Sovereign spreads, currency crises, and fundamentals: A non-linear analysis

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Abstract

This paper presents a model showing how sovereign bond spreads in emerging markets may be affected by market expectations on exchange rate dynamics. This allows to explain why most of previous studies fail to relate sovereign spreads to underlying fundamental variables, especially over periods of financial distress. A Markov-switching VAR model deriving from the theoretical framework is estimated using Brazilian data. The results show that the non-linear multivariate specification of sovereign spreads dynamics outperforms its linear counterpart as well as a number of other univariate models.

Keywords: Sovereign spreads, Currency crises, Multiple equilibria, Emerging markets, Markov-switching VAR.
JEL Classification: F31, F32, F34, C32, C51.

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1 Introduction

In this paper, I use the framework developed in Boschi and Goenka (2007) to show that the pattern of sovereign bond spreads in emerging markets is affected by expectations on exchange rate dynamics and, more specifically, by the probability of a currency crisis occurring. This allows to reconcile the behaviour of economic fundamentals with that of sovereign spreads over periods of financial distress. Previous literature suggests that this relationship is non-linear, breaking down during crisis periods, thus casting doubts on the efficiency of international financial markets. The model of Boschi and Goenka (2007) can account for such non-linearity explaining it as the outcome of rational investors optimizing behaviour, allowing at the same time multiple equilibrium solutions. Further, I show how this model can be implemented empirically through the specification of a Vector Autoregression (VAR) model subject to Markov-switching (MS) regimes.

Given the growing market size of emerging countries' bonds, understanding the dynamic behaviour and the determinants of sovereign spreads has important implications for research as well as for portfolio management. However, while the literature has devoted a lot of attention to emerging markets' equities (e.g. Harvey (1995), Bekaert and Harvey (1995)), only recently research has leaned towards bond returns. The resurgence of capital flows to developing countries after the debt crisis of the mid-eighties has renewed the interest in sovereign bonds, one of the main instruments of the recent international financial integration.

Early studies on emerging markets' spreads focused on bank loans. During the nineties these were the almost exclusive channel of access to international capital markets for developing countries. In a seminal paper, Edwards (1984) developed a framework commonly used in subsequent research on the valuation of sovereign risk premium. Emerging markets are modelled as small borrowers on perfectly competitive financial markets. Their fair value spreads are determined by the probability of default, which in turn depends on domestic macroeconomic fundamentals and on external shocks. Edwards (1984) applies this framework to estimate the determinants of primary yields on bank lending to emerging countries, while Edwards (1986) extends the analysis to estimate separately the default risk premia on international bank loans and bond markets. His results support the presumption that bonds are riskier than bank loans.

A subset of the literature on sovereign spreads is concerned on comparing the relative importance of domestic and foreign factors. Fernàndez-Arias (1996), for example, finds that US government bond yields have a larger impact than domestic fundamentals. Using the same methodology as Edwards (1986), Min (1998) investigates the primary bond market finding that the terms of trade are the only external factor affecting spreads, while the role of US short-term interest rates and the oil price is insignificant. Kamin and Kleist (1999) study the influence of borrower's creditworthiness, as measured by the credit ratings issued by the major rating agencies, on bond and loan launch yield spreads. They find that creditworthiness is an important determinant of emerging markets' spreads, while they cannot identify any robust and statistically significant relationship between sovereign spreads and industrial countries' interest rates. By contrast, Eichengreen and Mody (1998) find that the US government bond yield is negatively related to the issuance of sovereign bonds by emerging countries and, through the larger supply, to sovereign spreads.

Dungey *et al.* (2003) look at the impact of global risk aversion on emerging market debt in several crisis events. By decomposing global risk aversion in three factors, volatility, credit, and liquidity risks, they find that the Russian crisis was characterized by a sharp increase in global credit risk, while the relative size of global risk factors is mixed for the Brazilian crisis.

Jostova (2006) develops an error correction model to investigate the predictability of changes in Brady bond spreads as a function of a broad set of macroeconomic variables reflecting economic conditions and, in turn, the default probability of local governments. She finds evidence of return predictability in Brady bonds returns through a number of approaches to the evaluation of out-of-sample performance of the model. An important predictor is the deviation from the long-run equilibrium between spreads and the underlying economic fundamentals, while controlling for contagion effects from other emerging markets and the correlation with equity markets does not produce the expected results.

Ferrucci (2003) studies empirically the determinants of emerging markets' sovereign spreads in order to disentangle the relative contribution of country-specific fundamentals while controlling for global interest rates, market risk and liquidity in bond markets. One of the interesting findings of his paper is that while sovereign spreads reflect broadly market fundamentals, non-fundamental factors play a more important role. Ferrucci interprets this result as evidence against the efficient market hypothesis, according to which asset prices always reflect the information publicly available as processed by rational investors. By comparing the market-based spreads against their fundamental-based counterparts he concludes that the generalized fall in emerging markets spreads between 1995 and 1997 was due to capital market imperfections as reflected in higher investors' risk appetite resulting from lower global interest rates. The largest misalignments between actual and predicted spreads are reported for countries which are experiencing a financial crisis. Ferrucci comes to the conclusion that the break-down of the fundamental relationship during crisis periods represents evidence of market inefficiency. By contrast, in this paper I show that this can be the outcome of the optimization behaviour of a fully rational international investor with decreasing risk aversion in absence of market frictions.

Boschi and Goenka (2007) provides a theoretical framework to analyse the interaction between currency crises and investors' beliefs in the context of a second generation model of currency crises *à la* Jeanne (1997). It shows that the devaluation probability of pegged currencies affects international investors' wealth via portfolio returns. If investors are risk averse, a decrease in wealth will increase the risk premium they require on assets and this, in turn, will increase the government's cost of defending the peg. For certain values of the fundamentals and parameters, multiple equilibria and self-fulfilling crises may arise as a consequence of inherent circularity - investors' formulate rational expectations on the government commitment to defend the peg and, conversely, the government's behaviour is derived optimally conditioning on expected investors' beliefs. This argument helps to understand that the relationship between sovereign spreads, as a proxy for the country-specific risk premium, and fundamentals weakens when the pressure exerted by investors on the foreign exchange market increases.

