# Sovereign bond ratings and market spreads. a dynamic panel analysis.

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# SOVEREIGN BOND RATINGS AND MARKET SPREADS: A DYNAMIC PANEL ANALYSIS

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Abstract

This paper applies a measure of country risk to determine the evolution of credit spreads on secondary market sovereign bonds issued by emerging countries. After the Mexican financial crisis in 1995, this market has been characterised by a sharp decline of spreads which, by mid-1997, brought them to a level which was thought not to adequately cover risk. The episode has been followed in successive years by a new increase of spreads, accompanied by high volatility in concomitance with the Asian and Russian crises. In order to tackle the issue of how spreads are determined, we concentrate on sovereign risk as measured by spreads on Brady bonds and specify a dynamic panel model including seven countries that are large issuers of these instruments. The analysis reveals a significant effect for economic fundamentals, but we also found that spreads are significantly affected by shock factors: besides general financial crises, we isolated a role for commodity prices. We found an asymmetric effect for core countries interest rates, which signals the limited role for core rates in affecting the decline in spreads, that we instead attribute, besides a bettering of fundamentals, to a spreading of globalisation. In the post '97 period we found spreads grossly in line with fundamentals but we have no specific explanation to offer for the occurrence of repeated financial crises save that a general recourse to the argument of interdependence. We think that the analysis of contagion or interdependence problems that has recently attracted much attention obviously deserves further work and possibly a different econometric technique using data at a higher frequency than the monthly data employed in this study.

JEL Classification: C23, F34, G12, G15

Keywords: Brady bonds, bond spreads, sovereign ratings, emerging markets

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# SOVEREIGN BOND RATINGS AND MARKET SPREADS: A DYNAMIC PANEL ANALYSIS

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## 1. THE DETERMINANTS OF EMERGING MARKET SPREADS

In recent years, much attention has been given to the issue of how spreads on emerging market bonds are determined: the question arises from the observed turmoil in emerging countries during the second half of the nineties. After the explosion of spreads in the course of the Mexican financial crisis in 1994-95, a systematic and sharp reduction of spreads took place during the following years until, by the half of 1997, they had plunged at levels which were the same or lower than those observed before the crisis. This decline was so sustained that it induced concern among investors because it was felt that spreads had perhaps fallen too far. The following episodes, starting with the Asian crisis in 1997, passing through the Russian in the summer of 1998, until the recent Brazilian troubles at the beginning of 1999, seem to confirm this suspicion. It is difficult, anyway, to determine what the equilibrium level of the price of emerging market bonds should be. The question so becomes twofold: on the one side it implies the need to investigate on the economic determinants that affect the price - and hence the spread - of emerging country bonds; on the other hand it gives rise to the issue of which factors forced the spreads to such low levels in those years.
With regard to the first problem, there are several works focusing on it which began to be published since the beginning of the eighties, and of which we will present a review. These papers focus on the relation between the price of emerging market debt and economic fundamentals, with the aim of finding specific variables able to predict spreads in lending to less developed countries. Besides the endogenous variables representing macro fundamentals of the economies considered (among these a significant role is usually found, as an example, for per-capita income, inflation rate and external or fiscal unbalances), exogenous factors of shock were also considered when it was thought that they might have an impact on the degree of solvency of the

country borrower, as for example the interest rates in core economies, or the price of oil.

Recently, the role of industrialised country interest rates has begun to be examined with renewed attention; this brings back to the second aspect of the question: the pronounced narrowing of the spreads in 1995.

Among the variety of explanations that has been proposed, the basic claim reported from financial observers was that core country interest rates affect the "appetite for risk" of investors. The idea is that an expansive monetary stance, hence a high level of global liquidity, may have increased the appetite for riskier investments in those years, forcing spreads to levels lower to those that adequately cover risk.

Many other interpretations were actually suggested: a first one is the possibility of a resumption of a longer-term trend toward low levels of emerging credit spreads which was interrupted by the Tequila crisis in 1995: this trend could reflect, again on the lenders' side, the maturing of core financial markets, with improved access to information, liberalisation of regulatory frameworks regarding specific investor classes (in particular pension funds), improved risk management capabilities. This general trend of globalisation would have fostered the supply of lending to emerging markets, which, on the other hand, were implementing in those years important programs of stabilisation and structural reform.

In contrast with this market-efficiency claim, other argued the view of possible herding behaviours which may result as a consequence of the globalisation process, mainly due to information asymmetries among investors[1]. Actually, another effect of globalisation can consist in a reduction of the incentives to incur the costs of acquiring country-specific information in the asset allocation activity: financial contagion becomes the "rational" result of an optimal portfolio allocation when securities markets grow and the costs of acquiring information are "sufficiently" great.

Another hypothesis is related to the claim of a moral hazard problem in lending to emerging economies, which would be induced by the IMF-led rescue package after the Tequila crisis which, through the bailing out of the Mexican government, would have strengthened the incentives for investing in risky markets.

[1] This issue is tackled for example in Calvo [1]; uninformed investors may incorrectly extract signals from the informed investor set: in case of negative signalling, they may not discern if this is due to negative information about markets or because of margin calls of leveraged investors. In this sense, see also Fernández-Arias [10] for a brief discussion on the role of the hedge funds in the Russian crisis.

What is generally shared among all the views previously expressed, is the limited role, from the borrowers' side, for the evolution in economic fundamentals, which seem to be overshadowed by the financial events occurred in past years. The volatility of spreads that followed the Russian crisis, characterised by temporary high peaks followed by drastic declines, has further highlighted the role of financial contagion or interdependence and possibly the diminished weight of fundamentals in the market assessment of country risk.

In the following we analyse data on a set of emerging countries during the years from 1992 to 1999. We solve the problem of measuring the role of fundamentals by developing a rating index that is constructed, as we explain in the text, as to give a simplified replication of Moody's ratings.

Section 2.1 gives a brief survey of the structure of the emerging markets of bonds and presents the EMBI Index, which refers to the specific class of bonds that we will model in this analysis. Section 2.2 reviews the literature. Section 3.1 reports, after a preliminary empirical examination of the time series properties of the aggregate EMBI spreads, the estimates of a regression equation on the aggregate EMBI time series. Section 3.2 details the procedure and the macroeconomic variables used to construct our rating index, and reports the estimates of a disaggregated dynamic panel model for seven countries that are large issuers of Brady bonds. Section 4 gives some conclusions and suggestions for further work.

## 2.1 THE EMERGING BOND MARKETS

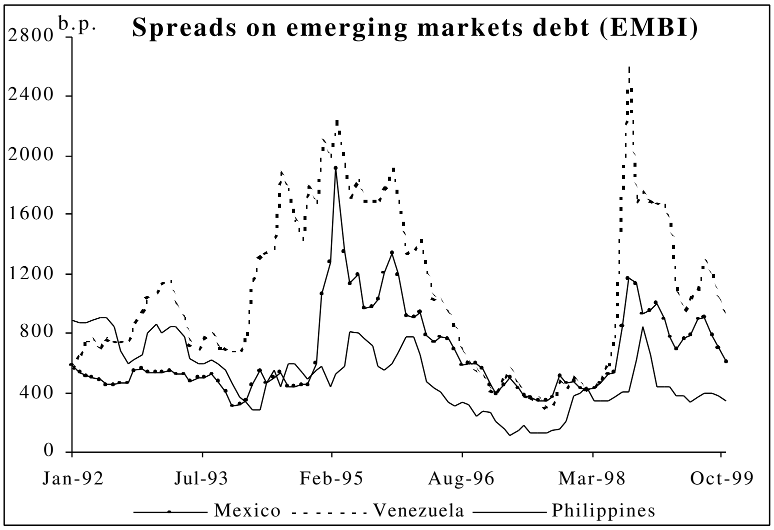
The analysis in this paper will be restricted to debt instruments exchanged on the secondary markets of bonds. The credit spread is defined as the excess yield which a risky security pays over an industrial country government security[2] of the same currency denomination and maturity: it represents the assessment of the market on the degree of riskiness associated to the country issuer. We examine secondary market spreads[3] associated to sovereign issues, not focusing on private issues or other forms of

[2] For floating coupon bonds, the spread is measured over Libor.

