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EMERGING MARKET SOVEREIGN SPREADS, GLOBAL FINANCIAL CONDITIONS AND US MACROECONOMIC NEWS

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Abstract

This paper investigates the impact of global financial conditions, US macroeconomic news and domestic macroeconomic fundamentals on the evolution of Emerging Market Bond Index EMBI spreads for a panel of 18 emerging markets (EM) using daily data. To this end, we employ not only the conventional panel data estimation procedures, but also the recently developed common correlated effects panel mean group method which incorporates heterogeneity by allowing country-specific coefficients whilst accounting for the effects of common global shocks such as contagion. The results strongly suggest that the long-run evolution of EMBI spreads depends on external factors such as changes in global liquidity conditions, risk appetite and crises contagion. Domestic macroeconomic fundamentals, proxied by sovereign country ratings, are also found to be important in explaining the spreads. The results from panel equilibrium correction models suggest that EMBI spreads also respond substantially to US macroeconomic news, as well as changes in the Federal Reserve's target interest rates. However, the magnitude and the sign of the effect of US macroeconomic news crucially depend on the state of the US economy, such as the presence of inflation dominance.

ملخص

تدرس هذه الورقة- بناءا علي بيانات يومية- مدي تأثير الظروف المالية العالمية والتقارير المتعلقة بالإقتصاد الكلي للولايات المتحدة الأمريكية وأسس الإقتصاد الكلي المحلي علي انتشار مؤشر سندات الأسواق الناشئة علي عينة من 18 سوقا ناشئة.

ولهذا الغرض لا نستخدم الإجراءات التقليدية لتقدير البيانات فحسب بل نستخدم كذلك المنهج الشائع والذي طور حديثا لمتوسط مجاميع الـتأثيرات المتلازمة الشائعة وهو منهج يقوم علي دمج الخواص المتباينة عن طريق استخدام معاملات تخص كل دولة مع تفسير تأثيرات الصدمات العالمية الشائعة مثل انتقال العدوي بين الإقتصاديات.

وتوحي النتائج الناجمة بأن النمو طويل الأجل لمؤشر سندات الأسواق الناشئة يعتمد علي عوامل خارجية مثل التغيرات في ظروف السيولة العالمية والميل نحو المخاطرة وعدوي الأزمات.

وقد برزت أهمية أسس الإقتصاد الكلي المحلي في تفسير هذا الانتشار استرشادا بما ورد في التقديرات الخاصة ببعض الدول ذات السيادة.

وتشير نتائج نماذج تصويب التوازن إلي أن انتشار مؤشر سندات الأوراق المالية يستجيب وبقوة للتقارير الواردة عن الإقتصاد الكلي الأمريكي والتغيرات في سعر الفائدة المستهدف الذي يحدده بنك الإحتياطي الفدرالي الأمريكي, إلا أن مدي تأثير التقارير الواردة عن الإقتصاد الكلي الأمريكي يعتمد بالأساس علي حالة الإقتصاد الأمريكي مثل هيمنة التضخم.

1. Introduction

International financial integration presents countries with both opportunities and challenges. One of the basic challenges of the deepening financial integration over the last decades has been the increase in the role of global financial conditions on the macroeconomic fluctuations of emerging markets (EM). Calvo (2002, 2005), for instance, argue that capital mobility has made EM economies more vulnerable to exogenous shocks coming from global capital markets. Along the same lines, according to some recent studies including Neumeyer and Perri (2005), Uribe and Yue (2006), Mackowiak (2007) and Izquierdo, Romero and Talvi (2007), real output fluctuations in EM economies have often been triggered by changes in global financial conditions represented by international interest rates or US monetary policy shocks.

The cost of borrowing faced by EM economies in international financial markets is often represented by emerging market bond index (EMBI) spreads. The EMBI spread, which is the difference between the yields on emerging country sovereign bonds and bonds issued by a government of the industrialized world with identical currency denomination and maturity, is a standard measure of sovereign default risk. Movements in EMBI spreads have usually been associated with large business cycle swings in EM economies (Neumeyer and Perri, 2005; Izquierdo *et al*., 2007). Furthermore, some key financial variables including exchange rates and domestic interest rates also tend to be driven by EMBI spreads as shown by Blanchard (2004) and Favero and Giavazzi (2004) for Brazil and by Özatay (2005) for Turkey.

Understanding the contributions of global (external) and domestic factors in the evolution of EMBI spreads has crucial policy implications. Evidence suggesting that the spreads are mainly driven by domestic macroeconomic fundamentals can be interpreted as "good news" for EM economies implementing sound macroeconomic policies. This is because such a policy stance should decrease the default risk and thus the spread. However, evidence suggesting that the spreads are mainly driven by global financial conditions and macroeconomic performance of developed countries may imply that EM economies are highly vulnerable to external shocks. A significant increase in EMBI spreads (default risk) due to a tightening cycle in industrial countries, for instance, could lead to a rise in the debt-to-GDP ratio by depreciating the domestic currency and raising domestic interest rates, which has the potential to ignite a self-fulfilling prophecy as in the second generation crisis models, even in the presence of sustainable domestic macroeconomic fundamentals. Furthermore, based on positive correlations between domestic interest rates, exchange rates and EMBI spreads, Favero and Giavazzi (2004) and Blanchard (2004) argue that the central bank of an EM country may be vulnerable to an exogenous upward shift in the spread as this may lead to domestic currency depreciation and deterioration in inflation expectations. All these do not necessarily reduce the crucial importance of domestic fundamentals even if the spreads are predominantly determined by external conditions because domestic vulnerabilities provide the main magnifying mechanisms through which the impacts of exogenous shocks are transmitted.

The literature on the determinants of EMBI spreads is considerable and growing. One line of literature maintains that shocks originating in developed countries are the main drivers of the EMBI spreads and thus emphasizes external factors, such as international interest rates, global risk aversion and liquidity conditions (Kamin and Kleist, 1999; Calvo, 2002; García-Herrero and Ortíz, 2006; Gonzales-Rozada and Levy-Yeyati, 2006). A related literature stresses contagion effects of shocks originating in other EM economies on financial portfolios (Broner, *et al*., 2006) or on EMBI spreads (Kaminsky and Schmukler, 2002). Another line of the literature focuses on the effects of domestic economic fundamentals, indicating the importance of country default risk or creditworthiness in the determination of the country spreads (Arora and Cerisola, 2001; Kamin, 2002 and Çulha *et al*., 2006).

Financial markets often react to macroeconomic news as documented by a large body of literature – the bulk of which is based on the advanced industrial countries. The number of studies considering the impact of daily or intra-daily industrial country macroeconomic news on EM financial asset returns is extremely limited and the recent exceptions include Robitaille and Roush (2006); Wongswan (2006) and Andritzky, *et al.* (2007). Our study aims to contribute to this literature as well, by investigating the impacts of US scheduled macroeconomic announcements and surprises on EMBI spreads. Previous literature typically treats the interpretation of a given piece of macroeconomic news as "good" or "bad", regardless of the state of the economy. Our paper seeks to extend this literature by considering the case that EMBI spreads react differently to a piece of US macroeconomic news depending on the presence of inflationary concerns in the US economy.