Another contribution of this paper is methodological in nature. I show that a second generation model of currency crises with multiple equilibria has in a MS-VAR model a natural empirical counterpart. Jeanne and Masson (2000) characterize the equilibria of a reduced form version of a class of "escape clause" or "second generation" models of currency crises. They identify the conditions under which multiple equilibria can arise and those under which an arbitrarily large number of equilibria, rather than only three as in most other models, can occur. Most importantly, they show how this class of models can be consistently brought to data through the use of Markov-switching regimes econometric techniques, in which the switch across regimes corresponds to jumps between different equilibria. They provide an illustration of this modelling strategy by applying it to the 1993 French franc crisis and find that allowing for sunspots to influence the devaluation probability improves the model performance.

The results of the estimation of the MS-VAR as well as several other univariate models of Brazilian sovereign spreads support the hypothesis of a non-linear relationship between spreads and their fundamental determinants due to expectations about currency devaluations.

The rest of the paper is structured as follows. Section 2 sketches the theoretical framework. Section 3 describes the empirical methodology. Section 4 presents the results, and Section 5 concludes.

2 Theoretical framework

By including risk averse international investors in a second generation model of currency crises, the model developed in Boschi and Goenka (2007) shows that the probability of a currency devaluation and the investor beliefs are determined jointly through a complex interdependence between the government's and the private sector's expectations.

The representative international investor maximizes his expected utility over time.

$$U_t = E_t \left\{ \sum_{s=t}^{\infty} \beta^{s-t} u(C_s) \right\} \tag{1}$$

where E_t is the expectation operator conditional on all information available up to and including date t, $\beta \in (0, 1)$ is the constant subjective time-preference factor, $u(\cdot)$ is the period utility function which is assumed to be twice-continuously differentiable, strictly increasing, and strictly concave.

The investor's portfolio includes a risk-free asset as well as risky assets, either local currency-denominated and dollar-denominated, issued by M countries. Therefore, the investor's budget constraint is given by:

$$\begin{aligned} B_{s+1}^{*f} + \sum_{m=1}^{M} x_{s+1}^{*m} B_s^{*m} + \sum_{m=1}^{M} x_{s+1}^{m} \frac{B_s^m}{e_s^m} &= (1 + r_s^{*f}) B_s^{*f} \\ &+ \sum_{m=1}^{M} x_s^{*m} \left(I_s^{*m} + B_s^{*m} \right) \\ &+ \sum_{m=1}^{M} x_s^{m} \left(\frac{I_s^m + B_s^m}{e_s^m} \right) - C_s \quad (2) \end{aligned}$$

where B_s^{*f} is the real net risk-free bond purchase at time $s-1$, x_s^{*m} and x_s^m are respectively the fractional shares of country m's dollar- and local currency-denominated debt purchased by the agent in period $s-1$, B_s^{*m} and B_s^m denote respectively the date s real market value of country m's dollar- and local currency-denominated debt, r_s^{*f} is the net real interest rate on the risk-free bond B_s^{*f} between period $s-1$ and s, I_s^{*m} and I_s^m are the coupons paid on country m's securities at time s, and e_s^m is time s spot exchange rate (price of dollars in terms of country m's currency).

By maximizing the expected utility subject to the budget constraint, the investor obtains a set of Euler equations, one for each type of asset. Manipulating these three equations, one obtains the standard expression for the risk premium in the consumption-based international CAPM. By taking a second order Taylor expansion around the consumption mean, the risk premium can be written as follows (see Boschi and Goenka (2007) for details):

$$E_t(r_{t+1}^{*m}) - r_{t+1}^{*f} \approx (1 + r_{t+1}^{*f}) \beta \left(\frac{-C_t u''(C_t)}{u'(C_t)} \right) Cov \left\{ \frac{C_{t+1}}{C_t}, r_{t+1}^{*m} \right\}. \quad (3)$$

If we assume a decreasing relative risk aversion, $-u''(C_t)/u'(C_t)$, then a negative relationship between the risk premium and aggregate consumption is established.

By equating the right-hand sides of the Euler equations for the local currency-denominated and the dollar-denominated assets, assuming that the covariance term is equal across assets, and taking logarithms, one obtains the following version of the Uncovered Interest Parity (UIP) relationship under risk aversion:

$$E_t(r_{t+1}^m) = r_{t+1}^{*f} + \rho_{t+1}^m + \pi_{t+1}^m \Delta e \quad (4)$$

where $\rho_{t+1}^m \equiv E_t(r_{t+1}^{*m}) - r_{t+1}^{*f}$, π_{t+1}^m is the probability of devaluation of country m's currency occurring at time $t+1$, and $\Delta e = \ln\left(e_{t+1}^m / e_t^m\right)$ is the proportional, time-invariant, extent of such devaluation, where Δ is the difference operator.

In order to link the time-varying risk premium to the probability of a currency devaluation in one of the countries whose assets the investor holds, I rewrite the budget constraint (2) as follows:

$$W_t^* = \left(W_{t-1}^* - C_{t-1} \right) \left(1 + r_t^{*w} \right) \quad (5)$$

where W_t^* denotes total real wealth (denominated in dollars) and $(1 + r_t^{*w})$ is the gross real return on wealth invested from period $t-1$ to period t. Given international portfolio diversification, the *ex post* gross return can be decomposed as follows:

$$(1 + r_t^{*w}) = q_t^{*f}(1 + r_t^{*f}) + \sum_{m=1}^{M} q_t^{*m}(1 + r_t^{*m}) + \sum_{m=1}^{M} q_t^m \frac{e_{t-1}^m}{e_t^m}(1 + r_t^m) \qquad (6)$$

where q_t^{*f} is the proportion of wealth invested in the risk-free bond and q_t^{*m} and q_t^m are, respectively, the proportions of wealth invested in country m's dollar- and local currency-denominated assets at time $t-1$, implying that $q_t^{*f} + \sum_{m=1}^{M} q_t^{*m} + \sum_{m=1}^{M} q_t^m = 1$.