[3] Secondary market spreads follow different patterns than launch spreads over time: factors that increase the perception of risk associated to emerging markets debt, while negatively affecting secondary market spreads, may have opposite effects on first issues because only the low-risk borrowers are left to tap the market, while riskier borrowers are rationed out. See Eichengreen-Mody [8]and Min [18] for an analysis on launch spreads; Cline-Barnes [5] and Zhang [22] consider secondary market spreads, the first on a

lending such as syndicated loans[4]; we consider the spreads associated to the J.P. Morgan Emerging Markets Bond Index (EMBI), which is a total return index of U.S. dollar-denominated Brady bonds and other sovereign restructured bonds.

Chart 1




The EMBI Index is composed of a set of country sub-indexes whose spreads are a weighted average of those of the instruments included in each country basket; the weights are given by the capitalisation of the instruments considered.

Brady securities follow the reconstruction of commercial bank loans after the LDCs debt crisis during the eighties. The first Brady plan, accredited to former U.S. Treasury Secretary, Nicholas Brady, was launched in Mexico in 1989: the Brady bond market has then quickly become the largest and most liquid of the emerging fixed-income markets[5]. These bonds are usually collateralised by Treasury zero-coupons, hence the risk they carry blends pure country risk and the zero credit risk of the assets used as guarantee: sovereign spreads are computed stripping the influence of collaterals to capture the residual country risk.

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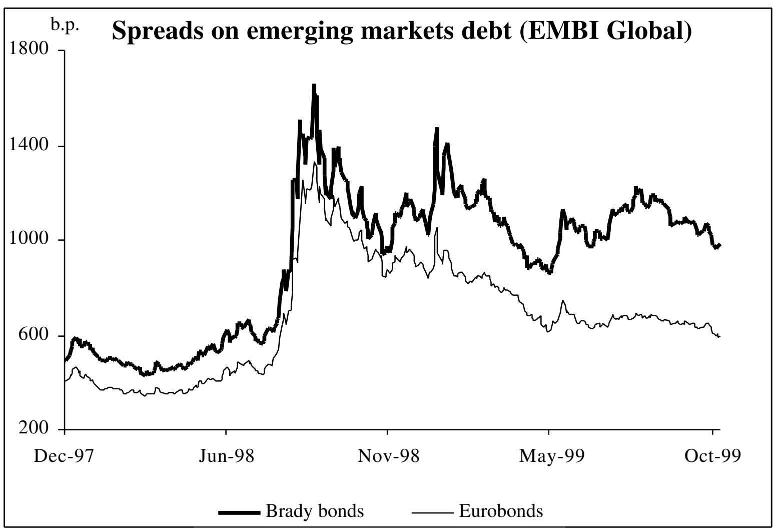
pooled sample of emerging markets and industrial countries; Kamin-Kleist [15] conduct a study on both the kind of issues.

[4] See Kamin-Kleist [15] or Cline-Barnes [5] for an exhaustive description of the characteristics of these markets and the recent evolution of the associated spreads. Eichengreen-Mody [8] examine the differences in the issuing decision of borrowers and pricing decision of lenders through a sample selection model à la Heckman for private and non-private borrowers. Durbin-Ng [6] study the correlation between secondary sovereign spreads and secondary corporate spreads to test the "sovereign ceiling" rule that no corporate debt should be better rated than that of the home government.

[5] For a more detailed historical perspective on Brady bonds and a description of their basic features see EMTA [9].

To date Brady bonds are not, anyway, fully representative of the market: in the first place their quote on the total secondary market has been reducing since the beginning of the '90s; on the other hand the countries issuing bonds in more recent years are generally better rated than the countries that issued Brady bonds and, as a consequence of being perceived as riskier instruments, the evolution of the related spreads can be different from that of other international bond classes. Differences in the spreads' dynamics can also be attributed to the fact that they generally have longer tenors than Eurobonds or Global bonds; in addition, the collateralisation may be not well perceived by investors for the transaction costs that it induces: for all these reasons the EMBI spread cannot be considered as fully representative of the cost of financing of the whole of emerging countries[6].

Chart 2




The EMBI Global, published by J.P. Morgan since 1999, is based on a broader country set than that of the EMBI; it adds to Brady bonds other classes of instruments such as Eurobonds, traded loans and local market securities.

The selection of the EMBI Index is motivated by the fact that its spread is the only readily available synthetic measure of spread for each country – associated to the average maturity of each country/sub-index – over a sufficiently long span of years. Tracking individual bond spreads implies the need of verifying that they satisfy

[6]See Kamin-Kleist [15], page 4-6. As countries go through the process of structural reform and the perception of their sovereign credit improves, Brady bonds are expected to eventually be replaced by debt which is cheaper to fund, or repurchased.

conditions of sufficient liquidity, so as to obtain fair prices of the bonds considered; the tendency of spreads to decline as bonds approach maturity may represent another source of bias toward measuring spreads over time. Furthermore, and to conclude, Brady bond spreads represent anyway an homogeneous class — the low-credit rating one — of emerging market bonds.

## 2.2 PAST STUDIES

Interest in the pricing of emerging country bonds, as opposed to bank loans, was revived by Edwards [7] in the aftermath of the international debt crisis erupted in 1982. Major concern of this paper was the difference in behaviour between bonds and loans, based on the conjecture that bond spreads may reflect risk more accurately than loans, due to the syndicated nature of the latter as opposed to dispersed nature of holders of the former[7]. In the above paper Edwards developed a model that can be regarded as the model conventionally followed by many authors for the determination of the spreads in the case of risk-neutral lender. Let's write with Edwards

$$(1-p)\left[1+\left(i^{*}+s\right)\right] = \left(1+i^{*}\right),$$

where $i^*$ is the international risk-free rate, $s$ the spread and $p$ the probability of default, possibly a linear function of some macroeconomic indicators $x_i$; we can then derive the regression equation

$$\log(s) = \log\left(1+i^{*}\right) + \sum \beta_i x_i + u,$$

that has been used by Edwards and many others as a basis for the specification of models for spreads. The main finding of his analysis was to indicate in the debt-output ratio and in the investment-GNP ratio the two major determinant of the spreads, the former negative and the latter positive. On the other hand Edwards did not find significant differences in the determinants of both loans and bonds.

[7] In the case of crisis, bondholders would have no other choice than declare default, while banks should usually be able to reschedule the payments. Without entering deeply into the problem, Miller-Zhang [17] show how the lack of coordination among bondholders, not covered by Paris or London clubs, may induce moral hazard both on the lenders' side, in case of IMF bailing out of borrowers in trouble, and on the debtors' side, in case of authorized payment standstill. The two events have potential heavy repercussions of opposite sign on the price of bonds.

The basic approach set forth by Edwards has been followed in the '90s by a number of authors that we briefly examine.