In this context, this paper aims to investigate the impacts of global financial conditions, US macroeconomic news and domestic macroeconomic fundamentals on the evolution of EMBI spreads for a panel of 18 EM economies by using daily data dating from December 31, 1997 to December 31, 2006. The literature often employs conventional panel data estimation procedures without providing the individual country estimates, even though these often contain useful information. Therefore, in this study we consider not only the conventional panel data and panel cointegration procedures, but also employ the panel mean group estimation procedure proposed by Pesaran and Smith (1995) that incorporates heterogeneity by allowing for both country-specific intercepts and slopes. Moreover, omitted common variables or global shocks such as contagion may induce cross-section dependence and lead to inconsistent regression coefficient estimates if they are correlated with the explanatory variables. To account for the cross-sectional dependence in the data, we employ the common correlated effects mean group estimator by Pesaran (2006) which is robust in a general non-stationary framework where the regressors and errors share common factors (for example global shocks). The empirical modeling contributions of this paper includes panel equilibrium correction estimates of the EMBI spreads which allow us to assess the adjustments to deviations from the long-run equilibrium relationship along with the short-run impact of the stationary US macroeconomic news surprises and changes in the Federal Reserve (FED) target rates.

The plan for the rest of the paper is as follows. The following section presents a brief review of the empirical literature on the determinants of EMBI spreads. This section also argues that the interpretation (and thus the impact) of a macroeconomic news as "good" or "bad" may not be invariant to the state of the economy. Section III is devoted to the empirical analysis of the determinants of EMBI spreads. In section III.1. we report the individual country estimates and the corresponding panel mean group estimation results. Section III.2. takes into account the potential cross-sectional dependence in the data and presents the results for the common correlated effects mean group estimation procedure. Panel cointegration and equilibrium correction mechanisms are also considered in this section. Section III.3. is devoted to the investigation of the impacts of US macroeconomic announcements and surprises on the EMBI spreads. Finally, section IV presents some concluding remarks.

2. The Determinants of EMBI Spreads

A general model for the empirical determinants of sovereign spreads (S) can be defined as

$$S_{it} = c + \theta X_t + \Phi Z_{it} + u_{it} \qquad (1)$$

where *c* is a constant term, X and Z are vectors of foreign and domestic explanatory variables, respectively, θ and Φ are the transposes of the corresponding coefficient vectors and u is the disturbance term. The subscripts i and t stand for country and time. In the literature, the set of variables in X contains an industrial country (namely the US), interest rates or the FED target rate to proxy global liquidity and some alternative measures to capture global risk appetite or financial conditions. The spread of high yield corporate bonds in developed markets or the volatility implicit in US stock options (*VIX*) compiled by the Chicago Board Options Exchange are taken as measures of risk appetite of international investors –or alternatively the price of risk (Gonzales-Rozada and Levy-Yeyati, 2006). As argued by Kamin and Von Kleist (1999), country spreads are expected to increase with international interest rates since increases in these rates deepen the borrowing country debt burden and the probability of debt default, thus raising the risk premium. Furthermore, increases in international interest rates can decrease the risk appetite of investors, reducing the demand for risky assets and thus increasing the country spread. It is worth noting that the distinction between "liquidity" and "risk appetite" impacts of international interest rates and high yield spreads may not be very clear as the former implies the latter, or *vice versa*. Consequently, long-term US interest rates and high yield spreads are also often used to proxy the global risk appetite (Dailami, *et al.,* 2005) and global liquidity (Gonzales-Rozada and Levy-Yeyati, 2006; Fostel and Kaminsky, 2007), respectively.

The results of the empirical literature on the impact of international interest rates appear to be inconclusive. Cline and Barnes (1997), Kamin and Von Kleist (1999) and Eichengreen and Mody (2000) suggest that the effect of US interest rates on new-issue bond spreads (launch spreads) are either statistically insignificant or theoretically inconsistent with a negative coefficient. Eichengreen and Mody (2000) justify the estimated negative interest rate coefficient by arguing that a reduction in the US government bond yield appears to increase the supply of emerging countries' sovereign bonds, thereby raising sovereign spreads. Arora and Cerisola (2001), however, find that the impact of the long-term US interest rates is significantly positive when spreads for bonds actively traded in secondary markets are considered instead of launch spreads. Arora and Cerisola (2001) further argue that the FED target rate, which is a direct measure of US monetary policy, tends to positively influence sovereign spreads. Recent research often considers secondary market spreads and finds that both domestic and international factors play a role in their evolution. According to Dailami *et al.* (2005) the impact of US interest rates and high yield spreads increases significantly with the level of indebtedness of the borrowing country and is not invariant to the contagion effects of crises. This result is consistent with Kaminsky and Schmukler (2002) suggesting that economic fragility, captured by country ratings, makes countries more sensitive to changes in international markets. Kaminsky and Schmukler (2002) find that changes in US short-term interest rates increase country spreads, and that this impact is more severe in countries with low ratings. The results by Gonzales-Rozada and Levy-Yeyati (2006) suggest that the variability of emerging market spreads is significantly explained by global financial conditions such as the spread of high yield corporate bonds in developed markets, 10-year US Treasury rates and systemic crisis, representing the risk appetite, global liquidity and contagion, respectively.

The set of variables in Z in (1) contains domestic economic fundamentals indicating default risk or creditworthiness of the country. Country debt, current account deficit, net foreign assets, fiscal balance and gross reserves (all expressed as ratios to the GDP), debt default history, debt service ratios, sovereign credit ratings, terms of trade volatility are among the most commonly employed domestic default indicators. Studies considering country specific variables, including Cline and Barnes (1997); Kamin and von Kleist (1999); Eichengreen and Mody (2000); Arora and Cerisola (2001); Kaminsky and Schmukler (2002); Dailami *et al.*, (2005) and Çulha *et al.* (2006), all find that domestic macroeconomic fundamentals are significant determinants of sovereign spreads.

Financial markets often react to macroeconomic news as documented by a wide and growing literature, the bulk of which is based on the US economy or advanced industrial countries. Gürkaynak *et al.* (2005) find that the US short-term interest rate increases (decreases) when releases that are pro-cyclical (countercyclical) have a higher realized value than expected. The results by Faust *et al.* (2007) suggest that stronger than expected U.S. real activity announcements tend to appreciate the dollar and raise interest rates in the U.S. Similarly, Clarida and Waldman (2007) show that higher than expected inflation appreciates exchange rates in inflation targeting countries implementing a Taylor rule. To date, the number of studies considering the effects of industrial country macroeconomic news using high frequency data on emerging market countries' financial asset returns is very limited and the recent notable exceptions include Robitaille and Roush (2006) and Wongswan (2006), Andritzky, *et al.* (2007). The results by Robitaille and Roush (2006) suggest that US macroeconomic surprises and FOMC interest rate increase announcements prompted an increase in the Brazilian bond yield spread and a decline in the stock price index. Wongswan (2006) finds that macroeconomic announcements in the US and Japan have a significant impact on intraday return volatility of Korean and Thai equity markets. Andritzky, *et al.,* (2007) find that macroeconomic announcements basically affect the volatility of emerging market bond spreads by reducing uncertainty. In all these studies, higher than expected real releases are interpreted as "good news" for the strength of the US economy.

However, interpretation of macroeconomic news surprises as "good" or "bad" for a given financial asset variable of interest may not be invariant to the state of the economy. Results of some recent studies based on the US or advanced industrial country data provide empirical support for this view. For the US, McQueen and Roley (1993) find that, when the economy is strong, the stock market responds negatively to news about higher real economic activity. Boyd *et al.* (2005) consider the US unemployment data and find that an announcement of rising unemployment is good news for stocks during economic expansions and bad news during economic contractions. Andersen *et al.* (2007) consider real time interactions between U.S., German and British stock, bond and foreign exchange markets and find that equity markets react differently to the same news depending on the state of the U.S. economy, with bad macroeconomic news having a positive impact during expansions and negative impact during recessions.