Taking logarithms of expectations of both sides of (6) gives:

$$\begin{aligned} E_{t-1}\left(r_t^{*w}\right) \approx & \log\{q_t^{*f}(1 + r_t^{*f}) + \sum_{m=1}^{M} q_t^{*m} \exp\left[E_{t-1}\left(r_t^{*m}\right)\right] \\ & + \sum_{m=1}^{M} q_t^m \exp\left[E_{t-1}\left(r_t^m\right) - \pi_t^m \Delta e\right]\} \end{aligned} \qquad (7)$$

In order to relate expectations about currency devaluations to unexpected changes in consumption and, thus, to wealth effects, Boschi and Goenka (2007) extends to such an internationally diversified portfolio framework the log-linear approximation to the budget constraint proposed by Campbell (1993). Thus:

$$\begin{aligned} c_t - E_{t-1}c_t = & (E_t - E_{t-1}) \sum_{j=0}^{\infty} \eta^j r_{t+j}^{*w} \\ & -(E_t - E_{t-1}) \sum_{j=1}^{\infty} \eta^j \Delta c_{t+j} \end{aligned} \qquad (8)$$

Therefore, an unexpected change in the portfolio's rate of return, r_t^{*w}, due to an increase of the devaluation probability in one country, π_{t+1}^m, will decrease the investor's wealth and make the investor more risk averse, thus raising the risk premium ρ_{t+1}^m. This relationship between the devaluation probability and the risk premium is now embedded in a second generation model of currency crises close to the one proposed by Jeanne (1997). As a result, the devaluation probability depends on the risk premium in a context where multiple equilibria may arise.

Consider an emerging market's government pegging the national currency to the US dollar. As long as the country's debt is partly denominated in local currency, the government faces a policy dilemma as it has an incentive to devalue in order to reduce the cost of debt. It will take any decisions about either maintaining the peg or not by minimizing the following quadratic loss function:

$$L_{t+1} = (u_{t+1})^2 + (\Delta d_{t+1})^2 - (\Delta tb_{t+1})^2 + \delta\Gamma_{t+1} \qquad (9)$$

where u_{t+1} is the unemployment rate, Δd_{t+1} is the growth in government real debt proportional to GDP, Δtb^a_{t+1} is the trade balance proportional to GDP, δ is a dummy variable which is equal to 1 if devaluation occurs and 0 if the peg is maintained, and Γ_{t+1} denotes the exogenous cost of devaluing.

The net benefit of maintaining the peg is given by:

$$\overline{V}_{t+1} = L^d_{t+1} - L^f_{t+1}$$

where L^d_{t+1} and L^f_{t+1} are the loss functions corresponding to the devaluation and to the fixed peg hypothesis respectively. The gross benefit of maintaining the peg, V_{t+1}, summarizes the fundamentals of the emerging economy. Assuming that V_{t+1} is governed by the process:

$$V_{t+1} - \phi_{t+1} = \epsilon_{t+1} \sim \text{i.i.d.N}(0, \sigma_\epsilon) \tag{10}$$

where:

$$\phi_{t+1} = E_t\left(V_{t+1}\right) \tag{11}$$

then, the investor's rational behaviour implies that his expectations about the devaluation probability depend on the net benefit of keeping the peg, i.e.:

$$\pi_{t+1} = \Pr[\overline{V}_{t+1} < 0]$$

Proposition 4 in Boschi and Goenka (2007) shows that this probability may be rewritten as:

$$\pi_{t+1} = \Pr[\epsilon_{t+1} < \alpha\pi_{t+1} + \varphi\rho_{t+1} - \phi_{t+1}] \tag{12}$$

or:

$$\pi_{t+1} = F[\alpha\pi_{t+1} + \varphi\rho_{t+1} - \phi_{t+1}] \tag{13}$$

where $F(\cdot)$ is the cumulative distribution function of $f(\cdot)$, with the latter being the density function of ϵ_{t+1}. Boschi and Goenka (2007) show that under certain conditions equation (13) may have multiple solutions.

3 Empirical methodology

In this Section I derive an empirical formulation of the theoretical framework sketched above, and discuss briefly the econometric methodology employed to estimate it. Extending Jeanne and Masson (2000), I show how the class of second generation models of currency crises exemplified by (13) can be estimated through a Markov-switching regime modelling approach.

Besides Jeanne and Masson (2000), MS regime models have been widely and increasingly used in the empirical literature on currency and financial crises. Martinez Peria (2002), for example, studies the European Monetary System crises using a regime-switching approach. She estimates a univariate and a VAR model of the exchange rate, both with time varying transition probabilities. The VAR model includes exchange rates, interest rates differentials, and foreign exchange reserves as dependent variables. Allowing for time-varying probabilities makes it possible to estimate the determinants of probabilities of switching from a tranquil to a crisis state. Her main results show that both fundamentals and expectations (proxied by interest rate differentials) affect the likelihood of switching from one state to another. Fratzscher (2003) estimates a univariate Markov-switching model and a panel data model for a sample of 24 emerging market and East European transition countries. He shows how the inclusion of contagion variables measuring the financial and real *proximity* of countries helps to improve the performance of Markov-switching models of currency crises over models that only consider domestic fundamentals. Cipollini *et al.* (2007) extend the work of Jeanne and Masson (2000) by allowing for time-varying transition probabilities to study the causes of the European Monetary System crisis of 1992-1993. They find that fundamental variables were important in understanding the currency crises, beyond the effect of market expectations driven by external uncertainty.

3.1 Empirical specification

Jeanne and Masson (2000) argue that in a model similar to (13) jumps between states with different devaluation expectations may be driven by extrinsic uncertainty – sunspots which coordinate the private sector beliefs on a given state. Under sunspots the devaluation probability, π_t, depends jointly on the state, s_t, on the risk premium, ρ_t, and on the fundamental variable, ϕ_t, according to the following equation:

$$\pi_t = \sum_{s=1}^{n} \theta(s_t, s) F(\rho_t, \phi_t, \phi_s^*) \tag{14}$$

where s_t is the current state, $s = 1, ..., n$ are the possible future states in which the economy can be, each differing from the others according to the level of fundamentals triggering devaluation, ϕ_s^*, $\theta(s_t, s)$ is the probability of transition from state s_t to state s, and F is defined above.