Cantor and Packer [4] examined the relationship between the sovereign credit ratings assigned by Moody's and Standard & Poor's [8] and the secondary market spreads on sovereign bonds, issued by both developed and developing countries, through a cross-section regression analysis. In a first equation they put a set of macroeconomic variables as regressors for the spreads, finding an important role for the per capita income, inflation and the external debt-exports ratio; when the credit rating variable is added to the equation none of the economic variables remains statistically significant: the authors conclude stressing the ability of credit ratings to provide the market with information about country risk which goes beyond that available in public macro data[9].

Cline and Barnes[5] used a pool of 11 emerging market bonds and 6 industrial country bonds, over the period 1992-96, to estimate a model of secondary market spreads based on the fundamentals of the countries considered: they pointed out strong significance for the external debt-exports ratio, followed by the international reserves-imports ratio and the inflation rate. The sharp decline of emerging market spreads during the period 1995-97 exceeds what could be explained by the macroeconomic performance of the emerging markets in question: the authors ascribe the phenomenon to the supply-side forces which have arisen in recent years in the international capital markets.

Eichengreen and Mody [8] conducted a study on launch spreads of about a thousand developing-country bonds over the period 1991-96: they paid special attention to the problem of sample selection bias associated to the borrowers' issuing decision; they considered both the issuing decision of debtors and the pricing decision of the investors in an à la Heckman model where they are treated simultaneously. They also tested the impact on the behaviour of the spreads over time of shifts in market sentiment toward emerging markets: they found that the variation in spreads in 1995-97 seems to be

[8] See Cantor-Packer [3] for a survey of the historical and the institutional aspects of the ratings industry.
[9] Other papers which focus on the interaction between sovereign credit ratings and market spreads are the more recent Reisen and von Maltzan [21] and Larrain, Reisen and von Maltzan [16]: through a Granger-causality test they discuss whether the rating of the agencies lead or lag markets with respect to sovereign risk. Their findings lead to a cautious conclusion about the existence of an independent long-run impact of credit ratings on the markets assessment of country risk: anyway they find a significant announcement effect in case of a put on review (in particular if the outlook is negative) by the rating agencies. This makes them suggest a favourable policy conclusion about the ability of the rating industry in dampening excessive capital flows into the emerging markets.

dominated more by arbitrary shifts in pricing behaviour[10], originated by good liquidity conditions in core markets, than by a better performance of the systems considered during that period.

Min [18] examines the primary market of a sample of about 500 issues of Asian and Latin American countries from 1991 to 1995: inflation rate, debt-GDP ratio and reserves-GDP ratio result as the most relevant variables for predicting the evolution of the spreads. Special attention is given to the analysis of external shocks as measured by the real oil price and the international interest rates, as proxied by the 3-month US Treasury Bill rate: both variables were found to be scarcely significant.

Kamin and Kleist [15] estimate a panel data model, over the period 1991-97, on about a thousand launch spreads on new bond issues and new bank loans: by exploiting the evidence of Cantor and Packer[4]'s analysis, they use the ratings assigned by Moody's and Standard & Poor's to new loan and bond issues as measures of country risk in the place of fundamentals. They test the significance of industrial country rates in the determination of the spreads, considering both the primary and the secondary market: the result points to a little role for industrial rates of United States, Germany and Japan in explaining the spreads' dynamics. The authors qualify the phenomenon of falling spreads after 1995 more as a consequence of the increased financial globalisation over the course of the beginning of the nineties, trend halted by the Mexican crisis but taken up after its dissipation.

Zhang [22]'s work, based on a panel of eight Eurobonds and Brady bonds of developing countries from 1992 to 1997, discusses the issue of a possible moral hazard in lending to emerging markets during the nineties, induced by the large official support program in Mexico during 1995, which would have taken to overlending to emerging markets in the successive years and to the observed pronounced decline in spreads after that crisis. Zhang tests the moral hazard hypothesis inserting the spreads on B-rated US corporates among the usual economic explanatory variables: the evidence of a positive correlation between the two variables and the similar declining

[10] Calvo and Mendoza [2] explain why, in globalised securities markets, investors can be incentivated to imitate arbitrary "market" portfolios: the herding behaviour can "rationally" be the result of optimal portfolio diversification when securities markets grow in the presence of information costs.

pattern during 1996-97 showed by corporate spreads would not support the moral hazard claim[11].

## 3. EMPIRICAL ANALYSIS

In this paragraph we will examine empirically the relationship between spreads, fundamentals and international short term rates as suggested by the existing literature.

### 3.1 A PRELIMINARY ANALYSIS OF THE TIME SERIES PROPERTIES OF THE EMBI SPREADS

The behaviour of the aggregate EMBI spreads is highly erratic over the entire sample considered in this analysis (1992.01 1999.10). Their higher volatility with respect to most other measures of credit spreads (loans and bonds) that have been produced (see Kamin and von Kleist [15]) is explained by the fact that they represent the low-rating class of emerging market bonds.

As in previous studies we chose to model the logarithm of the spread over the risk-free rate and on a measure of country creditworthiness plus dummies to reflect special events. Before presenting such regression analysis at the disaggregated level we wish to examine some properties of the aggregate time series.

Formal unit root tests indicate that the spread appears in the sample as an I(1) variable. ADF with a constant and six lags gives:

| | | | |
|---|---|---|---|
| ADF Test Statistic | -1.669 | 1% Critical Value* | -3.507 |
| | | 5% Critical Value | -2.895 |
| | | 10% Critical Value | -2.584 |

and similar results can be obtained with the Phillips-Perron test.

We decided not to take at face value the above sample information but rather to interpret it as the need to include in the empirical analysis appropriate dummy variables to reflect the presence of aberrant observations[12]. The spreads are in fact obviously affected by the numerous financial crises that have troubled the '90s, starting with the Mexican crisis of 1995. We treat the effect of crises as innovation outliers (see Box 1).

[11] See Miller-Zhang [17] for a more detailed discussion of the issue.

BOX 1

A simulation experiment to assess the impact of IO on DF statistic

Neglected innovation outliers will not generally have large effects on the estimated coefficients of an ARMA process (see Franses [12]); they have however some effect in making it difficult for near unit root processes to be properly assessed. On the other hand it is possible to show that taking care of the i.o. with an appropriate dummy produces estimates that converge to the true value as the size of the i.o. increases. At the same time the ADF statistic (including the dummy) will uniformly increase. To show such an effect we report in the table below the result of a simulated experiment.
The generating process is:

$$y_t=0.9*y_t(-1)+u_t+\omega*io_t\ (\tau=50),\ t=2,100,\ u_t=N(0,1),\ y_1=0,$$
$$\omega \text{ takes values } 1,2,3,...,10.$$

| Dependent var. is $\Delta(y_t)$ | Test equation is $\Delta(y_t)=c_1+c_2\times y_t(-1)+c_3\times\Delta(y_t(-1))+u_t+\omega\times io_t\ (\tau=50)$ Average of 20 replications | | |
|---|---|---|---|
| $\omega$ | $c_2$ | se_$c_2$ | t_$c_2$ |
| 1 | 0.86 | 0.06 | -2.59 |
| 2 | 0.86 | 0.06 | -2.60 |
| 3 | 0.86 | 0.07 | -2.64 |
| 4 | 0.86 | 0.07 | -2.68 |
| 5 | 0.87 | 0.07 | -2.74 |
| 6 | 0.87 | 0.07 | -2.81 |
| 7 | 0.87 | 0.07 | -2.90 |
| 8 | 0.88 | 0.07 | -2.99 |
| 9 | 0.88 | 0.06 | -3.09 |
| 10 | 0.88 | 0.06 | -3.20 |

The table is self illuminating. Note that for true unit root processes the ADF will be little affected and the AR parameter will go to one.

