.000]	0.863*** [0.000]	0.614*** [0.000]	0.770*** [0.000]	0.681*** [0.000]
Tax revenues	0.023 [0.889]	0.532*** [0.002]	0.687*** [0.002]	0.513*** [0.000]	0.170 [0.269]	-0.007 [0.965]	0.410** [0.030]	-0.665*** [0.000]	-0.120 [0.528]

Note: Variables are in levels; p-values are in brackets; ***, **, and * denote significance at 1, 5, and 10 percent confidence level, respectively.

4. Estimation

4.1. Methodology

4.1.1. Pooled Mean Group Estimator

Existing empirical literature offers either panel data or country-by-country estimation results, ignoring either differences between countries or universalities that arise from general theoretical concepts. This study tries to combine both, because we use a PMG estimator that enables us to explore data in a panel setting, still allowing for short-run country-specific deviations. PMG is a panel version of the error-correction model, and provides an opportunity to obtain more efficient estimation results, while preserving some group-specific heterogeneity. It is appropriate to use PMG whenever there is reason to believe that countries differ on the matter in the short-run, but comply in the long-run.

A dynamic panel can be estimated using different procedures, with each one offering both advantages and limitations in comparison to alternative methods. For example, if we pool the time-series data for each group and allow only the intercepts to differ across groups, we are using a dynamic fixed effects framework. However, if the slope coefficients are actually heterogeneous, then fixed effects estimation results would be inconsistent and misleading. An alternative method is one proposed by Pesaran and Smith (1995), who suggest using the mean group (MG) estimator in which intercepts, slope coefficients and error variances are allowed to differ across groups. In this setting, the panel coefficients are given as simple averages of the coefficients obtained by estimating each group separately. This method implies no restrictions on the coefficients whatsoever, and is therefore more flexible than the suggested alternatives. Finally, Pesaran et al. (1999) suggest using PMG, a combination of previously described methods, as it uses both pooling and averaging. Just as in MG, PMG allows intercepts, short-run coefficients and error variances to differ, but analogously to fixed effects, it restricts the long-run coefficients to be equal across groups.

The original Pesaran et al. (1999) paper starts with an autoregressive distributed lag (ARDL) dynamic panel specification, with p being the number of lags for the dependant variable, and q the number of lags for the explanatory variables. The specification takes the following form:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}^{'} X_{i,t-j} + \mu_i + \varepsilon_{it} \text{ , } t = 1,2,...,T \text{ , } i = 1,2,...,N \quad (12)$$

where λ_{ij} are coefficients of the lagged dependent variable, X_{it} is a set of regressors and δ_{ij} is a $(k \times 1)$ vector of its coefficients. Group-specific effects are represented by μ_i while ε_{it} is the error term. This model specification requires that T is large enough, rule of thumb being the possibility to estimate the model for each group separately. Due to the fact that T specified in equation (12) is rather large, there is reason to expect that some variables might not be stationary. In case the variables are integrated of order one, and consequently cointegrated, then we expect that the error term is stationary for all i. Typically, cointegrated variables react to deviations from their long-run equilibrium, and adjust in the short-run. These sorts of deviations and reactions are usually presented as error-correction models that take the form presented by equation (13), obtained by reparameterization of (12).

$$\Delta y_{it} = \phi_i \left(y_{i,t-1} - \theta_i^{'} X_{it} \right) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-1} + \sum_{j=0}^{q-1} \delta_{ij}^{'*} \Delta X_{i,t-j} + \mu_i + \varepsilon_{it} \quad (13)$$

where

$$\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij} \right),\ \theta_i = \frac{\sum_{j=0}^{q} \delta_{ij}}{1 - \sum_{k} \lambda_{ik}},\ \lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im},\ j = 1, 2, ..., p-1, \text{ and } \delta_{ij}^{*} = -\sum_{m=j+1}^{q} \delta_{im}.$$

The error coefficient, or the speed of adjustment term, is presented here as ϕ_i. In case ϕ_i is statistically significant, there is evidence of a long-run relationship, while a negative ϕ_i implies that the variables return to the equilibrium after a deviation in the short-run. Equation (13) reveals that $\theta_i^{'}$ is the vector of long-run coefficients, while λ_{ij}^{*} and $\delta_{ij}^{'*}$ are short-run coefficients of the lagged dependent and explanatory variables, respectively.