Taking a first order Taylor approximation of equation (14) gives:[1]

$$\pi_t = \gamma_{s_t} + \varphi_1 \rho_t + \boldsymbol{\psi}_1' \mathbf{x}_t + v_t, \qquad s_t = 1, ..., n \tag{15}$$

where γ_{s_t} is an intercept term depending on the state, φ_1 is a coefficient, $\boldsymbol{\psi}_1 = (\psi_{11}, ..., \psi_{k1})'$ is a vector of coefficients, $\mathbf{x}_t = (x_{1t}, ..., x_{kt})'$ is a vector of relevant economic fundamentals, and v_t is an i.i.d. shock..

Equation (15) differs from the Jeanne and Masson (2000) model for the risk premium term, $\varphi_1 \rho_t$ deriving from the explicit consideration of the risk averse international investor's behaviour. This term is endogenously determined and depends itself on the devaluation probability, π_t, according to equations (3) and (8). This relationship can be linearized as follows:[2]

$$\rho_t = \mu_{s_t} + \varphi_2 \pi_t + \boldsymbol{\psi}_2' \mathbf{x}_t + e_t, \qquad s_t = 1, ..., n \tag{16}$$

[1] See the Appendix for details.

[2] See the Appendix for details.

where μ_{s_t} is an intercept term depending on the state, φ_2 is a coefficient, $\boldsymbol{\psi}_2 = (\psi_{12}, ..., \psi_{k2})'$ is a vector of coefficients, $\mathbf{x}_t$ is defined above, and e_t is an i.i.d. shock. Equations (15) and (16) can be stacked in a VAR with the following structural form:

$$\begin{bmatrix} 1 & -\varphi_1 \\ -\varphi_2 & 1 \end{bmatrix} \cdot \begin{bmatrix} \pi_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} \gamma_{s_t} \\ \mu_{s_t} \end{bmatrix} + \begin{bmatrix} \boldsymbol{\psi}_1' \\ \boldsymbol{\psi}_2' \end{bmatrix} \cdot \mathbf{x}_t + \begin{bmatrix} v_t \\ e_t \end{bmatrix} \tag{17}$$

or

$$\mathbf{B}\mathbf{y}_t = \Gamma_0(s_t) + \boldsymbol{\Psi}\mathbf{x}_t + \boldsymbol{\epsilon}_t$$

where

$$\begin{aligned} \mathbf{B} &= \begin{bmatrix} 1 & -\varphi_1 \\ -\varphi_2 & 1 \end{bmatrix}, \quad \mathbf{y}_t = \begin{bmatrix} \pi_t \\ \rho_t \end{bmatrix}, \quad \Gamma_0 = \begin{bmatrix} \gamma_{s_t} \\ \mu_{s_t} \end{bmatrix} \\ \boldsymbol{\Psi} &= \begin{bmatrix} \boldsymbol{\psi}_1' \\ \boldsymbol{\psi}_2' \end{bmatrix}, \quad \boldsymbol{\epsilon}_t = \begin{bmatrix} v_t \\ e_t \end{bmatrix}. \end{aligned}$$

Premultiplying by $\mathbf{B}^{-1}$ allows us to obtain the VAR model in reduced form:

$$\mathbf{y}_t = \boldsymbol{\nu}(s_t) + \boldsymbol{\Pi}\mathbf{x}_t + \boldsymbol{\varepsilon}_t$$

where $\boldsymbol{\nu}(s_t) = \mathbf{B}^{-1}\Gamma_0(s_t)$, $\boldsymbol{\Pi} = \mathbf{B}^{-1}\boldsymbol{\Psi}$, and $\boldsymbol{\varepsilon}_t = \mathbf{B}^{-1}\boldsymbol{\epsilon}_t$.

In proceeding further to the empirical specification, note that the theoretical model sketched above does not block a dynamic, i.e. lagged, relationship between the devaluation probability, π_t, and the risk premium, ρ_t. Therefore, the MS-VAR model that I estimate is a reduced-form, dynamic generalization of the model (17) given by:

$$\mathbf{y}_t = \boldsymbol{\nu}(s_t) + \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \sum_{j=0}^{q} \boldsymbol{\Pi}_j \mathbf{x}_{t-j} + \boldsymbol{\varepsilon}_t, \quad s_t = 1, ..., n \tag{18}$$

where the intercept $\boldsymbol{\nu}(s_t)$ is regime-dependent.

This configures a VAR model whose intercepts are allowed to switch between regimes which are assumed to follow a hidden 2-state Markov chain stochastic process. In the terminology of Krolzig (1997) this is a MSI(2)-VARX(p, q) model, where X denotes the inclusion of exogenous variables, and with p and q being the lag orders of the endogenous and exogenous variables respectively. Note that the number of states follows from the theoretical model, according to which there are two alternatives regimes: crisis and non-crisis.

3.2 Markov-switching VAR models

The Markov-switching model has been popularized in time series analysis by Hamilton (1988, 1989, 1994). Its multivariate version extends a standard linear VAR model by allowing its parameters to be subject to regime shifts. In such situation, rather than time-varying, the VAR process can be modelled as time-invariant conditional on an unobservable regime variable s_t which indicates the regime prevailing at time t. Therefore, Markov-switching vector autoregressions are generalizations of the basic VAR model of order p:

$$\mathbf{y}_t = \boldsymbol{\nu} + \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{u}_t \tag{19}$$

where $\mathbf{y}_t = (y_{1t}, ..., y_{Kt})'$ is a K-dimensional vector, $\boldsymbol{\nu}$ is an intercept term, $\mathbf{A}_i$, for $i = 1, ..., p$, are $K \times K$ matrices of coefficients, and $\mathbf{u}_t$ is a vector of residuals. Denoting $\mathbf{A}(L) = \mathbf{I}_K - \mathbf{A}_1 L + ... + \mathbf{A}_p L^p$ as the lag polynomial of dimension $K \times K$, I assume that there are no roots on or inside the unit circle $|\mathbf{A}(z)| \neq 0$ for $|z| \leq 1$, where L is the lag operator. Under the additional assumption that $\mathbf{u}_t \sim NID(\mathbf{0}, \boldsymbol{\Sigma})$, equation (19) is the intercept form of a stable *Gaussian* VAR model of order p.