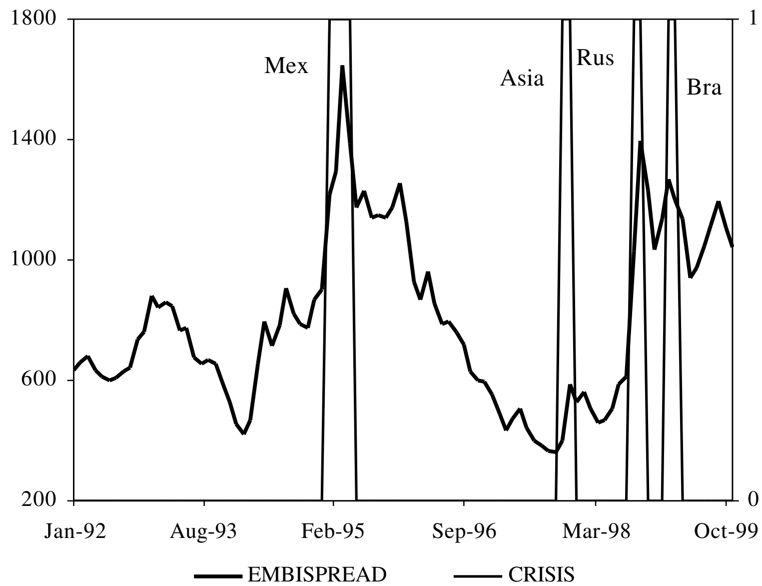
The interpretation of crises as innovation outliers can be appraised by inspection of Chart 3, in which we plot the behaviour of the spreads together with the timing of the four major crises that have occurred in the nineties. If we insert a dummy variable that takes values of 1 at the dates indicated by the above graph, the simple DF statistics rises to -2.76, which is indicative of the size of the effects that particular observations have on the test. Coherently with the results of the Montecarlo experiment reported in Box 1 we have retained the above dummy variable named CRISIS in all following analyses.

The majority of the studies previously presented, based on a shorter sample than ours, have concentrated on the explanation of the fall of spreads following the Mexican crisis. The basic explanation put forward by market observers was to be looked in the rather low credit rates prevailing in advanced economies. After looking for the evidence that short term interest rates in industrial country affect spreads on emerging market bonds (Brady bonds included), a quite general result of the above studies was

[12] The Jarque-Bera statistics for the first difference of the logs takes the value 47.95 with SK=1.3 and K=5.44 thus confirming the presence of aberrant observations.

the finding of a little role for them. On the contrary, by looking at a longer period that includes another rise in the post-98 period, we find significant effects for the US short term rate, as measured by the 3-month Fed Funds rate.

Chart 3




The following equations simply illustrate the result for the aggregate EMBI. In the first equation we regress the first difference of log spread on the US short term rate (FED3M) plus dummies and a measure of oil prices changes (BRENT).

Table 1

Equation 1: Emerging market spread and 3-month Fed Funds rate

| Dependent Variable: DLOG(EMBISPREAD) | | | | |
|---|---|---|---|---|
| Method: Least Squares | | | | |
| Sample(adjusted): 1992:03 1999:09 | | | | |
| Included observations: 91 after adjusting endpoints | | | | |
| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
| C | -0.020 | 0.012 | -1.646 | 0.103 |
| DLOG(1+FED3M/100) | 13.332 | 6.037 | 2.208 | 0.029 |
| CRISIS | 0.215 | 0.037 | 5.767 | 0.000 |
| DLOG(BRENT(-1)) | -0.253 | 0.129 | -1.960 | 0.053 |
| R-squared | 0.297 | Mean dependent var | | 0.006 |
| Adjusted R-squared | 0.273 | S.D. dependent var | | 0.129 |
| S.E. of regression | 0.110 | Akaike info criterion | | -1.539 |
| Sum squared resid | 1.048 | Schwarz criterion | | -1.429 |
| Log likelihood | 74.01 | F-statistic | | 12.28 |
| Durbin-Watson stat | 1.624 | Prob(F-statistic) | | 0.000 |

The short term rate effect is clearly significant even when controlling for financial crises and oil prices hikes. More interestingly, in the second equation we show that there is a significant asymmetric effect in the role of the US short term rate variable. In fact the coefficient for a rate increase is much more significant than for the simple rate, while the coefficient for a rate decrease is almost insignificant, even if it retains the right sign.

Equation 2: Emerging market spread and 3-month Fed Funds rate asymmetric effect.

Dependent Variable: DLOG(EMBISPREAD)
Method: Least Squares
Sample(adjusted): 1992:03 1999:09
Included observations: 91 after adjusting endpoints

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| C | -0.040 | 0.017 | -2.404 | 0.018 |
| DLOG(1+FED3M/100)*(D(FED3M)>0) | 24.819 | 8.903 | 2.788 | 0.006 |
| DLOG(1+FED3M/100)*(D(FED3M)<0) | -5.352 | 12.292 | -0.435 | 0.664 |
| CRISIS | 0.217 | 0.037 | 5.880 | 0.000 |
| DLOG(BRENT(-1)) | -0.249 | 0.128 | -1.951 | 0.054 |
| R-squared | 0.321 | Mean dependent var | | 0.006 |
| Adjusted R-squared | 0.289 | S.D. dependent var | | 0.129 |
| S.E. of regression | 0.108 | Akaike info criterion | | -1.551 |
| Sum squared resid | 1.012 | Schwarz criterion | | -1.413 |
| Log likelihood | 75.58 | F-statistic | | 10.18 |
| Durbin-Watson stat | 1.696 | Prob(F-statistic) | | 0.000 |

The result is important in the present setting, since most analysts expect a rise in short term instruments in the US and Europe: according to the above result, such interest rate hike ought to be transmitted also to emerging market debt instruments. We have allowed for this asymmetric effect also in the disaggregated analysis and found it equally significant.

### 3.2 DISAGGREGATED ANALYSIS: BRADY BONDS SPREADS FOR THE MAJOR 7 ISSUERS.

The aggregated analysis of the previous paragraph does not allow to insert measures of creditworthiness in the equation since creditworthiness is an idiosyncratic phenomenon. Save for contagion effect, the credit worth or default probability must be appreciated at the level of a single country. From the aggregated analysis we have nevertheless a number of suggestions that we want to test at the individual country level.

In order to do so we have to produce in the first place a measure of credit worth or default probability at the country level. Previous studies have used either a series of indicators of economic fundamentals (e.g. Min [18]) or a series of ratings expressed by major agencies like Moody's or S&P (e.g. Kamin and von Kleist, [15]). We do not want to enter into the problem of assessing whether agency rating or fundamentals are better indicators of credit worth than direct use of indicators (see page 8). In our own work we have found that ratings and fundamentals are strictly correlated and that a valuable measure of country risk can be obtained by calibrating a function of macro fundamentals to a restricted classification derived from Moody's ratings. In practice we have restricted the 16 categories given by Moody's to 5 classes according the following table:

Table 2
Assignment of numerical values to bond ratings

| Numerical value | Moody's |
|---|---|
| 5 | Aaa |
| 4 | Aa1, Aa2, Aa3 |
| 3 | A1, A2, A3, Baa1, Baa2, Baa3 |
| 2 | Ba1, Ba2, Ba3 |
| 1 | B1, B2, B3 |

We have then used an ordered probit specification, using a sample of 74 countries, to construct an index of country risk according to the following regression equation:

$$IR= Lf\,(INFL, GDPPC, CAEXP, DEFGDP, INT3M, DEBT)^{13},$$

where IR means index of risk and *Lf* stands for a linear function of:

*GDPPC*: per-capita income in dollars, as a measure of the state of the economic development; as the per capita income increases, we expect a rise in the probability to belong to the less risky class.