3. Empirical Analysis

EMBI spreads can be specified as determined by domestic macroeconomic fundamentals (Z) and variables representing global financial conditions (X) as already discussed in the context of eq. (1). In this study, we consider daily data which indeed restrict severely the availability of data for domestic macroeconomic variables. Following the literature using high frequency data[1] we consider country credit ratings as a proxy for the domestic macroeconomic fundamentals. As shown by Cantor and Packer (1996, pp.49), "sovereign ratings effectively

[1] See, Eichengreen and Mody (1998), Kamin and Kleist (1999), Dailami, *et al.* (2005), Gonzales-Rozada and Levy-Yeyati (2006) and Andritzky, *et al*. (2007) for studies using country spreads as a proxy for domestic macroeconomic fundamentals.

summarize and supplement the information contained in macroeconomic indicators". In the same vein, Afonso, Gomes and Rother (2007) find that the core set of variables that are relevant for the determination of the ratings include real GDP, government debt, government effectiveness, external debt, external reserves and default history, which are indeed among the main fundamentals explaining sovereign spreads.

Global financial conditions are proxied by the volatility implicit in US stock options (*VIX*) compiled by the Chicago Board Options Exchange as a measure of risk appetite of international investors –or alternatively the price of risk (Gonzales-Rozada and Levy-Yeyati, 2006). For robustness, following Blanchard (2004) and Favero and Giavazzi (2004), we also consider the spread of US corporate bonds with a Moody's rating of Baa with a maturity of 10 years over and 10-year US treasuries (HYS) as an alternative measure[2] of global risk appetite and thus liquidity conditions.

We start by estimating the following equation for each country (i) in our sample:

$$s_{it} = \gamma_{0i} + \gamma_{1i} rt_{it} + \gamma_{2i} vix_t + u_{it} \qquad (2)$$

where s_i is the log of the EMBI+ spreads provided by JP Morgan[3], rt_i is the log of sovereign ratings by Standard and Poor's (S&P) which cover changes in the actual ratings and rating outlooks[4], and vix is the log of the VIX index. The sample covers 18 emerging market countries (Argentina, Brazil, Bulgaria, Colombia, Ecuador, Egypt, Mexico, Malaysia, Morocco, Panama, Peru, Philippines, Poland, Russia, South Africa, Turkey, Ukraine and Venezuela) for the period from December 31, 1997 to December 29, 2006 (period coverage varies across countries, as reported in Table 1).

3.1. Individual Country and Panel Mean Group Estimations

The recent empirical studies on the determinants of the country spreads often employ fixed effects estimation procedures to allow heterogeneity between the panels of countries under consideration. These methods, however, impose a common slope coefficient disregarding the information provided by the county-specific coefficients. Therefore, in this paper, we also employ the panel mean group (MG) method developed by Pesaran and Smith (1995) which permits heterogeneity in both intercept and slope coefficients. Phillips and Moon (1999) show that the cross-sectional variation in a non-stationary panel may be helpful in obtaining consistent estimates of a long-run average parameter even if there is no time series cointegration at the individual level. As argued by Coakley *et al.* (2006), this insight justifies the use of the MG procedure which provides consistent estimates for non-stationary, heterogeneous panels. Furthermore, standard t-tests for the MG estimator based on the N(0,1) distribution have reasonably good size properties irrespective of I(0) or I(1) errors as shown by Coakley *et al.* (2006).

[2] The logs of the VIX and HYS variables are found to be both integrated of order one but are also cointegrated. Consequently, these variables are not considered jointly in our long-run equation specifications.

[3] The EMBI+ index by JP Morgan covers the US dollar and other external currency denominated Brady bonds, loans, Eurobonds, and local market instruments. The details for the index are provided by JPMorgan (2004).

[4] The assignment of numerical values to credit ratings is as in Kamin and Kleist (1999), with 1 being the worst credit risk and 22 the best. Following Gonzales-Rozada and Levy-Yeyati (2006) we interpret the outlook as a five-notch grading scale around the credit rating: positive, positive watch, neutral, negative watch, and negative. The outlook –augmented ratings are computed by giving each notch a 0.2 value and adding to the credit rating. We also considered the ratings provided by Institutional Investor as an alternative proxy for domestic factors and obtained virtually the same results presented in this paper. The results obtained using these ratings are reported despite the high correlation between the ratings (Afonso, *et al*., 2007) and the evidence preferring the S&P rating (Gande and Parsley, 2005).

To obtain the MG estimators, we first estimate equation (1) for each of the countries. The MG estimator ($\hat{\gamma}_{\mathrm{MG}}$) and its standard error (se($\hat{\gamma}_{\mathrm{MG}}$)) for N cross-sectional units, are calculated as follows:

$$\hat{\gamma}_{\mathrm{MG}} = \sum_{i=1}^{N} \hat{\gamma}_i / N \text{ and se}(\hat{\gamma}_{\mathrm{MG}}) = \sigma(\hat{\gamma}_i)/\sqrt{N}$$

where $\hat{\gamma}_i$ and $\sigma(\hat{\gamma}_i)$ are the estimated individual country time-series coefficients and their standard deviations, respectively.

Table 1 reports the OLS estimates of the equations for each of the countries. The table also reports the augmented Dickey-Fuller (ADF) statistics to test the non-stationarity of equation residuals[5] (Engle and Granger, 1987). The results suggest non-rejection of the null of no-cointegration for all the countries except Philippines. An increase in the price of risk (an increase in vix) substantially and significantly increases the EMBI spreads for each of the countries. Better domestic macroeconomic fundamentals, as represented by sovereign ratings, lead to a decrease in EMBI spread for all the countries except Philippines.

The panel MG method yielded the following results (standard errors in parentheses):

$s_{it} = 8.315 - 2.00rt_{it} + 0.998vix_t$

(1.273) (0.577) (0.105)

According to the panel MG results, both the domestic fundamentals and global financial conditions are significant in explaining the spreads.

Common global shocks which are not fully represented by the global liquidity condition and risk appetite variables such as VIX or HYS arising from contagion of a crisis in one or a group of EM countries or from shocks originated in financial centers may induce cross-section dependence in the data and thus lead to inconsistent regression coefficient estimates if they are correlated with the explanatory variables. To account for the cross-sectional dependence in the data, we employ the common correlated effects mean group (CCE-MG) estimator by Pesaran (2006). The CCE-MG estimator yields consistent estimates also in the presence of common factors and appears to be the most efficient (Kapetanios and Pesaran, 2007) and robust to alternative hypotheses of non-stationarity of variables (Coakley *et al.*, 2006).