4.1.2. PMG Estimation

Pesaran et al. (1999) recommend using maximum likelihood (ML) estimation method for estimating equation (13), as it is nonlinear in its parameters. However, prior to ML estimation, one must make a few assumptions about this specification. Firstly, we rewrite equation (13) by stacking the time-series observations for each group:

$$\Delta y_i = \phi_i \left(y_{i,-1} - \theta_i X_i \right) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,-j} + \sum_{j=0}^{q-1} \Delta X_{i,-j} \delta_{ij}^{*} + \mu_i \iota + \varepsilon_i \quad (14)$$

where $\iota = (1,...,1)^{'}$ is a $(T \times 1)$ vector of ones, and the disturbance term is $\varepsilon_i = \left(\varepsilon_{i1}, ..., \varepsilon_{iT} \right)^{'}$. Next, we assume that the error terms are distributed independently across groups, time, and of the regressors. The last assumption is necessary for a consistent estimate of short-run coefficients, as we allow them to differ across groups. To control for the long-run relationship, and for the adjustment to the long-run equilibrium, it is needed that the speed of adjustment term, ϕ_i, is negative. This is ensured when the model given by equation (12) is stable in its roots that lie outside the unit circle, or that $\sum_{j=1}^{p} \lambda_{ij} z^j = 1$. Then we can say that there is a long-run relationship between the dependant variable and regressors, with θ_i the long-run coefficients, and η_{it} a stationary process:

$y_{it} = -\theta_i X_{it} + \eta_{it}$.

Once we confirm there is a long-run relationship, we can rewrite its coefficients without the group-specific term:

$\theta_i = \theta$ (15).

Having made these assumptions, we can rearrange equation (14):

$$\Delta y_i = \phi_i \xi_i(\theta) + W_i \kappa_i + \varepsilon_i \qquad (16)$$

where

$$\xi_i(\theta) = y_{i,-1} - X_i \theta, \qquad (17)$$

$W_i = \left(\Delta y_{i,-1}, ..., \Delta y_{i,-p+1}, \Delta X_i, \Delta X_{i,-1}, ..., \Delta X_{i,-q+1}, \iota\right)$, and $\kappa_i = \left(\lambda^*_{i,1}, ..., \lambda^*_{i,p-1}, \delta^{*'}_{i,0}, \delta^{*'}_{i,1}, ..., \delta^{*'}_{i,q-1}, \mu_i\right)'$.

Additional to no restrictions on the short-run coefficients, we allow the error variances to differ across groups as well, given by $\operatorname{var}(\varepsilon_{it}) = \sigma_i^2$. Now, the nonlinearity of the parameters θ and ϕ_i from equation (16) implies that we should take a likelihood approach to estimating the described panel. Just for these purposes, it is assumed that the error term is normally distributed. Together with all the assumptions, it is possible to express the likelihood of the panel as the product of each, separate, group-specific likelihoods. Taking logarithms, one gets:

$$l_T(\varphi) = -\frac{T}{2}\sum_{i=1}^{N} \ln\left(2\pi\sigma_i^2\right) - \frac{1}{2}\sum_{i=1}^{N} \frac{1}{\sigma_i^2}\left[\Delta y_i - \phi_i \xi_i(\theta)\right]' H_i \left[\Delta y_i - \phi_i \xi_i(\theta)\right] \qquad (18)$$

where $H_i = I_T - W_i\left(W_i' W_i\right)^{-1} W_i'$, $\varphi = \left(\theta', \phi', \sigma'\right)'$, $\phi = (\phi_1, \phi_2, ..., \phi_N)'$, and $\sigma = (\sigma_1^2, \sigma_2^2, ..., \sigma_N^2)'$. In order to get consistent and asymptotically normal estimators, it is necessary to add some further assumptions that can be found in the original paper (Pesaran, 1999, p. 624-625).

As implied by equation (18), $\varphi = \left(\theta', \phi', \sigma'\right)'$ is estimated in the following order. Maximizing equation (18) with respect to φ gives us estimates of the long-run coefficients, $\hat{\theta}$, and of the error-correction coefficients, $\hat{\phi}_i$:

$$\hat{\theta} = -\left[\sum_{i=1}^{N} \frac{\hat{\phi}_i^2}{\hat{\sigma}_i^2} X_i' H_i X_i\right]^{-1} \times \left[\sum_{i=1}^{N} \frac{\hat{\phi}_i}{\hat{\sigma}_i^2} X_i' H_i \left(\Delta y_i - \hat{\phi}_i y_{i,-1}\right)\right], \qquad (19)$$

$$\hat{\phi}_i = \left(\hat{\xi}_i' H_i \hat{\xi}_i\right)^{-1} \hat{\xi}_i' H_i \Delta y_i \qquad (20)$$

where $\hat{\xi}_i' = \xi_i\left(\hat{\theta}\right) = y_{i,-1} - X_i\hat{\theta}$. Estimators in (19) and (20) are called pooled mean group estimators because they reflect both pooling, inherent in $\hat{\theta}$, and averaging across groups, inherent in $\hat{\phi}_i$. Solving the maximization problem provides the error variance estimate as well:

$$\hat{\sigma}_i^2 = T^{-1}\left(\Delta y_i - \hat{\phi}_i\hat{\xi}_i\right)' H_i\left(\Delta y_i - \hat{\phi}_i\hat{\xi}_i\right). \quad (21)$$