*INFL*: yearly inflation rate; the presence of high inflation as a sign of an incorrect economic policy implies greater risk for the economy considered. This variable should then present negative coefficient.

[13] The equation is fully reported in Appendix 1.

*CAEXP*: the current account-exports ratio points out possible external unbalances as a function of the hard currency revenues deriving from the exports; positive coefficient.
*DEFGDP*: nominal fiscal balance-GDP ratio, signalling the general quality of the economic policy; negative coefficient.
*INT3M*: real three-months interest rates, whose high level is to be interpreted as a sign of low financial stability. As such it presents negative coefficient.
*DEBT*: binary variable (dummy) which takes value 1 for countries that present a public debt-GDP ratio over 80% or external debt-GDP ratio over 60% and 0 for the others; for this dummy too we expect a negative coefficient.[14]

The variables considered in the model specification are those usually found in the above studies. We checked the stability of the coefficients on a selected number of dates of Moody's releases. We then computed the value of the index for every month of the sample and further normalised the index to the interval 0-1 by using a logit transformation. We then used this constructed variables in the sequel: they are bounded by 0 and 1 by the logit transformation. They appear as near unit processes. We do not interpret them as integrated variables and carry on the analysis specifying a conditional ECM between spreads and our index of macro-economic conditions.

At the disaggregated level we consider 7 countries that represent the main issuers of Brady Bonds. They include 4 Latin American countries - Mexico, Argentina, Brazil and Venezuela -, two eastern European – Bulgaria and Poland - and the Philippines. The countries form an unbalanced panel, since we have less observations for Poland, Bulgaria and Brazil, which start in 1995. The overall sample goes from 1992.01 to 1999.10: we have thus 530 unbalanced observations.

We now explore the properties of the panel with a LSDV specification.[15]

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[14] See Hardy and Pazarbasioglu [13] as an example of works in which is used a set of variables of regional relevance; in particular the authors apply bank crisis indicators to the analysis of the Asian crisis.

[15] We are aware of the shortcomings and possible biases of using the LSDV specification on time series cross section panels of inadequate length. We have 8 years of monthly observations: relying on Montecarlo evidence reported by many authors and most recently by Judson and Owen [14], ours can be regarded as an adequate sample. However we also are aware of the problems of heterogeneity raised by Pesaran, Smith and Im [20]. On this aspect we have relied heavily on a battery of Wald test, not always reported in the main body of the paper, that give us some confidence that slope heterogeneity is not a major problem of our specification. In fact we found an extremely homogenous set of coefficients even when estimating the fully disaggregated version of the panel. We do not consider the newer tests on cointegrated panels since we do not accept the hypothesis that spreads are integrated.

In order to appraise the correlation of our proposed measure of country risk we first consider the static specification between spreads and individual country risk.

Results are in Table 3: they show a strong negative association (as expected) of our index with the spreads; such association is clearly present in the data controlling for the crises. Individual effects are remarkably close.

Of course the above equation is badly specified: dynamics is absent, residuals show strong positive autocorrelation, other variables that we already found significant in the aggregate index are also missing.

Table 3
Static LSDV specification

| | | | | |
|---|---|---|---|---|
| Dependent Variable: LOG(SPREAD) | | | | |
| Method: GLS (Cross Section Weights) | | | | |
| Sample: 1992:01 1999:10 | | | | |
| Included observations: 94 | | | | |
| Number of cross-sections used: 7 | | | | |
| Total panel (unbalanced) observations: 530 | | | | |
| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
| LOGIT(IR) | -1.149 | 0.176 | -6.529 | 0.000 |
| CRISIS | 0.267 | 0.054 | 4.976 | 0.000 |
| Fixed Effects | | | | |
| MEX--C | 6.73 | | | |
| ARG--C | 7.06 | | | |
| BRA--C | 6.70 | | | |
| VEN--C | 7.03 | | | |
| POL--C | 6.03 | | | |
| BUL--C | 7.15 | | | |
| PHI--C | 7.00 | | | |
| R-squared | 0.893 | Durbin-Watson stat | | 0.148 |
| Adjusted R-squared | 0.891 | | | |
| S.E. of regression | 0.423 | | | |

A more complete LSDV specification is the following (Table 4), where we have imposed the restriction that the long run coefficient between the spreads and the risk is one. The Wald test for such restriction gives:

| | | | |
|---|---|---|---|
| F-statistic | 0.5899 | Probability | 0.4428 |
| Chi-square | 0.5899 | Probability | 0.4424 |

Examination of the coefficients of the equation shows that the process of adjustment to the fundamentals is rather slow. Only 10% of the distance from the equilibrium value is eliminated in the current month. On the other hand it is confirmed also at the individual country level that there is a significant impact effect of international rates. Oil price is significant with the expected negative sign since a number of the countries included are important oil producers. This last result is open to question, given that the sign is right for oil producing countries but it could be opposite for oil importing countries. The sample is dominated by countries like Mexico and Venezuela that are oil producers, but includes other countries like Bulgaria and the Philippines which are not. The assumption can be tested removing the restriction of the common coefficient for the oil price variable.

Table 4
Dynamic LSDV specification

Dependent Variable: DLOG(SPREAD)
Method: GLS (Cross Section Weights)
Sample: 1992:03 1999:10
Included observations: 92
Number of cross-sections used: 7
Total panel (unbalanced) observations: 517

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| DLOG(SPREAD(-1)) | 0.158 | 0.041 | 3.864 | 0.000 |
| D(LOGIT(IR)) | -0.354 | 0.122 | -2.898 | 0.004 |
| LOG(SPREAD(-1))+ LOGIT(IR(-1)) | -0.079 | 0.014 | -5.655 | 0.000 |
| DLOG(BRENT(-1)) | -0.194 | 0.061 | -3.197 | 0.001 |
| CRISIS | 0.172 | 0.018 | 9.340 | 0.000 |
| DLOG(1+FED3M/100) | 0.563 | 0.190 | 2.962 | 0.003 |
| Fixed Effects | | | | |
| MEX--C | 0.51 | | | |
| ARG--C | 0.52 | | | |
| BRA--C | 0.51 | | | |
| VEN--C | 0.54 | | | |
| POL--C | 0.43 | | | |
| BUL--C | 0.53 | | | |
| PHI--C | 0.52 | | | |

| | | | |
|---|---|---|---|
| R-squared | 0.264 | Mean dependent var | -0.006 |
| Adjusted R-squared | 0.247 | S.D. dependent var | 0.157 |
| S.E. of regression | 0.136 | Sum squared resid | 9.326 |
| Log likelihood | 319.05 | F-statistic | 36.24 |
| Durbin-Watson stat | 1.921 | Prob(F-statistic) | 0.000 |

The CRISIS variable can be similarly tested for heterogeneity. Finally, the hypothesis of asymmetry in the effect of international short rates that was apparent in the aggregate EMBI equation can also be tested.

In order to pursue a better understanding of the basic model we have followed the strategy of specifying it as a system of regression equations estimated by WLS. We have then tested heterogeneity and asymmetry with a battery of Wald tests. In the last step the equation has been estimated with SUR.

The final result of such specification strategy is reported in the following table. It deserves a number of comments, to facilitate the interpretation it may be useful to refer to part b) of the table looking to coefficient codes.