Since I assume that the parameters of the *observed* time series vector $\mathbf{y}_t$ depend on the *unobservable* regime variable s_t, a model for the regime generating process is required. In the Markov-switching VAR model the regime $s_t \in \{1, ...M\}$ is assumed to be governed by a discrete time, discrete state Markov stochastic process characterized by the following transition probabilities:

$$p_{ij} = \Pr(s_{t+1} = j \mid s_t = i), \ \sum_{j=1}^{M} p_{ij} = 1 \ \forall i, j \in \{1, ...M\}. \tag{20}$$

The transition probabilities can be represented by the following transition matrix:

$$\mathbf{P} = \begin{bmatrix} p_{11} & p_{12} & \cdots & p_{1M} \\ p_{21} & p_{22} & \cdots & p_{2M} \\ \vdots & \vdots & \ddots & \vdots \\ p_{M1} & p_{M2} & \cdots & p_{MM} \end{bmatrix} \tag{21}$$

where $p_{iM} = 1 - p_{i1} - ... - p_{i,M-1}$ for $i = 1, ..., M$. A crucial assumption for the theoretical properties of MS-VAR models is that s_t follows an irreducible ergodic M state Markov process with transition matrix given by (21).

Therefore, if time series are subject to shifts in regime, the M regimes Markov-switching form of the VAR(p) model of equation (19) is given by:

$$\mathbf{y}_t = \boldsymbol{\nu}(s_t) + \sum_{i=1}^{p} \mathbf{A}_i(s_t) \mathbf{y}_{t-i} + \mathbf{u}_t. \tag{22}$$

where $\mathbf{u}_t \sim NID(\mathbf{0}, \boldsymbol{\Sigma}(s_t))$ and $\mathbf{A}_1(s_t), ..., \mathbf{A}_p(s_t), \boldsymbol{\Sigma}(s_t)$ are shifts functions describing the dependence of the parameters $\mathbf{A}_1, ..., \mathbf{A}_p, \boldsymbol{\Sigma}$ on the realized regime s_t.

The parameters of the model are estimated with the maximum likelihood method (see Hamilton (1989, 1994)). The maximization of the likelihood function of an MS-VAR requires an iterative estimation of the parameters of the autoregression and the transition probabilities governing the Markov chain of the unobserved states. This is usually obtained through the implementation of the *Expectation Maximization* (EM) algorithm introduced by Dempster *et al.* (1977) and proposed by Hamilton (1990) for this class of models.

4 Empirical illustration

The Brazilian economy represents an interesting case study given the frequent recurrence of financial and currency crises. Brazil has undergone three main periods of financial distress over the last 15 years.

First, the Mexican crisis that occurred after December 1994 triggered a sudden stop of capital flows to other Latin American countries, including Brazil and Argentina (see Sachs *et al.* (1996)), causing the sovereign spread to increase sharply, and some tension on the foreign exchange market. Baig and Goldfajn (2001) report that at the end of 1994 and the beginning of 1995 the net flow of capital to Brazil was insufficient to finance its current account, while the loss of official reserves reached slightly less than $10 billion. The second episode occurred in August and September 1998, after the default of Russia on its external debt. The widespread loss of investors' confidence on emerging markets' assets induced withdrawal of capital from Brazil along with conspicuous loss of foreign exchange reserves. This led the authorities to let the domestic currency float in January 1999 (see Baig and Goldfajn (2001)). The third period of turmoil for Brazilian financial markets goes from the second half of 2001 to the end of 2003, with a short recovery at the beginning of 2002 and a peak in October 2002. The interpretations of these events are debatable. For example, on one hand Goretti (2005) attributes most of the crisis effect to the events unfolding in Argentina from the end of 2001, and thus to regional contagion, with only a limited role for political events in Brazil itself. On the other hand, Miller *et al.* (2005) focus on the role of political uncertainty arising from the fears for unilateral debt restructuring as the left-wing candidate was the favourite in the months preceding the presidential elections of 2002. Miller *et al.* (2005) show that political events and contagion are not mutually exclusive, but, rather, the former can act as a channel of transmission for the latter.

4.1 Data issues

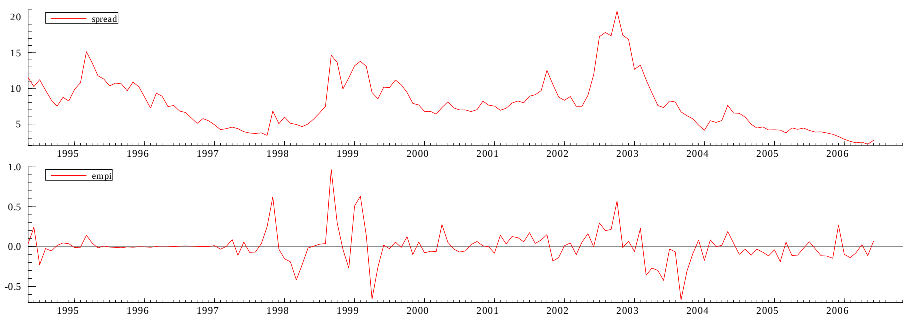
I proxy the risk premium, ρ_t, required by international investors on emerging markets' assets using the Emerging Markets Bond Index + (EMBI+) developed by JP Morgan. The EMBI+ is the spread over the US Treasury interest rate of comparable features of traded external debt instruments in the emerging markets. Specifically, the instruments include external currency-denominated Brady bonds, loans and Eurobonds, as well as US dollar-denominated local markets instruments. Instruments in the EMBI+ must satisfy certain inclusion criteria, e.g. have a minimum face value of 500 million. A country's spread is then calculated as the average of the spreads of all the bonds that satisfy the inclusion criteria, weighted by the market capitalization of the instruments.