The PMG estimators are computed using the algorithm of back-substitution or in other words, by iterating the obtained estimates. Using the initial estimate $\hat{\theta}$, one can get $\hat{\phi}$ and $\hat{\sigma}_i^2$ from equations (20) and (21), respectively. These estimates can then be replaced into equation (19) to get a new estimate of θ, used to get new estimates of ϕ and σ_i^2, repeated until convergence is achieved. Pesaran et al. (1999) distinguish between stationary and nonstationary regressors, but here we will focus only on the nonstationary part, because our time span is rather long and the regressors we use turn out to be integrated of order one. It is important to note that the ML estimates of long-run and short-run parameters are asymptotically distributed independently of each other, implying that once we get the long-run parameters, we can use them to consistently estimate the short-run and the error-correction coefficients. Although the parameters obtained by iteration are identical to those obtained from full-information maximum likelihood, the covariance matrix is not. Nevertheless, we can recover the covariance matrix for all estimated parameters, because we know the distribution of the PMG parameters. The covariance matrix can be estimated by the inverse of:

$$\begin{pmatrix}
\sum_{i=1}^{N}\frac{\hat{\phi}_i^2 X_i' X_i}{\hat{\sigma}_i^2} & \frac{-\hat{\phi}_1 X_1'\hat{\xi}_1}{\hat{\sigma}_1^2} & \cdots & \frac{-\hat{\phi}_N X_N'\hat{\xi}_N}{\hat{\sigma}_N^2} & \frac{-\hat{\phi}_1 X_1' W_1}{\hat{\sigma}_1^2} & \cdots & \frac{-\hat{\phi}_N X_N' W_N}{\hat{\sigma}_N^2} \\
 & \frac{\hat{\xi}_1'\hat{\xi}_1}{\hat{\sigma}_1^2} & \cdots & 0 & \frac{\hat{\xi}_1' W_1}{\hat{\sigma}_1^2} & \cdots & 0 \\
 & & \vdots & \vdots & & \vdots & \vdots \\
 & & & \frac{\hat{\xi}_N'\hat{\xi}_N}{\hat{\sigma}_N^2} & 0 & \cdots & \frac{\hat{\xi}_N' W_N}{\hat{\sigma}_N^2} \\
 & & & & \frac{W_1' W_1}{\hat{\sigma}_1^2} & \cdots & 0 \\
 & & & & & \vdots & \vdots \\
 & & & & & & \frac{W_N' W_N}{\hat{\sigma}_N^2}
\end{pmatrix}.$$

Additionally, the MG parameters are then simply unweighted means of individual coefficients, or:

$$\hat{\phi}_{MG} = \frac{\sum_{i=1}^{N}\hat{\phi}_i}{N} \text{ and } \hat{Var}\left(\hat{\phi}_{MG}\right) = \frac{\sum_{i=1}^{N}\left(\hat{\phi}_i - \hat{\phi}_{MG}\right)^2}{N-1}.$$

The homogeneity assumption in the PMG estimation or the restrictions on both long-run and short-run parameters in dynamic fixed effects can be tested using a Hausman-type test (Hausman, 1978). The null hypothesis of this test says that the poolability restrictions hold, so in case we cannot reject the null, PMG is the preferred estimator. It is important to run this test because the MG estimator is inefficient if coefficients are actually homogenous in the long-run, while the PMG estimator is efficient and consistent in that case. The same holds for the fixed effects estimators; in case the long-run restriction binds, the fixed effects estimator is inefficient.

4.2. Empirical Specification

Following the theory suggested by Ferrucci (2003) and Berganza et al. (2004) that is presented in section "Theoretical framework", we model sovereign spreads. It has been observed that the dynamics of sovereign spreads complies with theory in the long-run, but demonstrate different dynamics in the short-run. We detect possible long-run and short-run drivers and incorporate them into a nonstationary heterogeneous panel inspired by the pooled mean group estimator of Pesaran et al. (1999). In addition to PMG, we use two alternative estimators, mean group estimator of Pesaran and Shin (1995) and a traditional dynamic fixed effects estimator. We then test the efficiency of all of these estimators and decide on the most efficient and consistent one.
The long-run relationship that we explore is similar to equation (1) and is given by:

$$spread_{it} = \theta_i + \sum_{j=1}^{k} \theta_{ji} LR_{jit} + u_{it} \qquad (22)$$

where k denotes the number of long-run variables, while LR_{it} represents long-run regressors. We assume and test the hypothesis that the long-run variables together with the dependent variable are nonstationary, and cointegrated. If that is the case, then the error term u_{it} is stationary for all i. Setting maximum lags to one, and reformulating equation (13), we get the following ARDL equation:

$$spread_{it} = \lambda_{it} spread_{it-1} + \sum_{j=1}^{k} \delta_{1ji} LR_{jit} + \sum_{j=1}^{k} \delta_{2ji} SR_{jit-1} + \mu_i + u_{it} \qquad (23).$$