The CCE-MG procedure suggests approximating the linear combinations of the unobserved factors by cross section averages of the dependent and explanatory variables and then estimating the regressions of interest augmented with these cross section averages. Therefore, to obtain the CCE-MG estimator, we estimate the following equation for each country (i):

$$s_{it} = \gamma_{0i} + \gamma_{1i}rt_{it} + \gamma_{2i}vix_t + c_{1i}m\_rt_t + c_{2i}m\_s_t + u_{it} \quad (3)$$

In (3) m_rt and m_s denote the cross-sectional means of the ratings (rt) and EMBI spreads (s). Note that, the coefficients of the cross–sectional means (CSMs) need not have any economic meaning; their augmentation simply aims to improve the coefficient estimates of interest. However, in our specific case, the CSMs may contain some important information for the evolution of our main variable of interest –EMBI spreads. It is arguably plausible to represent the effect of common global shocks such as contagion by the CSMs of EMBI

[5] Table A1 of the Appendix presents the ADF test statistics for the individual, country-specific variables. All the country specific variables appear to be integrated of order one. The ADF statistics (lag lengths) for vix, Δvix, hys and Δhys are –0.70(3), -31.9(2), -0.71(2) and –39.4(1), respectively. Accordingly, the ADF statistics in Table 1 can be interpreted as valid to test for the null of no cointegration between the variables in the corresponding equations (Engle and Granger, 1997).

spreads. Therefore we expect the estimated c_{2i} to be positive. To the extent that, the ratings are determined solely by domestic macroeconomic fundamentals, the impact of the CSMs of the ratings for the spread of the country may be ambiguous.

Table 2 reports the estimation results for equation (3) for each of the countries. The ADF test results suggest that all the equations, except for Philippines, can be interpreted as representing the long-run equilibrium relationships. The results are essentially the same as those reported in Table 1 except the cases that the country rating variable coefficients become positive (for Egypt, Panama, Venezuela and Poland), and vix coefficients become negative (for Argentina and Malaysia) when the equations are augmented with the CSMs of ratings and spreads.

The common correlated effects panel mean group (CCE-MG) method yielded the following results (standard errors in parentheses):

$$s_{it} = \underset{(2.575)}{0.870} - \underset{(0.525)}{1.187}rt_{it} + \underset{(0.086)}{0.505}vix_t + \underset{(0.107)}{0.554}m\_s_t + \underset{(0.786)}{1.331}m\_rt_t$$

All the coefficients except that for m_rt are strongly significant and theory-consistent. The statistical insignificance of the m_rt coefficient may not be unexpected as the impact of the CSM ratings on the individual country spreads may be negative or positive depending on their relative strength of domestic fundamentals to the rest of the countries. The contagion impact of crises or financial turbulence in one, or a group of EM countries appears to be an important determinant of EMBI spreads. The overall impact of the external factors represented by vix and m_s_t appears to be comparable with the effect of domestic macroeconomic fundamentals. The significance of the external factors arising from the interrelated global liquidity conditions, risk appetite and crises contagion is consistent with the recent findings of Gonzales-Rozada and Levy-Yeyati (2006) suggesting that “a large fraction of the variability of emerging market bond spreads is explained by the evolution of global factors” and thus EM countries “do remain vulnerable to sudden shifts in financial market conditions”.

To check the robustness of our results, we also consider high yield spread (HYS) as an alternative measure of global risk appetite and thus liquidity conditions. Following Blanchard (2004) and Favero and Giavazzi (2004), we define HYS as the spread of US corporate bonds with a Moody’s rating of Baa with a maturity of 10 years over and 10-year US treasuries (HYS). Table A2 and Table A3 in the Appendix respectively reports individual country estimates of equations (2) and (3) with hys instead of vix. The results are virtually the same as those reported in Tables 1 and 2. The panel mean group (PMG) and the common correlated effects panel (CCE-MG) method yielded the following results:

$$s_{it} = \underset{(2.834)}{15.133} - \underset{(1.185)}{3.911}rt_{it} + \underset{(0.206)}{1.126}hys_t$$

and

$$s_{it} = \underset{(2.514)}{0.896} - \underset{(0.556)}{1.806}rt_{it} + \underset{(0.143)}{0.721}hys_t + \underset{(0.105)}{0.724}m\_s_t + \underset{(0.887)}{2.012}m\_rt_t$$

The magnitudes of the estimated coefficients for hys in the equations are only slightly and insignificantly greater than those obtained with vix, suggesting the robustness of our result with respect to the use of an alternative indicator for global financial conditions. The coefficient of rt considerably increases in the equations with hys. Furthermore, the cross-sectional means of the ratings (m_rt) become statistically significant in the CCE-MG equation with hys. However, the main message from the use of these two alternative global indicators, namely the crucial importance of external factors in determining the EMBI spreads, remains empirically valid and robust.

3.2. Panel Cointegration and ECM Estimations

The recent literature, including Gonzales-Rozada and Levy-Yeyati (2006), often employs panel data estimation procedures in investigating the determinants of EMBI spreads. Equation (4.1) in Table 4 presents the results of the cross-section fixed effects regression for our unbalanced panel data of 18 countries. Both the rt and vix variables have the expected coefficient signs and are statistically significant. Compared to the MG and CCE-MG estimations, the absolute magnitude of the estimated coefficients appears to be smaller for the rt coefficient and larger for the vix coefficient. However, the basic idea that the EMBI spreads are largely determined by global financial conditions, along with domestic fundamentals, remains to be strongly supported. For a robustness check, we also consider high yield spread (HYS) as an alternative measure of global risk appetite and thus liquidity conditions. Equation (4.2) in Table 4 presents the results with hys (log. of HYS) instead of vix. The results from (4.2) are essentially the same as those from (4.1). In Equation (4.3) we consider vix and hys jointly. Accordingly, the inclusion of hys does not affect the magnitude of the estimated coefficient for vix significantly. The coefficient of hys, on the other hand, decreases substantially with the inclusion of vix. This evidence may lend support to the view that the use of the VIX index alone may not lead to a significant information loss in our analysis.

The results of the panel unit root tests presented in Table 5 suggest that all the variables in our panel data regressions are integrated of order one (I(1)). Consequently, we need to test whether these I(1) variables are not cointegrated. To this end, we consider Engle and Granger (1987) based procedures and test whether the residuals from the static regressions are not stationary[6]. All the tests suggest the stationarity of the equation residuals and thus the cointegration of the variables. Consequently, the equations represent long-equilibrium relationships and by the Granger representation theorem there is an equilibrium correction mechanism (ECM) for the evolution of EMBI spreads.

To estimate the panel ECM (PECM) representation which allows us to assess the adjustment mechanism to a deviation from the long-run equilibrium relationship, along with the short-run dynamics, we first consider the following specification:

$$\Delta s_{it} = b_{0i} + \alpha ec_{t-1} + b_1\Delta s_{it-1} + c_1\Delta rt_{it} + c_2\Delta rt_{it-1} + d_1\Delta vix_t + d_2\Delta vix_{t-1} + u_{it} \qquad (4)$$

where ec (equilibrium correction term) are the stationary residuals from equation 4.1 in Table 4. Considering the low sample variability of rt, we set the lag length as 2 for the general Autoregressive Distributed Lag (ARDL) relationship, a re-parameterisation of which gives (4). Equation (5.1) in Table 5 presents the estimation results. Accordingly, the equilibrium correction in the long-run evolution of the EMBI spreads appears to be significant, and considering the fact that the data are daily, the adjustment towards equilibrium is relatively rapid (around six months). The short-run impact of changes in the global financial conditions as represented by the Δvix coefficient appears to be significant.

3.3. US Macroeconomic News and EMBI Spreads

Financial markets often react to US macroeconomic news as documented by a large body of literature, the bulk of which is based on advanced industrial countries. In this section, we proceed by investigating the impacts of U.S. scheduled macroeconomic announcements and surprises on EMBI spreads. To this end, we consider nine major U.S. regularly scheduled

[6] The results from other panel cointegration tests such as Pedroni (2004) were essentially the same and thus not reported to save the space.

macroeconomic announcements basically concerning real activity (non-farm payroll employment NFP, retail sales RS, capacity utilization CU), consumption (new home sales NHS), forward looking (manufacturing index MAN, consumer confidence CCONF, leading indicators LEAD) and prices (core consumer price index CPI, core producer price index PPI).