In this paper, I term "currency crisis" a strong speculative pressure on the country's currency. Therefore, the probability of devaluation, π_t, is proxied by an index of the pressure exerted by participants on the foreign exchange market, namely the Exchange Market Pressure Index (EMPI). Indices such as this have been widely used in the literature on currency and financial crises (see, among others, Eichengreen *et al.* (1996), Kaminsky and Reinhart (1999) and Cerra and Saxena (2002)) since it does not only capture successful speculative attacks but also unsuccessful ones, i.e. exceptional pressure on the foreign exchange market which did not result in crises because of the effective defence by the monetary authorities. The index is computed as a weighted average of the change in foreign exchange reserves, the nominal exchange rate depreciation, and the change of the Brazilian - US interest rate differential, with weights such as the three components of the index have equal sample volatilities, according to the following formula:

$$empi_t = -\frac{\sigma_e}{\sigma_{RES}}\Delta RES_t + \Delta e_t + \frac{\sigma_e}{\sigma_{id}}\Delta id_t$$

where RES_t denotes foreign exchange reserves, e_t is the nominal exchange rate, id_t denotes the interest rate differential, and σ_i, with $i = RES$, e, id, is the standard deviation. Since large currency devaluations may be driven by hyperinflation, the computation of the EMPI uses different standard deviation weights in different sub-samples according to the magnitude of inflation. In particular, different standard deviations are used for sub-samples with 1, 2, or 3 digits inflation rates. Plotting the EMBI+ and the EMPI (Figure 1) confirms the periods of financial distress highlighted by the narrative evidence reported at the beginning of this Subsection.The exogenous determinants are selected

Figure 1: Brazil's spreads and empi



according to the theoretical model described in Section 2 and the vast literature on the fundamental determinants of sovereign spreads and currency crises (see, e.g., Kaminsky and Reinhart (1999), and Ferrucci (2003)).

The theoretical model suggests that the rate of unemployment, u_t, the government deficit, Δd_t, the change in the trade balance, Δtb_t, and the real exchange rate, $reer_t$, are potential determinants of a currency crises as they affect the net benefit of maintaining a peg through the government loss function. The risk-free real interest rate, r_t^{*f}, on the other hand, affects the probability of devaluation through the asset pricing equation

However, this is only one possible characterization of the government preferences, and other arguments may be included basing on previous literature (e.g. Cerra and Saxena (2002), Fratzscher (2003)). Therefore, in the empirical specification I also include the following explanatory variables: the rate of change of the stock of foreign exchange reserves, Δres_t,[3] the rate of inflation, Δp_t, the current account balance, ca_t, the rate of change of a broad money aggregate, Δm_t, the spread between the bank lending and deposit rates, ld_t, the rate of change of the amount of domestic credit, $\Delta cred_t$, the change in the local stock market index, Δeq_t. A time trend, t, is also included as a short cut to capture the effect of government reputation. A detailed analysis of variables source and construction is given in the Appendix.

4.2 Estimation results

Given the focus on the short-run dynamic relationship of spreads with its possible determinants, all non-stationary variables, except *spread* itself, enter the regression in first difference. Unit root Augmented Dickey-Fueller (ADF) and Phillips-Perron (PP) tests results are reported in Table 1.

Firstly, I estimate by ordinary least squares two single equation linear models (indicated as model (1) and model (2)) of the Brazilian spread as a function of fundamental factors, the second model differing from the first one for the inclusion of *empi* among regressors. Model (3) is a single equation model of the spreads with two regimes estimated with the EM algorithm.[4] All regressors of the single equation models are lagged once in order to avoid any issues of endogeneity. The estimation results of the three single equation models are reported in Table 2.

From the estimation of model (1) (first column of Table 2) we see that the only two fundamental variables affecting Brazilian spreads are the real effective exchange rate, *reer*, and the stock market index, Δeq. The latter becomes statistically insignificant once I include the *empi* among the regressors (model (2) shown in the second column of Table 2). This probably means that the expectations on the exchange market index encompass those on the stock market as a determinant of sovereign spreads. The model (2) performs only slightly better than (1), as evidenced by the standard measures of fit: the standard error σ is lower, as well as the statistics of the standard information criteria, i.e. Akaike (AIC), Hannan-Quinn (HQ), and Schwartz (SIC).[5] A standard F test reveals that the increase in R^2 from 0.311 to 0.330 due to the inclusion of *empi* as a regressor is insignificant as the value of the test statistics, 0.0002, is

[3] The variable RES used in the computation of the $EMPI$ differs from the variable res since the latter is obtained by dividing RES by the GDP volume. See the Appendix for details.

[4] Using again the terminology of Krolzig (1997) this is a MSI(2)-ARX(0).

[5] Standard information criteria are supposed to trade off in-sample fit with prediction accuracy by penalizing models with a large number of parameters. Better models ought to minimize each of the information criteria. These statistics do not suffer from issues related to nuisance parameters and can therefore be used to compare models with different as well as same number of regimes.

Table 1: Unit root test statistics

	ADF (AIC)	ADF (SIC)	PP
spread	-2.49	-2.49	-2.73
Δspread	-11.38	-11.38	-11.38
empi	-4.62 (i)	-8.68 (i)	-8.55 (i)
Δempi	-6.65	-12.91	-39.68
u	-1.86	-1.72	-1.71
Δu	-2.77	-9.67	-10.71
Δd	-1.88	-1.88	-11.05
Δ^2d	-12.19	-12.19	-69.64
Δtb	-4.30 (i)	-17.36 (i)	-17.76 (i)
Δ^2tb	-8.34	-8.34	-73.40
reer	-4.40 (i)	-4.75 (i)	-3.76 (i)
Δreer	-9.03	-9.03	-8.44
rf	-0.69 (i)	-8.09 (i)	-7.67 (i)
Δrf	-5.32	-7.60	-27.99
Δp	-3.19 (i)	-57.90 (i)	-30.44 (i)
Δ^2p	-5.95	-14.24	-229.50
ip	-3.11	-3.11	-3.05
Δip	-13.98	-13.98	-13.98
ca	-0.64 (i)	-1.62 (i)	-1.71 (i)
Δca	-3.52	-5.64	-5.98
Δm	-6.31 (i)	-13.21 (i)	-12.95 (i)
Δ^2m	-7.42	-7.42	-32.36
ld	-4.58 (i)	-5.15 (i)	-5.31 (i)
Δld	-6.33	-12.73	-15.16
Δcred	-14.64 (i)	-14.64 (i)	-15.61 (i)
Δ^2cred	-4.93	-8.56	-25.06
Δres	-6.19 (i)	-11.94 (i)	-11.94 (i)
Δ^2res	-7.01	-12.67	-50.02
Δeq	-11.80 (i)	-11.80 (i)	-11.88 (i)
Δ^2eq	-4.63	-9.40	-42.96

Notes: The sample period is 1994:5-2006:3. AIC (SIC) indicates Akaike (Schwarz) Information criterion. The regressions for all the level variables include a trend and an intercept, except where the symbol (i) appears, indicating that the only intercept is included. The regressions for variables in first difference include only an intercept. The 95% critical value for regressions with trend is -3.44 and for regressions without trend -2.88.