The error-correction equation is then given by:

$$\Delta spread_{it} = \phi_i \left(spread_{it-1} - \sum_{j=1}^{k} \theta_{ji} LR_{jit} \right) + \sum_{j=1}^{k} \delta_{2ji} \Delta SR_{it} + \mu_i + \varepsilon_{it} \qquad (24)$$

where $\phi_i = -(1-\lambda_i)$ and $\theta_{ji} = \dfrac{\delta_{1ji} + \delta_{2ji}}{1-\lambda_i}$.

Note that in equation (24) we call the term in differences, SR_{it}, the short-run term, regardless of the fact that it can be the differenced long-run variable or some different variable used to explain short-run dynamics. The term given in brackets represents the long-run relationship with θ_{ji}'s being the long-run elasticities. The speed of adjustment coefficient and short-run elasticities are given by ϕ_i and δ_{2ji}, respectively. Actually, the PMG estimator restricts the long-run coefficients to be equal, so we can rewrite equation (24) such that $\theta_{ji} = \theta_j$:

$$\Delta spread_{it} = \phi_i \left(spread_{it-1} - \sum_{j=1}^{k} \theta_j LR_{jit} \right) + \sum_{j=1}^{k} \delta_{2ji} \Delta SR_{it} + \mu_i + \varepsilon_{it} \qquad (25).$$

We estimate equation (25) using the PMG estimator, and after choosing the baseline model, we test it against two alternative estimators, the MG and the dynamic fixed effects. The reason we use PMG as our preferred estimator is in the nature of the data, economic reasoning, and econometric considerations. Firstly, PMG is a dynamic model, and as such, it reflects the nature of sovereign spreads more realistically. The long-run coefficient homogeneity provides stability of the model, and consequently, corroborates theoretical predictions that coefficients should be identical for all countries. On the other hand, short-run heterogeneity allows for country-specific characteristics and gradual adjustment to the equilibrium, which keeps us from misleading conclusions about the dependent variable (Haque et al., 2000). Lastly, it has been shown that pooled estimators outperform alternative heterogeneous estimators, because increasing efficiency, brought by pooling, offsets the occurrence of biases, brought by short-run heterogeneity across groups (Boyd and Smith, 2000).

4.3. Panel Unit Root Tests

We assume that the long-run variables we use are nonstationary, and in line with that, we have to test these assumptions. We apply five different unit root tests on the dependent variable and the long-run variables that were eventually chosen in the baseline model. The large number of different tests we use is argued by the fact that all of these tests have disadvantages, so using more different tests will lead to robust results (Enders, 1995, p. 243). We perform the following tests: Im-Pesaran-Shin (Im et al., 2003), Fisher-type (Choi, 2001), Levin-Lin-Chu (Levin et al., 2002), Breitung (Breitung, 2000; Breitung and Das, 2005), and Hadri (Hadri, 2000) tests. The first four tests test the null hypothesis that all panels contain a unit root, while the Lagrange multiplier-based Hadri test assumes that all panels are stationary under the null. These panel unit root tests are designed with options of including a trend and fixed effects. Only two tests allow for an unbalanced dataset, Im-Pesaran-Shin and Fisher-type tests, while the other tests require a balanced set of data.
Im-Pesaran-Shin test provides a test statistic based on augmented Dickey–Fuller statistics averaged across sectors. Therefore, the test is based on averages of individual unit root statistics. Fisher-type test however is more general, as it assumes that T can be different for each sector, and it allows that some sectors can contain a unit root, while some others can not. This test combines *p*-values from each sector-specific unit root test. The Levin-Lin-Chu test fits an augmented Dickey-Fuller regression to each sector, using the AIC criterion to find the optimal lag length. The main disadvantage of this test is that it has a common autoregressive factor for all sectors, implying that it does not allow a unit root for one sector and not for the other. Levin-Lin-Chu test with panel-specific means included, is suitable for panels in which the ratio of sectors to time periods tends to zero, therefore in cases where T grows faster than N (as is the case in this study). The main advantages of the Breitung test (a robust version of the Dickey-Fuller *t*-statistic), are that it allows for contemporaneous correlated errors, and performs well with respect to size and power. Finally, the Hadri test is easy to apply, and it has performed well for panel data models with fixed effects, individual deterministic trends and heterogeneous errors across groups.
To sum up, four out of five tests are based on the null of nonstationarity (unit root), while the Hadri test has stationarity defined under the null. Levin-Lin-Chu, Breitung and Hadri tests require a balanced panel, so they were applied to a truncated version of the dataset.