We measure expectations on U.S. macroeconomic fundamentals using the median market forecasts provided by Bloomberg. For a given macroeconomic variable M_t, the "news" or "surprise" is defined as the difference between the actual macroeconomic announcement (M^a_t) and the survey expectations (M^e_t). The units of measurement differ across variables. Therefore, following Balduzzi *et al.* (2001) and Andersen *et al.* (2003), we use standardized news for the ease of interpretation. The standardized news for M_t (M^s_t) is obtained by dividing each macroeconomic news variable (M^a_t - M^e_t) by its sample standard deviation. As the sample standard deviation is constant for each of the variables, such standardization does not affect the statistical properties of the estimators.

Table 6 lists the US macroeconomic announcements and reports the individual univariate statistical properties of the data using the ADF tests. All the forecast errors or surprises (less than expected) and five of the series (NFP, RS, LEAD, CPI, PPI), both announced (realized) and expected, appear to be zero-mean stationary. The order of integration for the expected and realized CU, NHS, MAN and CCONF is found to be unity. For these variables, the expected and realized values appear to be cointegrated with a unitary coefficient as suggested by the ADF tests for the surprises. Consistent with the rational expectations hypothesis, the stationarity of the forecast errors support the lack of a systematic bias in the surprises (Edison, 1997).

The stationarity of the news variables precludes them to be considered for a cointegration analysis. As a plausible alternative, we augment the PECM given by equation (4) with the US macroeconomic news variables to obtain:

$$\Delta s_{it} = b_{0i} + \alpha ec_{t-1} + b_1\Delta s_{it-1} + c_1\Delta rt_{it} + c_2\Delta rt_{it-1} + d_1\Delta vix_t + d_2\Delta vix_{t-1}$$

$$e_1\Delta ftr_t + e_2\Delta ftr_{t-1} + \sum f_j news_{jt} + u_{it} \qquad (5)$$

where $news_j$ is the j^{th} news. Note that, equation (5) also contains changes in the U.S. federal funds target rate (Δftr) which is found to be stationary (Table 5). We suppose that the variables acting as a proxy for global financial conditions, such as VIX and HYS, may also contain the impacts of the US monetary policy changes. The inclusion of the FED target rate, thus, maintains that changes in the US monetary policy may have an impact on EMBI spreads (Arora and Cerisola, 2001) in the short-run, apart from those already captured by VIX or HYS in the long-run specifications.

Equation (5.2) in Table 5 presents the results. As expected, an increase in the FED target rates leads to an increase in the EMBI spreads in the short-run. The results also suggest that the spreads respond to US macroeconomic news about non-farm payroll employment (NFP), retail sales (RS), new home sales (NHS), ISM manufacturing (MAN) and consumer confidence (CCONF)[7]. The negative estimated coefficients of the news variables lend support to the view that stronger-than-expected announcements for U.S. real activity lead EM country spreads to decline in the short-run. It may be plausible to expect spreads of EM economies to decline with a stronger global economy. However, such an interpretation maintains that higher than expected real releases are always good news for the strength of the US economy.

[7] The news about leading indicators (LEAD), capacity utilization (CU) and prices (CPI and PPI) are found to be jointly and individually insignificant in all the specifications reported in Table 5. The Likelihood Ratio (LR) test that these variables are jointly redundant in equation (5.1), for instance, yielded 5.26 with p=0.26.

The literature often maintains that the interpretation of given macroeconomic news as “good” or “bad” is invariant to the state of the economy. Under a positive inflation gap, during which inflation tends to be higher than the long-run or targeted inflation for instance, a higher than expected real activity may be interpreted as the economy is overheating and thus a “bad news” for monetary policy causing concerns about higher interest rates. Therefore, to consider the state of the US economy, we define the periods of positive deviations of inflation (based on seasonally adjusted core CPI series) from its Hodrick-Prescott detrended cyclical component as periods of “inflation dominance”. We define a dummy variable D taking unity when the observation belongs to the “inflation dominance” period and zero otherwise. We then interrelate the news variables with D, to obtain

$$\Delta s_{it} = b_{0i} + \alpha ec_{t-1} + b_1\Delta s_{it-1} + c_1\Delta rt_{it} + c_2\Delta rt_{it-1} + d_1\Delta vix_t + d_2\Delta vix_{t-1}$$

$$e_1\Delta ftr_t + e_2\Delta ftr_{t-1} + \sum f_j news_{jt} + \sum g_j D.news_{jt} + u_{it} \qquad (6)$$

In equation (6), the coefficient of $news_j$ (f_j) now gives the impact of the j^{th} news when there is no inflation dominance whilst (g_j) gives the change in the coefficient in the period of inflationary pressures on the economy. We expect the coefficient of D $news_j$ (g_j) to be positive as “positive news” or “stronger-than-expected macroeconomic announcements” may now mean that the economy is overheating rather than reflecting the strength of the US economy.

According to the results reported by equation (5.3) in Table 5, all positive news surprises, except CCONF, significantly decrease EMBI spreads in the absence of inflation dominance in the US economy. Positive surprises about a leading indicator variable consumer confidence CCONF, appears to be good news for the strength of the economy especially when there is an inflationary pressure. The response of EMBI spreads to positive news about the retail sales (RS), on the other hand, tends to be the same across the periods. Stronger-than-expected announcements for non-farm payroll employment (NFP), manufacturing (MAN) and new home sales (NHS) all lead to a significant decrease in the EMBI spreads during the periods of relatively lower inflation rates. Inflation dominance, however, tends to reduce this impact substantially. In the case of NHS, positive surprises can be interpreted as good news for EMBI spreads when there is no inflation dominance but turns out to be bad news otherwise. All these results suggest that investors’ response to news is not invariant to the state of the economy.

4. Concluding Remarks

“When it rains, it pours” holds true according to a recent study by Kaminsky, Reinhart and Végh (2004) investigating the impact of capital flows to EM countries. Along the same lines, according to Calvo (2002, 2005), with international financial integration, EM economies become more vulnerable to exogenous shocks coming from global capital markets which is referred to as “globalization hazard”. Consequently, both capital flows to EM economies and their sudden stops leading to financial crises during the last decade exhibit a degree of “globalization hazard” (Calvo, 2005). According to Uribe and Yue (2006), in a typical emerging market the price level and real output respond to US monetary policy shocks by more than the price level and real output in the US itself. Therefore, it may be argued that “when the U.S. sneezes, emerging markets catch a cold.” Our results, strongly suggesting that the long-run evolution of EMBI spreads crucially depends on external factors arising from the interrelated global liquidity conditions, risk appetite, crises contagion and US macroeconomic news provide a further support to the argument that real output fluctuations in EM economies have been significantly triggered by global financial conditions.

The crucial importance of exogenous global factors in the determination of interest rates that EM face in international financial markets does not necessarily relegate the importance of domestic macroeconomic fundamentals. The significance of the fundamentals in the long run evolution of the spreads simply imply that a strong macroeconomic policy stance improving domestic fundamentals would decrease the default risk and hence the cost of borrowing. The domestic fundamentals are important even when the spreads are predominantly determined by external conditions because they represent the main magnifying mechanisms through which the impacts of exogenous shocks are transmitted.