Coefficients C(8) to C(14) are fixed effect coefficients. We have nothing to observe save that they are remarkably close to each other indicating that there should be no other identifiable idiosyncratic effects in the country risk premia beyond those envisaged by the equation.

Table 5
SUR specification with asymmetry and individual country effects
a) estimated coefficients and standard errors

Estimation Method: Seemingly Unrelated Regression
Sample: 1992:03 1999:10
Included observations: 92
Total system (unbalanced) observations 517

| | Coefficient | Std. Error | t-Statistic |
|---|---|---|---|
| C(1) | 0.181 | 0.043 | 4.222 |
| C(2) | -0.149 | 0.086 | -1.734 |
| C(3) | -0.085 | 0.015 | -5.734 |
| C(4) | -0.259 | 0.081 | -3.216 |
| C(5) | 0.166 | 0.028 | 5.887 |
| C(6) | 18.24 | 5.645 | 3.231 |
| C(7) | -2.088 | 0.925 | -2.256 |
| Fixed effects: | | | |
| C(8) | 0.54 | 0.10 | 5.41 |
| C(9) | 0.56 | 0.10 | 5.37 |
| C(10) | 0.54 | 0.10 | 5.36 |
| C(11) | 0.57 | 0.10 | 5.44 |
| C(12) | 0.57 | 0.10 | 5.43 |
| C(13) | 0.46 | 0.09 | 5.15 |
| C(14) | 0.57 | 0.11 | 5.28 |

Coefficient C(1) and C(2) are the short run adjustment parameters for the dependent and risk index variables.

Coefficient C(3) measures the speed of adjustment to the equilibrium vector.

Coefficient C(1), C(2) and C(3) contribute to the dynamic process of adjustment of spreads to the country index of credit worth. Since in our metric the long run parameter is unity, a 1% change in credit worth is ultimately translated in 1% change in spreads.

b) country equations

Equation: DLOG(SPREADMEX) = C(8) + C(1)*DLOG(SPREADMEX(-1)) + C(2)*D(LOGIT(IR_MEX)) + C(3)*(LOG(SPREADMEX(-1)) + LOGIT(IR_MEX(-1)) + C(4)*DLOG(BRENT(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

| | | | |
|---|---|---|---|
| Observations: 92 | | | |
| R-squared | 0.327 | Durbin-Watson stat | 2.038 |
| Adjusted R-squared | 0.279 | | |

Equation: DLOG(SPREADARG) = C(9) + C(1)*DLOG(SPREADARG(-1)) + C(2)*D(LOGIT(IR_ARG)) + C(3)*(LOG(SPREADARG(-1)) + LOGIT(IR_ARG(-1)) + C(4)*DLOG(BRENT(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

| | | | |
|---|---|---|---|
| Observations: 76 | | | |
| R-squared | 0.339 | Durbin-Watson stat | 1.904 |
| Adjusted R-squared | 0.282 | | |

Equation: DLOG(SPREADBRA) = C(10) + C(1)*DLOG(SPREADBRA(-1)) + C(2)*D(LOGIT(IR_BRA)) + C(3)*(LOG(SPREADBRA(-1)) + LOGIT(IR_BRA(-1)) + C(4)*DLOG(BRENT(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

| | | | |
|---|---|---|---|
| Observations: 52 | | | |
| R-squared | 0.399 | Durbin-Watson stat | 2.116 |
| Adjusted R-squared | 0.319 | | |

Equation: DLOG(SPREADVEN) = C(11) + C(1)*DLOG(SPREADVEN(-1)) + C(2)*D(LOGIT(IR_VEN)) + C(3)*(LOG(SPREADVEN(-1)) + LOGIT(IR_VEN(-1)) + C(4)*DLOG(BRENT(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

| | | | |
|---|---|---|---|
| Observations: 92 | | | |
| R-squared | 0.329 | Durbin-Watson stat | 1.637 |
| Adjusted R-squared | 0.282 | | |

Equation: DLOG(SPREADPHI) = C(12) + C(1)*DLOG(SPREADPHI(-1)) + C(2)*D(LOGIT(IR_PHI)) + C(3)*(LOG(SPREADPHI(-1)) + LOGIT(IR_PHI(-1)) + C(6)*D(1+FED3M/100)*(D(FED3M)>0) + C(7)*DLOG(CRB)

| | | | |
|---|---|---|---|
| Observations: 92 | | | |
| R-squared | 0.157 | Durbin-Watson stat | 1.873 |
| Adjusted R-squared | 0.108 | | |

Equation: DLOG(SPREADPOL) = C(13) + C(1)*DLOG(SPREADPOL(-1)) + C(2)*D(LOGIT(IR_POL)) + C(3)*(LOG(SPREADPOL(-1)) + LOGIT(IR_POL(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

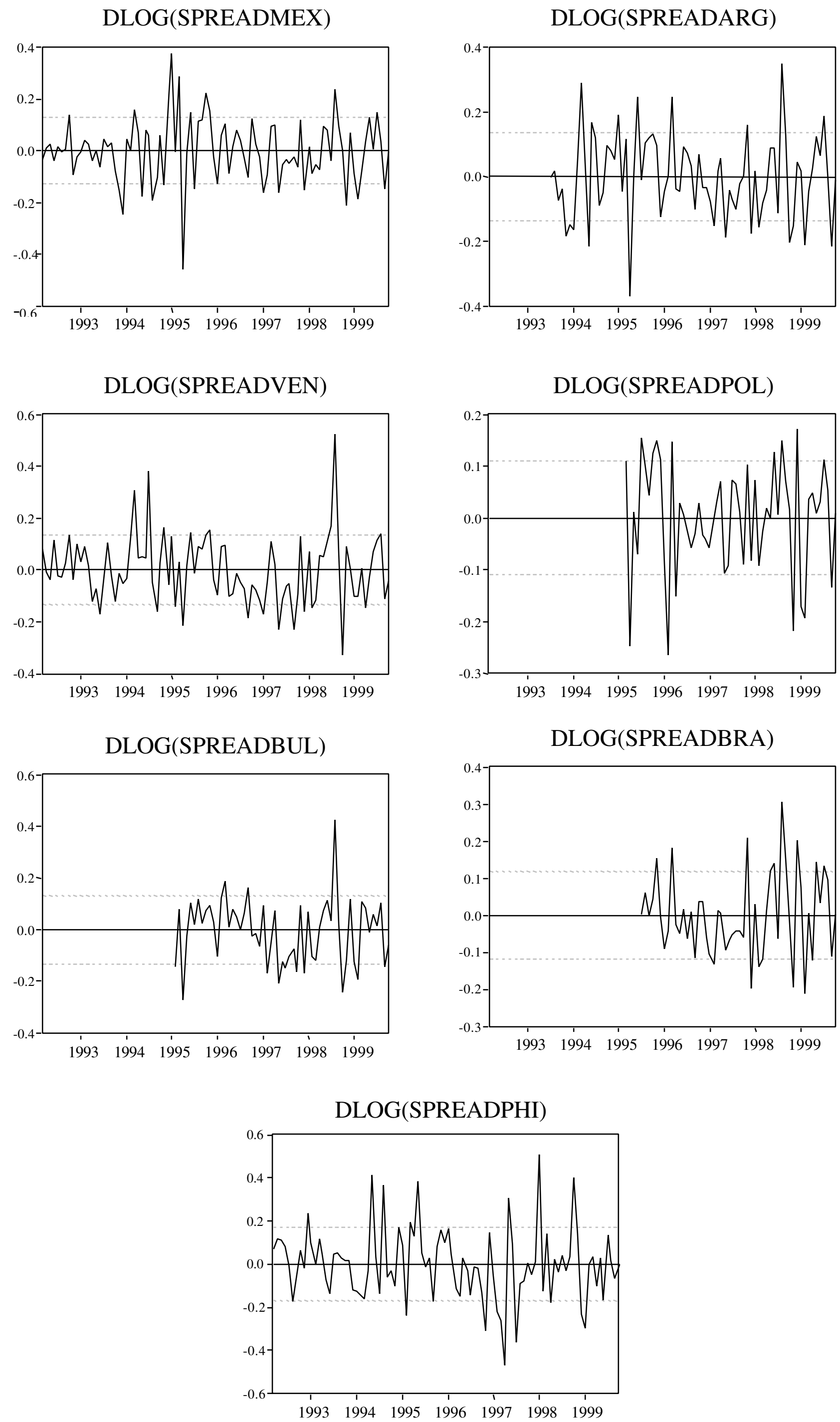
| | | | |
|---|---|---|---|
| Observations: 56 | | | |
| R-squared | 0.231 | Durbin-Watson stat | 2.252 |
| Adjusted R-squared | 0.154 | | |