5. Estimation Results

5.1. Panel Unit Root Testing

Prior to estimation, we apply panel unit root tests to the dependent and long-run variables: spread, external debt, current account and international reserves.[6] Results of panel unit root testing, presented in Table 3, show that the first four tests do not reject the null hypotheses of a unit root for the spread, external debt and international reserves. For the current account, only the Breitung test rejects the null, while other three tests can not reject the unit root hypothesis. The Hadri test with its different formulation, suggests that we can reject the null of stationarity for all long-run variables. These results imply that the long-run panel variables are not stationary, and that these variables could be cointegrated, therefore eligible for PMG specification.

Table 3. Panel unit root tests results

			p-values			
Test	Null hypothesis	Alternative hypothesis	Spread	External debt	Current account	International reserves
Im-Pesaran-Shin	All panels contain unit roots	Some panels are stationary	0.994	0.993	0.364	0.998
Fischer	All panels contain unit roots	At least one panel is stationary	0.860	0.847	0.153	0.987
Levin-Lin-Chu	All panels contain unit roots	All panels are stationary	1.000	0.108	0.156	0.843
Breitung	All panels contain unit roots	All panels are stationary	1.000	0.933	0.002	0.671
Hadri	All panels are stationary	Some panels contain unit roots	0.000	0.001	0.000	0.000

Note: The panels include nine countries; the overall sample covers the period from the first quarter of 2001 to the fourth quarter of 2011; Levin-Lin-Chu, Breitung and Hadri tests require a balanced panel and were therefore applied to a truncated version of the dataset.

5.2. Baseline Estimation

We estimate equation (25)[7] using maximum likelihood as presented in Pesaran et al. (1999). We start with a parsimonious version, because the PMG technique uses a big number of parameters that decrease the degrees of freedom. The first specification, presented in column 1 of Table 4, consists only of the most important long-run spread determinants. Note that spreads are defined in logarithms, so the coefficients in the table are semi-elasticities. Results imply that there is a long-run relationship between the variables, because the speed of adjustment coefficient is statistically significant and negative. The external debt, current account and international reserves are all significant long-run variables that have the expected signs. What is found here is that the share of external debt in GDP leads to higher sovereign spreads, a result consistent with theory and empirical work (Edwards, 1984; Alexopoulou et al., 2010; Bellas et al., 2010). On the other hand, international reserves and current account balance seem to work in the opposite direction. Higher international reserves and current account surpluses tend to decrease sovereign spreads, as previously found in Edwards (1984) and Strahilov (2006), respectively.
Following the balance-sheet effect point of view presented earlier in detail, we add our constructed variable to the short-run determinants, and find it to be statistically significant and positive (column 2), a result that accords with Berganza et al. (2004). Since market volatility is an important spread

[6] Different variables that potentially explain the long-run were added to the model, but none proved to be statistically significant. We tried adding variables that account for demographics (the share of citizens that are 65+ years old in total population), development (GDP per capita), primary fiscal balance, institutional framework (Worldwide governance indicators), and capital growth (as measured by gross capital formation). These robustness results are not presented here due to space considerations, but are available upon request.
[7] All models were estimated using a lag of one, though alternative specifications are possible.

driver, but only in the short-run, we also add the volatility index variable (column 3). Market volatility has the expected statistically significant positive impact on sovereign spreads, just as in Ebner (2009). Finally, we add a fiscal variable, but one that affects spreads in the short-run, tax revenues, and also find it to be statistically significant. Rising tax revenues seem to push spreads down, due to the fact that higher taxes persuade investors that sovereign debts will be rapid, but only in the short-run.

All models have appropriate explanatory power, expected signs and justifiable coefficient values. Additionally, all variables are statistically significant at conventional levels and most importantly, the coefficients are robust across models. Model 4 is the broadest model with three long-run determinants, and three short-run drivers: market sentiment and proxies for fiscal (tax revenues) and monetary policy (balance-sheet effect). Moreover, this model has the highest log likelihood and R-squared values that make it our preferred and baseline model.