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Table 1: The Determinants of EMBI Spreads: Individual Country Estimates

Country	Constant	rt	vix	R^2	ADF(l)	M_{RT}	σ_{RT}	N
Argentina	5.513** (0.057)	-0.570** (0.003)	0.723** (0.019)	0.928	-4.68(8)$^{++}$	5.74	4.83	2246
Brazil	10.958** (0.197)	-3.266** (0.073)	0.973** (0.016)	0.872	-6.65(0)$^{++}$	9.59	0.70	2246
Bulgaria	14.238** (0.144)	-4.109** (0.038)	0.411** (0.021)	0.949	-3.45(3)$^{+}$	11.16	2.13	2022
Colombia	4.930** (0.307)	-0.867** (0.122)	1.109** (0.017)	0.721	-4.87(1)$^{++}$	11.46	0.80	1893
Ecuador	4.976** (0.055)	-0.501** (0.022)	0.968** (0.013)	0.788	-6.20(0)$^{++}$	6.07	0.79	1600
Egypt	9.778** (1.518)	-4.019** (0.617)	1.780** (0.028)	0.781	-4.97(2)$^{++}$	12.34	0.52	1143
Mexico	15.006** (0.149)	-4.504** (0.049)	0.688** (0.011)	0.934	-7.06(0)$^{++}$	12.66	0.97	2246
Malaysia	13.813** (0.500)	-3.975** (0.164)	0.570** (0.023)	0.830	-3.94(1)$^{++}$	14.89	1.15	725
Morocco	22.016** (0.778)	-8.199** (0.291)	1.151** (0.032)	0.796	-4.87(1)$^{++}$	11.16	0.42	2185
Panama	2.875** (0.181)	0.335** (0.081)	0.734** (0.012)	0.704	-4.88(2)$^{++}$	11.32	0.57	2246
Peru	6.051** (0.297)	-1.604** (0.113)	1.257** (0.018)	0.796	-6.38(1)$^{++}$	10.69	0.56	2246
Philippines	3.962** (0.225)	0.224 (0.117)	0.517** (0.026)	0.389	-2.59(0)	11.22	0.79	2102
Poland	6.223** (0.538)	-2.314** (0.188)	1.624** (0.025)	0.726	-5.46(2)$^{++}$	14.69	0.63	2246
Russia	1.014** (0.116)	-0.315** (0.007)	1.919** (0.038)	0.768	-4.38(1)$^{++}$	8.40	4.93	2246
S. Africa	11.506** (0.454)	-3.242** (0.155)	0.674** (0.022)	0.811	-4.69(2)$^{++}$	13.35	1.00	1010
Turkey	8.113** (0.154)	-1.803** (0.046)	0.636** (0.021)	0.858	-4.47(2)$^{++}$	8.55	1.21	1851
Ukraine	6.942** (0.376)	-1.689 ** (0.130)	0.853** (0.038)	0.730	-4.27(1)$^{++}$	8.83	0.90	1252
Venezuela	2.666** (0.062)	-0.134** (0.012)	1.383** (0.012)	0.695	-5.97(0)$^{++}$	7.98	1.42	2246

Notes: t-ratios in parentheses. (*) and (**) denote significance at the 5 % and 1% level, respectively. ADF(l) are the ADF tests results for the residuals of the corresponding equation with the lag length (l) chosen by the Akaike Information Criteria (AIC). For the ADF tests, ($^{+}$) and ($^{++}$) denote the rejection of the unit root null hypothesis at the 5% and 1% level, respectively. M_{RT} is the sample mean of the outlook augmented rating and σ_{RT} is its' standard deviation. N is the effective number of observations.

Table 2: The Determinants of EMBI Spreads: Individual Country Estimates with CCE

Country	Constant	rt	vix	m_s	m_rt	R^2	ADF(l)
Argentina	5.129** (0.599)	-0.525** (0.004)	-0.121** (0.026)	0.586** (0.020)	-0.372 (0.214)	0.957	-5.08(8)$^{++}$
Brazil	8.743** (0.307)	-1.151** (0.085)	0.667** (0.017)	0.326** (0.014)	-1.580** (0.116)	0.919	-5.87(0)$^{++}$
Bulgaria	17.288** (0.693)	-4.397** (0.077)	0.508** (0.024)	-0.206** (0.025)	-0.561** (0.201)	0.952	-4.63(1)$^{++}$
Colombia	-0.570 (0.399)	-1.529** (0.122)	0.333** (0.015)	0.740** (0.016)	1.980** (0.156)	0.915	-4.92(0)$^{++}$
Ecuador	1.265** (0.596)	-0.191** (0.062)	0.589** (0.017)	0.370** (0.023)	0.810** (0.208)	0.854	-6.14(0)$^{++}$
Egypt	-4.350* (2.158)	2.988** (0.770)	1.136** (0.054)	0.302** (0.054)	-1.366* (0.572)	0.814	-4.66(1)$^{++}$
Mexico	1.466** (0.389)	-2.985** (0.053)	0.476** (0.011)	0.522** (0.013)	2.949** (0.091)	0.963	-7.30(0)$^{++}$
Malaysia	5.934** (0.976)	-2.423** (0.201)	-0.121* (0.042)	1.135** (0.058)	-0.747 (0.454)	0.889	-5.65(0)$^{++}$
Morocco	1.080 (0.735)	-7.104** (0.305)	0.521** (0.025)	0.942** (0.020)	5.974** (0.178)	0.904	-5.50(1)$^{++}$
Panama	2.532** (0.168)	0.857** (0.057)	0.156** (0.010)	0.355** (0.008)	-0.621** (0.077)	0.916	-4.90(1)$^{++}$
Peru	-0.319 (0.325)	-1.974** (0.118)	0.572** (0.020)	0.633** (0.016)	2.212** (0.160)	0.894	-5.21(1)$^{++}$
Philippines	5.746** (0.430)	-2.034** (0.121)	0.275** (0.023)	0.526** (0.022)	0.438** (0.143)	0.587	-2.81(0)
Poland	-15.77** (1.408)	0.431 (0.239)	1.351** (0.036)	0.631** (0.036)	4.814** (0.299)	0.759	-5.44(2)$^{++}$
Russia	-34.27** (0.692)	-0.215** (0.004)	0.609** (0.037)	1.886** (0.032)	11.398** (0.236)	0.911	-4.52(0)$^{++}$
S. Africa	5.829** (0.249)	0.042 (0.339)	0.573** (0.022)	0.117** (0.031)	1.461** (0.245)	0.835	-4.72(2)$^{++}$
Turkey	5.899** (0.641)	-1.726** (0.051)	0.428** (0.025)	0.220** (0.022)	0.536* (0.246)	0.873	-4.20(0)$^{++}$
Ukraine	4.615** (1.440)	0.543 (0.285)	0.804** (0.037)	0.249** (0.047)	-1.655* (0.658)	0.758	-5.24(0)$^{++}$
Venezuela	5.410** (0.309)	0.034** (0.007)	0.336** (0.017)	0.639** (0.014)	-1.712* (0.109)	0.918	-4.39(0)$^{++}$

Note: See Table 1.