Equation: DLOG(SPREADBUL) = C(14) + C(1)*DLOG(SPREADBUL(-1)) + C(2)*D(LOGIT(IR_BUL)) + C(3)*(LOG(SPREADBUL(-1)) + LOGIT(IR_BUL(-1)) + C(5)*CRISIS + C(6)*D(1+FED3M/100)*(D(FED3M)>0)

| | | | |
|---|---|---|---|
| Observations: 57 | | | |
| R-squared | 0.163 | Durbin-Watson stat | 1.840 |
| Adjusted R-squared | 0.081 | | |

Chart 4

Residuals of SUR system




Residuals are reasonably well behaved; around 1997 there is short span where a common overestimation of the spreads indicates that the model does not account completely for the sharp reduction in spreads after the Mexican crisis. Residuals also appear to exhibit some volatility, however we did not found significant ARCH effects except for Mexico.

Numerical values are such that approximately 14% of the permanent change happens in the same month, 75% within one year and 95% in two years.

Coefficient C(4) refers to the oil price variable. It is not significant for Poland, Bulgaria and the Philippines and it has been excluded from these country equations.

Coefficient C(5) refers to the common effect of international crises. It is not significant and it has been excluded from the Philippines equation.

Coefficient C(6) test the asymmetric effect of international short term rate. It is highly significant. The symmetric coefficient for negative variations of the same rate is insignificant (t-ratio=0.5) and has been excluded from the equation.

Coefficient C(7) refers to an index ex-oil of commodities, the CRB index. It is significant only for the Philippines, which is an important food producer.

Coefficients (C4) for oil price, C(6) for international rates and C(7) for food prices, reflect transient phenomena.

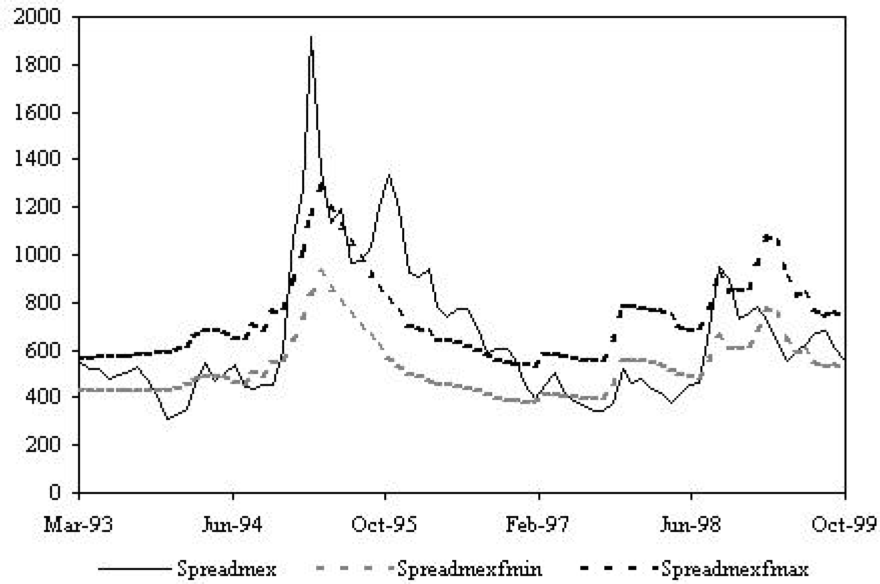
The fit of individual equations is not poor, averaging 30%; it appears somewhat better for Latin American countries. Residuals are well behaved (see Chart 4), they display some volatility which is however significant only for Mexico according to a LM(3) test.

Numerical values of the estimated coefficients imply a non marginal role for macro fundamentals. Let's illustrate it using data for Mexico.

Chart 5 graphs the behaviour of the actual spread for Mexico and two simulated series given by the results of a dynamic simulation over the entire period, using as values for the index of risk those corresponding to the range of the actual series. For Mexico the range goes from a minimum of 0.14 to a maximum of 0.48 with median value 0.33. The two simulated series are obtained in correspondence to the two extreme values kept constant.

The average distance between the two bands is approximately 200 basis points. Such is then the range of variation of the spreads that can be attributed to changes in risk rating in the case of Mexico. From the graphs one sees that the actual spread wanders in and out of the corridor reverting to it. Similar values hold for the other countries of the panel.

Chart 5
The spread for Mexico: the role of fundamentals




## 4. CONCLUSIONS

We take four main conclusions from the above empirical analysis. First, we have found a significant role in the behaviour of spreads for an indicator of macro fundamental conditions. Spreads tend to adjust to macro fundamentals with a rather long lag, however we estimate that 75% of the adjustment takes place in the same year.

Second, we have found that spreads are significantly affected by shock factors. We have isolated a role for commodity prices: interestingly, oil producing countries benefit from oil prices hikes and, similarly, food producing countries from the dynamics of the CRB index, which contains a large number of agricultural products and soft commodities.

Third, and this seems to constitutes a novel result, we have found a significant asymmetric effect for international rates. The result may help explain the contradictory evidence that has been reported in the literature: the sensitivity of spreads appears greater in the case of an upturn in international rates, possibly in anticipation of balance-of-payments difficulties. This asymmetric effect is also important for its potential implications for the asset allocation activity. The "appetite for risk" hypothesis seems instead refused by our model, confirming the conclusions of much literature about the little role for the loose monetary stance in industrialised countries on the decline of the spreads in 1996-1997. More convincingly the phenomenon seems to be attributed, as previously mentioned, to the longer-term process of globalisation in financial markets that has favoured a diversification of which the emerging markets have benefited.

Fourth, in examining the data we have matured the opinion that equations like ours and those similar reported in the literature are not apt to explain or predict the insurgence of crises in emerging market. Crises arise for other factors rather than a bad state of fundamentals and once they have started they spread to other countries[16]. In our equation, in order to isolate the role of fundamentals, we have neutralised the effect of crises by using a rather rough dummy variable that obviously mimics the interdependence. However the interdependence and contagion problems are important and cannot be studied by dummies alone: they obviously deserve further work and possibly a different econometric technique.[17]

[16] See Calvo [1] and Fernández -Arias [10].

[17] An interesting recent strand of literature considers the distinction between interdependence and contagion effect in the analysis of the transmission of financial crises. (Forbes-Rigobon [11] ).

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## DATA APPENDIX

*Dependent Variables*

RATINGS: Sovereign credit ratings were taken from Moody's Investors Service.