Table 4. Baseline estimates

	Model (1)	Model (2)	Model (3)	Model (4)
Speed of adjustment				
	-0.139*** [0.000]	-0.169*** [0.000]	-0.120*** [0.000]	-0.170*** [0.000]
Long-run coefficients				
External debt	0.019*** [0.000]	0.020*** [0.000]	0.022*** [0.000]	0.019*** [0.000]
Current account	-0.105** [0.013]	-0.084** [0.013]	-0.078** [0.043]	-0.054* [0.058]
International reserves	-0.033*** [0.000]	-0.040*** [0.000]	-0.046*** [0.000]	-0.037*** [0.000]
Short-run coefficients				
Δ balance sheet		0.084*** [0.009]	0.052** [0.019]	0.065* [0.087]
Δ volatility index			0.754*** [0.000]	0.755*** [0.000]
Δ tax revenues				-0.040*** [0.006]
Number of observations	338	326	326	295
Number of countries	9	9	9	9
Log likelihood	-85.1381	-62.5343	23.0836	42.7068
Within R-squared[#]	0.6777	0.7253	0.833	0.8468
Between R-squared	0.1560	0.0147	0.1974	0.5277
Overall R-squared	0.4215	0.4777	0.6637	0.6709
Hausman test	1.67 [0.645]	3.81 [0.283]	1.83 [0.969]	5.49 [0.704]

[#] R-squared values were obtained from models estimated by fixed effects.

Note: Estimations are performed using the PMG estimator of Pesaran et al. (1999); the reported short-run coefficients and the speed of adjustment are simple averages of country-specific coefficients; all equations include a constant term; p-values are in brackets; ***, **, and * denote significance at 1, 5, and 10 percent confidence level, respectively.

The overall speed of adjustment is equal to -0.170, which implies that half of the adjustment occurs in eight months. However, the PMG framework allows heterogeneity between countries, so Table 5 presents country-specific speed of adjustments for the baseline model. We can see that only for the case of Croatia there is no adjustment to deviations from equilibrium, while in other countries the half-life adjustments range from -0.258 or six months for the Czech Republic to -0.046 or two years and eight months for Hungary.

Table 5. Speed of adjustment coefficients

	Baseline model	Estimated half-life
Bulgaria	-0.064** [0.014]	2y
Croatia	-0.128 [0.184]	-
Czech Republic	-0.258* [0.083]	6m
Hungary	-0.046* [0.062]	2y 8m
Poland	-0.095*** [0.023]	1y 4m
Romania	-0.131*** [0.005]	1y
Serbia	-0.213** [0.028]	7m
Slovak Republic	-0.201*** [0.000]	7m
Turkey	-0.397*** [0.004]	4m

Note: Estimations are performed using the PMG estimator of Pesaran et al. (1999); p-values are in brackets; ***, **, and * denote significance at 1, 5, and 10 percent confidence level, respectively; “y” stands for years, and “m” for months.

As mentioned in the section “Methodology”, we test the preferred PMG specification and the long-run homogeneity restriction using the Hausman test. We compare the PMG estimation to MG and to the dynamic FE, and in that way we test the long-run homogeneity assumption. Hausman test results, as indicated in Table 6, give preference to PMG and confirm that we can impose homogenous coefficients in the long-run, while keeping heterogeneity between countries in the short-run.

Table 6. Tests on the homogeneity restriction

	Pooled mean group (PMG)	Mean group (MG)	Hausman test	Dynamic fixed effects (DFE)	Hausman test

Speed of adjustment					
	-0.170***	-0.279***		-0.656***	
	[0.000]	[0.002]		[0.000]	
Long-run coefficients					
External debt	0.019***	0.016***		0.010***	
	[0.000]	[0.002]		[0.000]	
Current account	-0.054*	-0.032	5.49	0.038***	0.04
	[0.058]	[0.511]	[0.704]	[0.000]	[0.998]
International reserves	-0.037***	-0.030		-0.018*	
	[0.000]	[0.256]		[0.070]	
Short-run coefficients					
Δ balance sheet	0.065*	0.094		0.002**	
	[0.087]	[0.155]		[0.014]	
Δ volatility index	0.755***	0.759***		0.884***	
	[0.000]	[0.000]		[0.000]	
Δ tax revenues	-0.040***	-0.038*		-0.023***	
	[0.006]	[0.063]		[0.006]	
Number of observations	295	295		295	
Number of countries	9	9		9	
Log likelihood	42.7068	64.5318		-184.0232	

Note: All equations include a constant term; p-values are in brackets; ***, **, and * denote significance at 1, 5, and 10 percent confidence level, respectively.

Our baseline estimation results indicate that external debt, international reserves and the current account affect sovereign spreads in the long-run, and that the balance sheet effect, market volatility and tax revenues influence spread dynamics in the short-run. A one percentage point change in the share of external debt in GDP tends to increase the spread by 1.9 percent in the long-run. At the same time, one-percentage point higher shares of current account and international reserves in GDP, decrease spreads by 5.4 and 3.7 percent respectively. In the short-run, we have found that the balance sheet effect is positive and stronger than any of the long-run coefficients. It implies that a one percentage point increase in the debt service cost in local currency terms, increases spreads by 6.5 percent. One percentage point change in tax revenues decreases spreads by 4 percent, while a 100 percent change in the market volatility index, leads to a 75.5 percent jump in sovereign spreads.