Table 3. Panel Unit Root Tests

Variables	MW	LLC	IPS
s	11.59 (4) [1.000]	3.38 (4) [1.000]	3.61 (4) [0.999]
Δs	3705.0^{++} (3) (0.000)	-101.1^{++} (3) (0.000)	-103.7^{++} (3) (0.000)
rt	27.05 (0) [0.859]	0.170 (0) [0.567]	-0.838 (0) [0.201]
Δrt	187.2^{++} (0) [0.000]	-194.0^{++} (0) [0.000]	-157.5^{++} (0) [0.000]
vix	31.94 (3) [0.663]		
Δvix	2442.7^{++} (2) [0.000]		
hys	31.61 (2) [0.680]		
Δhys	832.1^{++} (1) [0.000]		
fdtr	21.52 (0) [0.973]		
Δfdtr	331.6^{++} (0) [0.000]		

Notes: MW, LLC and IPS are the Maddala and Wu (1999), Levin, Li and Chu (2002) and Im, Pesaran and Shin (2003) panel unit root tests, respectively. We report the (t*) statistic of LLC and W statistic of IPS. For the global variables, the LLC and IPS tests are not considered as there is no cross-sectional variation in the data. The values in brackets [.] are the p-values and the optimum lag lengths for the tests, chosen by the AIC, are presented in parentheses. ($^{+}$) and ($^{++}$) denote the rejection of the unit root null at the 5% and 1% levels, respectively.

Table 4. The Determinants of the EMBI Spreads: Panel Data Estimations and Panel Cointegration

	Equation 4.1	Equation 4.2	Equation 4.3
constant	3.092**	6.452**	3.293**
	(0.018)	(0.089)	(0.020)
rt	-0.435**	-0.449**	-0.433**
	(0.002)	(0.003)	(0.003)
vix	1.287**		1.189**
	(0.006)		(0.007)
hys		1.306**	0.234**
		(0.011)	(0.011)
N	33751	33751	33751
R^2	0.857	0.748	0.819
LLC	-8.213^{++}	-15.17^{++}	-7.52^{++}
	[0.000]	[0.000]	[0.000]
IPS	-15.17^{++}	-3.05^{++}	-14.92^{++}
	[0.000]	[0.000]	[0.000]
MW	374.8^{++}	144.0^{++}	337.7^{++}
	[0.000]	[0.000]	[0.000]

Notes: t-ratios in parentheses. (**) denotes significance at the 1 % level. MW, LLC and IPS are the panel unit root tests (see Table 3) for the residuals of the corresponding equations. The values in brackets [.] are the p-values. ($^{+}$) and ($^{++}$) denote the rejection of the unit root null at the 1% level.

Table 5: Panel ECM Estimations and US Macroeconomic News

	(5.1)	(5.2)	(5.3)	
constant	-0.0007**	-0.0007**	-0.0007**	
	(0.0002)	(0.0002)	(0.0002)	
$\mathbf{ec_{t-1}}$	-0.0074**	-0.0073**	-0.0073**	
	(0.0006)	(0.0006)	(0.0006)	
$\mathbf{\Delta s_{t-1}}$	-0.1667**	-0.1672**	-0.1672**	
	(0.0053)	(0.0053)	(0.0053)	
$\mathbf{\Delta rt_t}$	-0.0003	-0.0005	-0.0005	
	(0.0038)	(0.0038)	(0.0038)	
$\mathbf{\Delta rt_{t-1}}$	0.0028	0.0027	0.0026	
	(0.0038)	(0.0038)	(0.0038)	
$\mathbf{\Delta vix_t}$	0.1197**	0.1185**	0.1191**	
	(0.0042)	(0.0042)	(0.0042)	
$\mathbf{\Delta vix_{t-1}}$	0.0598**	0.0601**	-0.0598**	
	(0.0043)	(0.0043)	(0.0043)	
$\mathbf{\Delta fdtr_t}$		0.0089	0.0077	
		(0.0137)	(0.0137)	
$\mathbf{\Delta fdtr_{t-1}}$		0.0307**	0.0289**	
		(0.0137)	(0.0136)	
		News	**News**	**D*News**
NFP		-0.2373**	-0.4330**	0.4967**
		(0.1023)	(0.1296)	(0.2103)
MAN		-1.4045**	-2.1722**	1.9459**
		(0.1026)	(0.1315)	(0.2098)
RS		-0.3507**	-0.3584**	0.0191
		(0.0999)	(0.1303)	(0.1303)
NHS		-0.1865**	-0.5208**	0.8879**
		(0.0978)	(0.1212)	(0.2056)
CCONF		-0.1912**	0.1593	-0.6371**
		(0.1024)	(0.1454)	(0.2048)
N	33713	33644		33644
R^2	0.055	0.061		0.064
DW	2.02	2.01		2.01
F	85.7	72.9		66.1

Notes: 1. See Table 1.
2. In equation (5.3), the estimates in the last column are the estimated coefficients of the corresponding news variables multiplied by the inflation dominance dummy variable D.
3. The coefficients of the news variables and their standard errors are multiplied by 100 for the ease of interpretation.

Table 6. US Macroeconomic Announcements

Announcement	**N**	**Source**	**ADF Tests**		
			Realized	Expected	Surprise
	Real Activity				
Nonfarm Payroll Employment (NFP)	108	BLS	-5.65**	-3.53**	-8.82**
Retail Sales (RS)	90	BC	-13.54**	-8.88**	-14.93**
Capacity Utilization (CU)	108	FRB	-1.29 [-8.79**]	-1.42 [-7.07**]	-10.75**
	Consumption				
New Home Sales (NHS)	108	BC	-2.98 [-13.94**]	-2.12 [-11.30**]	-11.29**
	Forward Looking				
ISM Manufacturing Index (MAN)	104	ISM	-2.33 [-9.29**]	-1.98 [-9.46**]	-9.69**
Consumer Confidence Index (CCONF)	108	BC	-1.71 [-10.20**]	-1.82 [-8.12**]	-10.20**
Index of Leading Indicators (LEAD)	98	BC	-9.90**	-8.26**	-9.85**
	Prices **(core)**				
Consumer Price Index (CPI)	104	BLS	-9.79**	-4.01**	-9.41**
Producer Price Index (PPI)	104	BLS	-15.02**	-7.40**	-15.49 **

Notes and abbreviations:
Number of observations (N), Bureau of Labor Statistics (BLS), Bureau of the Census (BC), Federal Reserve Board (FRB), The Institute for Supply Management (ISM). ** Denotes the rejection of the unit root null at the 1% level. The values in brackets [.] are the Dickey-Fuller (DF) test results for the first difference of the corresponding variable. The optimum lag for the DF regression equations with no constant is found to be zero for all the variables by the Akaike Information Criteria. The unit root test results are found to be robust to both an inclusion of a constant term or a higher lag length.