Table 2: Univariate models estimation results

	model (1)	model (2)	model (3)
ν(1)	**13.102**	**11.775**	2.97
	(2.954)	(3.010)	(2.129)
ν(2)	-	-	**7.913**
			(2.060)
empi (-1)	-	*3.388*	**4.732**
		(1.809)	(1.274)
Δu (-1)	-0.628	-0.623	**-0.867**
	(0.672)	(0.666)	(0.4306)
Δd (-1)	56.034	52.02	-4.244
	(39.46)	(39.13)	(27.738)
Δtb (-1)	7.314	11.944	1.236
	(73.18)	(72.51)	(46.767)
reer (-1)	**-14.861**	**-12.231**	**-6.102**
	(3.426)	(3.672)	(2.782)
rf (-1)	-3.809	0.202	54.946
	(101.0)	(100.1)	(67.811)
Δp (-1)	-12.318	-6.392	7.328
	(19.21)	(19.28)	(12.257)
Δip (-1)	-0.153	-0.098	0.05
	(0.141)	(0.143)	(0.096)
Δca (-1)	*85.654*	*80.731*	**71.352**
	(49.03)	(48.62)	(33.263)
Δm (-1)	-0.939	-5.706	**-23.789**
	(15.90)	(15.95)	(10.400)
ld (-1)	-1.081	-0.731	0.966
	(0.824)	(0.837)	(0.593)
Δcred (-1)	16.425	16.907	**18.877**
	(12.05)	(11.93)	(7.863)
Δres (-1)	-5.063	5.463	**17.903**
	(9.652)	(11.09)	(7.307)
Δeq (-1)	**-0.002**	*-0.002*	-0.001
	(0.001)	(0.001)	(0.001)
trend	**-0.025**	**-0.023**	-0.001
	(0.008)	(0.008)	(0.006)
σ	3.083	3.053	1.904
AIC	5.189	5.176	4.628
HQ	5.316	5.311	4.797
SIC	5.501	5.509	5.044
Log-lik	-353.427	-351.476	-308.583
LR linearity test statitstic			85.787
DAVIES			[0.000]**

Notes: Coefficients in bold are significant at the 5% significance level. Standard errors of estimated coefficients are in parentheses.

lower than the critical value F^1_{126}=3.92. Model (3) allows for two unobserved regimes in the underlying stochastic process. This two-state model is more satisfactory than its linear counterpart, i.e. model (2), by any measures of fit: the standard error and the information criteria statistics are all lower than for model (2), while the log-likelihood is much higher. However, Garcia (1998) showed that a formal testing procedure of the regime switching hypothesis - i.e. of the null hypothesis of $k = 1$, where k is the number of regimes, against the alternative of $k \geq 2$ - via likelihood ratio (LR) tests can be tricky as under the null a few parameters of the unrestricted model (under the alternative) - i.e. the elements of the transition probability matrix associated to the rows that correspond to "disappearing states" - can take any values without influencing the likelihood function. In presence of such parameters, said to be "nuisance" to the estimation, the LR statistics fails to have a standard chi-square distribution even asymptotically. However, Davies (1977) derives an upper bound for the significance level of the likelihood ratio test statistic under nuisance parameters. Table 2 reports the p-value of such test showing that the null hypothesis of linearity is rejected and the univariate Markov-switching model outperforms its linear counterpart. The better performance is also highlighted by the fact that now the exchange market pressure index, *empi*, is statistically significant, as well as a number of fundamental variables such as the unemployment rate, the real exchange rate, the current account, the monetary aggregate, the domestic credit expansion, and the foreign exchange reserves. The good performance of model (3) is also highlighted by Figure 2, showing the fitted and actual values of the dependent variable.

The main problem with model (3) is that it implies a lagged relationship between the Brazilian risk premium, indicated by *spread*, and the probability of devaluation, as proxied by *empi*. Therefore the latter is treated as an exogenous determinant of the country risk premium. The theoretical model of Section 2, on the other hand, clearly implies a contemporaneous relationship as well as feedback effects between these two variables which should thus be determined endogenously. As shown in Section 3, this relationship is better represented empirically by a VAR approach. I now turn to show that this specification is also supported by the Brazilian data.

The MSI(2)-VARX(p, q) model (18) is estimated over the period 1994:1-2005:12. The orders p and q are determined according to the Akaike (AIC), Hannan-Quinn (HQ), and Schwartz Bayesian (SBC) criteria among models with two regimes. Given the large number of parameters to be estimated and the available sample size, I limit the combinations to p, $q = 0,1,2$. All criteria select a MSI(2)-VARX(1,1) model (see Table 3), which is then estimated with the maximum likelihood method through the implementation of the EM algorithm.

Table 4 reports the estimated coefficients as well as measures of fit of the MSI(2)-VARX(1,1) model.

The estimated matrix of transition probabilities shows that two fairly per-

Figure 2: Univariate MS model: fitted and actual values of Brazil's spread

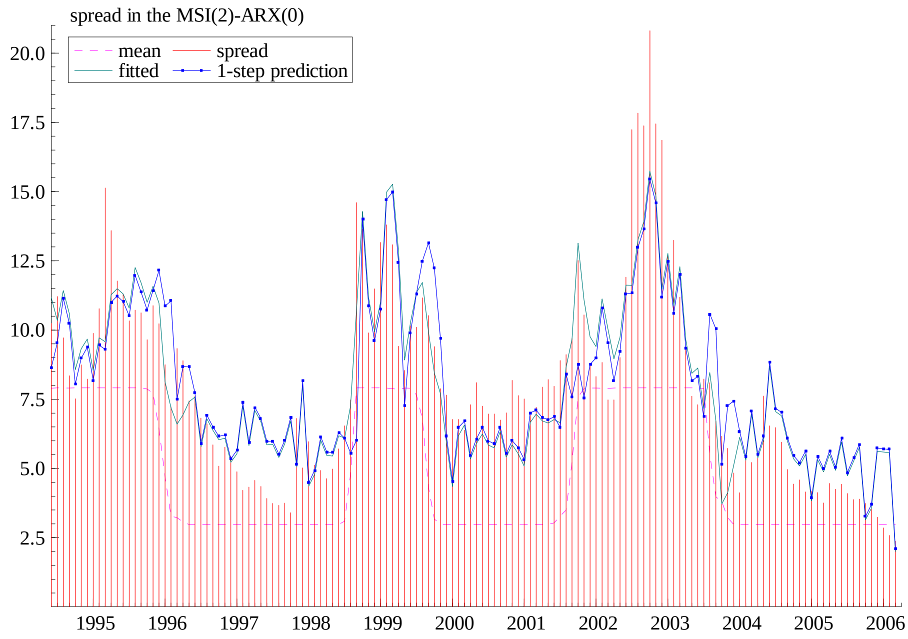


Table 3: Test statistics for selecting the lag order of the endogenous and exogenous variables in the MSI(2)-VARX(p,q) model