5.3. Robustness

Following Berganza et al. (2004), we run robustness checks for our baseline model (presented in the first column of Table 7). Model 2 presents the baseline model with the annual growth rate of exports added to the short-run determinants. The reason for including exports is to test for omitted variables, i.e. to check if there is a competitiveness effect in the aftermath of exchange rate depreciation. If exports increase significantly after the exchange rate depreciates, and by that affect sovereign spreads, then the balance sheet effect could be offset. We find that exports are not statistically significant, and that the balance sheet coefficient stays almost unchanged when we add exports to the specification. This implies that we do not need to keep the exports variable.

We also add external debt to our short-run determinants, to detect is the balance sheet variable significant only due to external debt accumulation, and not due to the presence of a real balance sheet effect. Thereby, we test the assumption of predetermined debt (Berganza et al., 2004). Results in the third column suggest that the external debt variable is not statistically significant, and that the balance sheet coefficient stays almost unaffected. From this we conclude that the increase in the size of external debt is not important for changes in spreads, but that spread movements are affected by increases in debt burden, caused by exchange rate movements (or balance sheet effects).

Finally, we tackle the question of the simultaneity bias.[8] As emphasized in Berganza et al. (2004), the equation we estimate may be only one of possible equations that determine the equilibrium. For example, the direction assumed here, namely that exchange rates affect the debt burden, might just be reversed. In that case, our balance sheet estimate is only a reduced one, and it can not be said that it reflects a true balance sheet effect on the cost of sovereign credit. To solve this problem, we need an instrument for our constructed variable, more specifically, an instrument for the real exchange rate change, while we leave external debt in the definition, as it is assumed to be predetermined. Berganza et al. (2004) suggest using inflation because inflation and real exchange rate are correlated, and inflation is not supposed to affect spreads on external debt (for evidence on this, see Figure 2 in section “Data”. We construct a new variable, “external debt*inflation”, and replace the balance sheet variable with its instrument. Column 4 of Table 7 presents the estimated coefficients, and suggests that the alternative variable is not statistically significant.

Table 7. Robustness checks for the baseline model

	Model (1)	Model (2)	Model (3)	Model (4)
Speed of adjustment				
	-0.170***	-0.152***	-0.207***	-0.164***
	[0.000]	[0.000]	[0.002]	[0.000]
Long-run coefficients				
External debt	0.019***	0.020***	0.018***	0.019***
	[0.000]	[0.000]	[0.000]	[0.000]
Current account	-0.054*	-0.067**	-0.007	-0.059*
	[0.058]	[0.040]	[0.694]	[0.053]
International reserves	-0.037***	-0.039***	-0.038***	-0.038***
	[0.000]	[0.000]	[0.000]	[0.000]
Short-run coefficients				
Δ balance sheet	0.065*	0.083*	0.103*	
	[0.087]	[0.087]	[0.074]	
Δ volatility index	0.755***	0.750***	0.748***	0.749***
	[0.000]	[0.000]	[0.000]	[0.000]
Δ tax revenues	-0.040***	-0.040***	-0.037**	-0.040***
	[0.006]	[0.006]	[0.025]	[0.007]
Δ export		-0.000		
		[0.864]		

[8] We ran two additional robustness checks. One excluding Bulgaria (since Bulgaria introduced a currency board exchange rate regime and might not reflect such a strong balance sheet effect), and one without extreme balance sheet values (without 5% extreme values). We find that the coefficients’ significance, signs, and values stay almost unchanged, once again conforming that our chosen model is preferred. These results are not presented here, but are available upon request.

Δ external debt			0.001 [0.707]	
Δ external debt*inflation				-0.002 [0.148]
Number of observations	295	295	295	295
Number of countries	9	9	9	9
Log likelihood	42.7068	50.9645	50.7357	42.0037
Within R-squared[#]	0.8468	0.8665	0.8731	0.8462
Between R-squared	0.5277	0.0304	0.1514	0.0008
Overall R-squared	0.6709	0.6732	0.7768	0.6468
Hausman test	5.49 [0.704]	1.42 [0.700]	7.70 [0.565]	3.29 [0.915]

[#] R-squared values were obtained from models estimated by fixed effects.

Note: Estimations are performed using the PMG estimator of Pesaran et al. (1999); the reported short-run coefficients and the speed of adjustment are simple averages of country-specific coefficients; all equations include a constant term; p-values are in brackets; ***, **, and * denote significance at 1, 5, and 10 percent confidence level, respectively.