SPREADS: Sovereign spreads were collected from J.P. Morgan's web site (http://www.morganmarkets.com). They refer to the Emerging Markets Bond Index (EMBI), which is a total-return index of U.S. dollar denominated Brady bonds and other sovereign restructured bonds.

*Dummy Variables*

CRISIS: it takes value 1 in coincidence with periods of major financial crises, 0 otherwise.

*Solvency Variables*

DEFGDP: ratio of nominal fiscal deficit to GDP; sources: OECD, DRI, IFS.

DEBT: dummy variable which takes value 1 if public debt/GDP is over 80% or external debt/GDP is over 60%, 0 otherwise; sources: OECD, BIS, IIF, IFS.

*Macroeconomic Fundamentals*

INFL: yearly inflation rate; sources: DRI, IFS.

GDPPC: per-capita income; sources: DRI, IFS.

CAEXP: current account-exports ratio; sources: DRI, IFS.

INT3M: real three-month interest rates; sources: Bloomberg, IFS.

*External Factors*

BRENT: nominal price of oil (US$ per barrel).

CRB: Commodity Research Bureau Index (spot price of commodities ex-oil).

FED3M: three-month Fed Funds rate.

Countries considered in the probit regression are:
Argentina, Australia, Austria, Bahamas, Bahrain, Barbados, Belarus, Belgium, Bolivia, Brazil, Bulgaria, Canada, Chile, China, Colombia, Côte d'Ivoire, Croatia, Cyprus, Czech Republic, Denmark, Ecuador, Egypt, Estonia, Finland, France, Germany, Greece, Hong Kong, Hungary, Iceland, India, Indonesia, Iran, Ireland, Israel, Italy, Japan, Jordan, Kazakhstan, Kuwait, Latvia, Lebanon, Lithuania, Luxembourg, Malaysia, Malta, Mauritius, Mexico, Morocco, Netherlands, New Zealand, Norway, Oman, Pakistan, Peru, Philippines, Poland, Portugal, Qatar, Romania, Russia, Saudi Arabia, Singapore, Slovakia, Slovenia, South Korea, Spain, South Africa, Sweden, Switzerland, Taiwan, Thailand, Tunisia, Turkey, UAE, Ukraine, United Kingdom, United States, Uruguay, Venezuela.

## APPENDIX 1: THE ORDERED PROBIT EQUATION FOR REPLICATING MOODY'S RATINGS

### A1.1 The ordered probit equation

As explained in the text, in order to gauge the importance of fundamentals in the spread equation of our panel, instead of using a number of economic condition indicators we have resorted to a summary measure offered by the index function of an ordered probit equation. The equation cited in the text is the following:

Dependent Variable: IR

| Numerical value | Moody's |
|---|---|
| 5 | Aaa |
| 4 | Aa1, Aa2, Aa3 |
| 3 | A1, A2, A3, Baa1, Baa2, Baa3 |
| 2 | Ba1, Ba2, Ba3 |
| 1 | B1, B2, B3 |

Method: ML - Ordered Probit
Included observations: 74 - Period 1998.Q1
Number of ordered indicator values: 5

| | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| GDPPC | 0.0001 | 2.E-05 | 5.6683 | 1.E-08 |
| INFL | -0.0301 | 0.0102 | -2.9565 | 0.0031 |
| CAEXP | 1.2456 | 0.5991 | 2.0792 | 0.0376 |
| DEBT | -0.8594 | 0.3552 | -2.4193 | 0.0156 |
| DEFGDP | -6.3527 | 3.3153 | -1.9161 | 0.0553 |
| INT3MRC | -0.1032 | 0.0446 | -2.3160 | 0.0206 |
| Limit Points | | | | |
| LIMIT_1:C(7) | -1.9061 | 0.5377 | -3.5448 | 0.0004 |
| LIMIT_2:C(8) | -0.4132 | 0.4358 | -0.9480 | 0.3431 |
| LIMIT_3:C(9) | 1.5623 | 0.5181 | 3.0154 | 0.0026 |
| LIMIT_4:C(10) | 3.3045 | 0.6729 | 4.9110 | 9.E-07 |

| | | | |
|---|---|---|---|
| Akaike info criterion | 2.0226 | Schwarz criterion | 2.3339 |
| Log likelihood | -64.834 | Hannan-Quinn criterion | 2.1468 |
| Restr. log likelihood | -114.10 | Avg. log likelihood | -0.8761 |
| LR statistic (6 df) | 98.539 | LR index (Pseudo-R2) | 0.4318 |
| Probability(LR stat) | 0.0000 | | |

Variables are defined in the Data Appendix.

Note: country variables corresponding to INT3MC have been censored at 0 and 12%.

## A1.2 Time series properties of the rating variables

We have constructed measures of credit worth for the individual countries as a linear combination of the macro-variables, using as weights the coefficients of the probit equation. We have then normalised the resulting series with a logit transformation. As a result, our measure of country risk (called LOIR) is bounded between zero and one. The time series properties of the resulting variable are somewhat mixed. ADF statistics reported in Table A2 indicate that they can be seen as near integrated processes (namely AR(1) with large autoregressive parameter). It is well known that ADF statistics are unreliable when the series are likely to contain level shifts. We believe this to be the case. We identify as level shift effects the changes in creditworthiness of the country induced by changes in the fundamentals.

Table A2
ADF statistics for the measure of country risk (LOIR)

| **Country** | Test equation.: $d(y_t)=c_1 + c_2y_t(-1) + c_3y_t(-1)) + u_t$ | | |
|---|---|---|---|
| Mexico | -2.174 | 1% Critical Value* | -3.502 |
| | | 5% Critical Value | -2.892 |
| Argentina | -4.761 | 1% Critical Value* | -3.502 |
| | | 5% Critical Value | -2.892 |
| Venezuela | -2.473 | 1% Critical Value* | -3.502 |
| | | 5% Critical Value | -2.892 |
| Brazil | -2.311 | 1% Critical Value* | -3.562 |
| | | 5% Critical Value | -2.919 |
| Poland | -3.800 | 1% Critical Value* | -3.550 |
| | | 5% Critical Value | -2.913 |
| Bulgaria | -1.291 | 1% Critical Value* | -3.503 |
| | | 5% Critical Value | -2.893 |
| Philippines | -2.656 | 1% Critical Value* | -3.502 |
| | | 5% Critical Value | -2.892 |

The assumption of stationarity for these variable is reinforced by the following panel regression:

Table A3
Panel estimation of the DF equation of the seven countries

| Dependent Variable: D(LOIR) | | | | |
|---|---|---|---|---|
| Method: GLS (Cross Section Weights) | | | | |
| Sample: 1992:03 1999:10 | | | | |
| Included observations: 92 | | | | |
| Number of cross-sections used: 7 | | | | |
| Total panel (unbalanced) observations: 566 | | | | |
| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
| LOIR | -0.124 | 0.019 | -6.503 | 0.000 |
| D(LOIR(-1)) | 0.072 | 0.041 | 1.738 | 0.083 |
| Fixed Effects | | | | |
| ARG--C | 0.059 | | | |
| BRA--C | 0.016 | | | |
| VEN--C | 0.027 | | | |
| MEX--C | 0.041 | | | |
| PHI--C | 0.103 | | | |
| POL--C | 0.055 | | | |
| BUL--C | 0.025 | | | |
| R-squared | 0.073 | Mean dependent var | | 0.003 |
| Adjusted R-squared | 0.059 | S.D. dependent var | | 0.048 |

In the following chart we show the time series profile of the series. One can see clear level shifts in most of the series:

Chart A1
Stacked line graph of the country risk measures