APPENDIX

Table A1. ADF Tests for the Country-Specific Variables

Country	s_t	Δs_t	r_t	Δr_t
Argentina	-1.13(13)	-10.3(12)**	-1.22(0)	-47.3(0)**
Brazil	-0.59(1)	-41.0(0)**	-0.63(0)	-47.4(0)**
Bulgaria	-0.15(3)	-33.1(0)**	-0.26(0)	-45.1(0)**
Colombia	-0.48(0)	-45.0(0)**	-1.91(0)	-47.4(0)**
Ecuador	-1.58(1)	-43.8 (1)**	-1.08(0)	-47.4(0)**
Egypt	-1.59(2)	-29.8(0)**	-0.63(0)	-47.4(0)**
Mexico	-0.71(1)	-43.4(0)**	1.53(0)	-47.3(0)**
Malaysia	-0.83(2)	-27.4(1)**	-0.22(0)	-47.3(0)**
Morocco	-0.47(3)	-37.7(2)**	-1.06(0)	-46.9(0)**
Panama	-0.73(0)	-47.9(0)**	-0.80(0)	-47.4(0)**
Peru	-0.43(1)	-33.1(0)**	-0.26(0)	-45.1(0)**
Philliphines	-0.59(1)	-53.7(0)**	0.58(0)	-47.3(0)**
Poland	-1.06(3)	-39.1(2)**	-2.03(0)	-47.4(0)**
Russia	-0.23(0)	-46.9(0)**	-1.42(0)	-47.3(0)**
S. Africa	-2.15(0)	-30.7(0)**	-0.33(0)	-47.4(0)**
Turkey	-0.62(0)	-41.5(0)**	-0.93(0)	-47.3(0)**
Ukraine	-2.06(1)	-47.5(0)**	-0.50(0)	-35.4(0)**
Venezuela	-0.37(1)	-43.9(0)**	-0.67(0)	-47.3(0)**

Notes: The values in parentheses are the optimum lag length for the ADF regressions chosen by the AIC. (**) denotes the rejection of the unit root null hypothesis at the 1 % level

Table A2. The Determinants of EMBI Spreads: Country Estimates with hys

Country	constant	rt	hys	R^2	ADF(l)
Argentina	7.410**	-0.569**	0.705**	0.904	-4.90(0) $^{++}$
	(0.014)	(0.004)	(0.031)		
Brazil	18.103**	-5.265**	0.747**	0.749	-3.50(0) $^{+}$
	(0.194)	(0.084)	(0.025)		
Bulgaria	14.699**	-3.948**	0.930**	0.967	-4.37(2) $^{++}$
	(0.070)	(0.026)	(0.023)		
Colombia	5.674**	-0.048	1.389**	0.477	-1.98(0)
	(0.428)	(0.176)	(0.036)		
Ecuador	6.334**	0.018	1.292**	0.607	-3.06(0)
	(0.175)	(0.096)	(0.026)		
Egypt	42.110**	-15.519**	4.119**	0.693	-3.43(1) $^{+}$
	(1.962)	(0.797)	(0.081)		
Mexico	20.887**	-6.097**	0.571**	0.888	-4.26(0) $^{++}$
	(0.128)	(0.050)	(0.016)		
Malaysia	13.813**	-3.975**	0.570**	0.830	-3.94(1) $^{++}$
	(0.500)	(0.164)	(0.023)		
Morocco	39.739**	-14.211**	0.574**	0.702	-3.53(0) $^{+}$
	(0.657)	(0.269)	(0.042)		
Panama	0.950**	1.946**	0.572**	0.398	-1.65(0)
	(0.253)	(0.105)	(0.022)		
Peru	14.869**	-3.885**	0.911**	0.464	-2.48(0)
	(0.422)	(0.175)	(0.038)		
Phillliphines	2.232**	1.527	0.344**	0.331	-0.59(0)
	(0.199)	(0.084)	(0.025)		
Poland	31.678**	-10.291**	2.223**	0.744	-4.68(1) $^{++}$
	(0.466)	(0.175)	(0.032)		
Russia	6.539**	-0.393**	0.853**	0.525	-1.70(1)
	(0.039)	(0.010)	(0.079)		
S. Africa	19.970**	-5.819**	0.951**	0.763	-4.15(1) $^{++}$
	(0.349)	(0.130)	(0.040)		
Turkey	10.664**	-2.257**	0.727**	0.857	-3.61(0) $^{+}$
	(0.083)	(0.036)	(0.025)		
Ukraine	10.474**	-2.409**	1.468**	0.796	-5.21(0) $^{++}$
	(0.189)	(0.083)	(0.045)		
Venezuela	6.278**	-0.120**	1.329**	0.329	-1.41(0)
	(0.040)	(0.019)	(0.040)		

Note: See Table 1.

Table A3: The Determinants of EMBI Spreads: CCE Country Estimates with hys

Country	constant	rt	cbs	M(embi)	M(rt)	R^2	ADF(l)
Argentina	3.879**	-0.536**	0.068*	0.544**	0.108	0.957	$-5.07(8)^{++}$
	(0.753)	(0.005)	(0.031)	(0.010)	(0.276)		
Brazil	5.648**	-2.011**	0.400**	0.629**	0.509**	0.878	$-3.35(2)^{+}$
	(0.378)	(0.110)	(0.021)	(0.013)	(0.155)		
Bulgaria	12.118**	-3.766**	0.961**	0.105**	0.618**	0.967	$-4.79(1)^{++}$
	(0.597)	(0.045)	(0.024)	(0.021)	(0.171)		
Colombia	0.711	-0.954**	0.432**	0.741**	1.191**	0.919	$-4.48(0)^{++}$
	(0.394)	(0.083)	(0.018)	(0.016)	(0.158)		
Ecuador	3.557**	-0.394**	0.714**	0.455**	0.379*	0.871	$-4.15(0)^{++}$
	(0.565)	(0.058)	(0.018)	(0.021)	(0.196)		
Egypt	4.085*	-1.120	2.357**	0.601**	-0.333	0.849	$-5.63(0)^{++}$
	(2.158)	(0.735)	(0.082)	(0.049)	(0.515)		
Mexico	0.624	-3.383**	0.389**	0.693**	3.807**	0.953	$-5.36(0)^{++}$
	(0.439)	(0.039)	(0.013)	(0.013)	(0.105)		
Malaysia	4.605**	-1.126**	0.471*	0.811**	-1.012	0.895	$-5.98(0)^{++}$
	(0.961)	(0.279)	(0.066)	(0.044)	(0.420)		
Morocco	-0.423	-8.385**	0.668**	1.126**	7.970**	0.918	$-5.77(1)^{++}$
	(0.683)	(0.266)	(0.023)	(0.017)	(0.161)		
Panama	1.511**	0.843**	0.109**	0.437**	-0.216*	0.911	$-4.47(1)^{++}$
	(0.187)	(0.065)	(0.011)	(0.009)	(0.097)		
Peru	-4.394	-3.126**	0.501**	0.945**	4.884**	0.891	$-4.53(1)^{++}$
	(0.339)	(0.110)	(0.019)	(0.011)	(0.153)		
Philliphines	2.766**	-2.011**	0.338**	0.674**	1.567**	0.606	-2.50(0)
	(0.412)	(0.114)	(0.022)	(0.021)	(0.151)		
Poland	2.951*	-5.577**	1.757**	0.726**	4.892**	0.793	$-4.87(1)^{++}$
	(1.396)	(0.258)	(0.040)	(0.032)	(0.277)		
Russia	-39.55**	-0.193**	0.490**	2.250**	13.321**	0.906	$-4.13(0)^{++}$
	(0.742)	(0.005)	(0.040)	(0.045)	(0.250)		
S. Africa	6.087**	0.334	0.898**	0.262**	-1.694**	0.830	$-4.68(1)^{++}$
	(1.266)	(0.345)	(0.036)	(0.033)	(0.246)		
Turkey	7.096**	-1.798**	0.586**	0.271**	0.392*	0.892	$-3.92(0)^{++}$
	(0.544)	(0.043)	(0.023)	(0.018)	(0.204)		
Ukraine	1.538	0.647**	1.606**	0.602**	0.681	0.868	$-4.82(1)^{++}$
	(1.075)	(0.203)	(0.037)	(0.034)	(0.481)		
Venezuela	3.226**	0.038**	0.241**	0.811**	-0.871	0.911	$-3.80(8)^{++}$
	(0.340)	(0.007)	(0.019)	(0.010)	(0.123)		

Note: See Table 1.