6. Conclusion

This study uses different theoretical and empirical sources to build a model of sovereign spread determinants that enables us to empirically test a relationship between spreads and financial imperfections that appear in the form of "original sin", a widely spread emerging market phenomenon. We investigate a positive relationship between a country's risk premium and balance sheet effects – increasing debt servicing costs caused by exchange rate depreciation. We apply this method to nine European emerging economies for the 2001-2011 period. We use a small open economy model and extend it with the collateral value concept of Kiyotaki and Moore (1997), and recent empirical findings on the balance sheet effect of Berganza et al. (2004). We place the model into a dynamic error correction setting introduced by Pesaran et al. (1999), and allow the short-run determinants to differ across countries, while we leave the long-run parameters to be equal for all countries. This allows more flexibility, brought by differentiation between short-run and long-run, but also provides estimation advantages, such as improved efficiency and model performance.

The results of the empirical model corroborate the differentiation between the short-run and the long-run, and suggest that there exists a strong positive relationship between spreads and balance sheet effects in the short-run. Besides the balance sheet effect, we find that market volatility and tax revenues also affect sovereign spreads in the short run. Estimation suggests that a 100 percent rise in the volatility index leads to a 75.5 percent jump in spreads, while a one percentage point change in tax revenues reduces spreads by 4 percent. On average, half of this deviation from long-run equilibrium is corrected in eight months. In the long-run, spreads increase by 1.9 percent when the share of external debt in GDP rises by one percentage point, but tend to decrease by 5.4 and 3.7 percent when the share of current account and international reserves in GDP increases by one percentage point.

Our empirical results have serious policy implications, as they emphasize the role and strength of short-run spread determinants, next to the extensively studied long-run drivers. We find evidence that external factors, such as market volatility, and balance sheet effects caused by financial imperfections of the inability to issue debt in local currency, can be responsible for severe short-run changes in sovereign spreads. In order to avoid significant spread volatility that could result in liquidity problems of refinancing sovereign debt, countries should avoid sudden and large real exchange rate depreciations, when their foreign currency external debt is large (as previously suggested by Hausmann et al. (2001), and Eichengreen et al. (2003)). Although European emerging countries did not experience larger exchange rate depreciations in the recent financial crisis, history has taught us that these events are not rare, and that countries can stand on the verge of devaluation for years before it finally comes about (Reinhart and Rogoff, 2009).

Besides external debt, further research should also include domestic debt denominated in foreign currency, as number of countries issue domestic debt that is indexed to the exchange rate. This would ensure a comprehensive measure of euroization, and provide a more realistic picture of the balance sheet effect. However, the primary research focus should be on building a theoretical model of the relation between country's risk premium and total debt euroization. As far as we are aware, this issue has only been investigated empirically so far.

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APPENDIX

Table A1. Variable description

Variable	Description	Expected sign	Data source
Dependent variable			

Spread	JP Morgan Euro EMBI Global indices equal the returns for US dollar-denominated Brady bonds, loans, and Eurobonds with an outstanding face value of at least $500 million, minus returns for U.S. Treasury bonds with similar maturity. The variable is in logarithms.		JP Morgan
Long-run determinants			
External debt	Gross external debt in millions of Euros, divided by real GDP (2005=100) in millions of Euros, and multiplied with 100.	positive	National central banks
Current account	Current account balance in millions of Euros, divided by real GDP (2005=100) in millions of Euros, and multiplied with 100.	negative	Eurostat and IMF IFS
International reserves	Official international reserves at the end of the quarter in millions of Euros (excluding gold) divided by real GDP (2005=100) in millions of Euros, and multiplied with 100.	negative	IMF IFS
Short-run determinants			
Balance sheet	Equals the product of external debt (see above) and the year-on-year difference in the real exchange rate, where the real exchange rate is defined as the ratio of the nominal bilateral exchange rate (local currency for 1 Euro) and the GDP deflator in national currency (2005=100), divided by 100.	positive	National central banks, Eurostat and own calculation
Volatility index	CBOE volatility index of investor sentiment and market volatility, calculated as an average quarterly value. The variable is in logarithms.	positive	Chicago Board Options Exchange
Tax revenues	General government tax revenues in millions of Euros, divided by real GDP (2005=100) in millions of Euros, and multiplied with 100.	negative	IMF IFS and national treasuries
Export	Export in Euros, calculated as a year-on-year growth rate.	negative	Eurostat and IMF IFS
External debt	Gross external debt in millions of Euros, divided by real GDP (2005=100) in millions of Euros, and multiplied with 100.	positive	National central banks
External debt*inflation	Equals the product of external debt (see above), and the GDP price index year-on-year growth rate in Euros (2005=100), divided by 100.	positive	National central banks, Eurostat and own calculation

Note: The sample covers the following emerging economies: Bulgaria, Croatia, Czech Republic, Hungary, Poland, Romania, Serbia, Slovak Republic, and Turkey.