Argentina			
Order (pi,qi)	AIC	HQ	SBC
(1,1)	**4.3043**	**4.8515**	**5.6510**
(1,0)	4.8961	5.2244	5.7041
(1,2)	4.3562	5.1259	6.2504
(2,1)	4.4100	4.9936	5.8463
(2,0)	4.8314	5.1951	5.7265
(2,2)	4.5627	5.3663	6.5402

Notes: statistics in bold indicate the order selected by the relevant criterion/test. AIC, HQ, and SBC denote the Akaike, Hannan-Quinn, and Schwarz Bayesian Information criterion, respectively.

Table 4: MSI(2)-VARX(1,1) estimation results

	spread	empi
ν(1)	**3.84**	**0.618**
	(1.078)	(0.139)
ν(2)	**5.636**	**0.836**
	(1.201)	(0.152)
spread (-1)	**0.752**	**-0.023**
	(0.036)	(0.005)
empi (-1)	**-1.184**	**0.147**
	(0.480)	(0.063)
Δu	0.285	-0.018
	(0.212)	(0.028)
Δd	**25.006**	1.036
	(12.372)	(1.621)
Δtb	-22.917	0.219
	(21.725)	(2.862)
reer	**-8.922**	**-1.484**
	(1.470)	(0.194)
rf	-1.775	-1.172
	(30.899)	(4.046)
Δp	**-27.182**	-1.020
	(5.975)	(0.788)
Δip	-0.044	*-0.010*
	(0.043)	(0.006)
Δca	24.572	1.412
	(15.226)	(1.990)
Δm	**10.086**	0.341
	(4.961)	(0.655)
ld	**-0.659**	**-0.131**
	(0.271)	(0.035)
Δcred	**11.902**	*0.822*
	(3.807)	(0.495)
Δres	**-12.081**	**-3.191**
	(2.892)	(0.382)
Δeq	**-0.002**	**-0.000**
	(0.000)	(0.000)
trend	*-0.005*	**-0.001**
	(0.003)	(0.000)
σ	0.883	0.117
AIC	2.014 [2.105]	
HQ	2.361 [2.418]	
SIC	2.867 [2.876]	
Log-lik	-101.992 [-112.481]	
LR linearity test statitstic	20.977	
DAVIES	[0.0006]**	

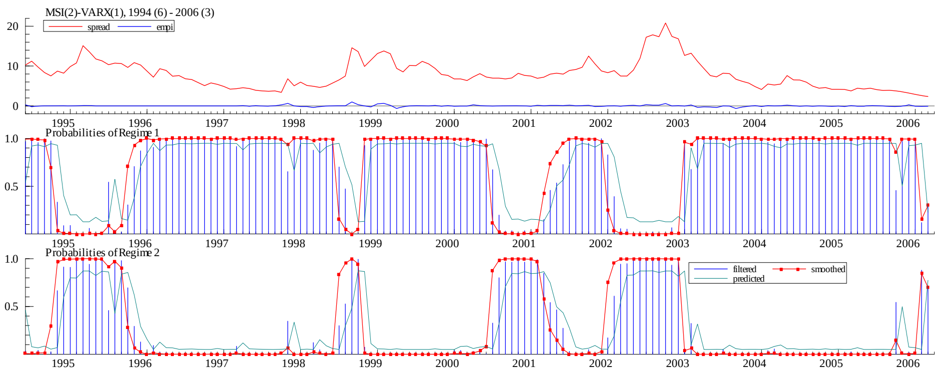
Notes: Coefficients in bold (italics) are significant at the 5% (10%) significance level. Standard errors are in parentheses. For the information criteria and the Log-likelihood, the value in square brackets is the corresponding statistics of the linear model.

sistent regimes do exist:

$$\widehat{P}=\begin{pmatrix} 0.9489 & 0.0511 \\ 0.1278 & 0.8722 \end{pmatrix}$$

Figure 3 exhibits the filtered, smoothed, and predicted probabilities of being in Regime 1 or Regime 2. As evidenced by the intercept coefficients of the estimated model, Regime 1 corresponds to "tranquil" periods, i.e. periods characterized by the absence of financial distress, while Regimes 2 represents times of high economic tension during which the risk premium on domestic assets and the probability of currency devaluation (or pressure on exchange rate market) are considerably higher. Regime 2 corresponds to the crises that have hit the Brazilian economy during the '90s and the beginning of the following decade. Specifically, the crises that seem to have been relevant are the Mexican devaluation of 1994-1995, the Russian default of 1998, the Argentinean turmoil of 2001-2002,[6] and, finally, the pressures on the Brazilian currency surrounding the Presidential elections of 2002.The model diagnostic analysis, conducted through

Figure 3: Regime probabilities



the analysis of plots of standardized and prediction errors reported in Figure 4, shows that the absence of severe problems of autocorrelation or departures from normality.Most of the fundamental variables coefficients (8 out of 13) are statistically significant in either one of the two equations of the model, and they have the right signs with the possible exception of the rate of inflation whose coefficient is expected to be positive. The model performs better than a linear, i.e. one state, VAR(1) featuring the same specification as can be seen from the lower values of information criteria statistics (in square brackets are the linear VAR statistics) and a significantly higher log-likelihood, as evidenced by the Davies (1977) test's p-value. Figure 5 shows the remarkable performance of the MS-VAR model by reporting the fitted and actual values of *spread* and *empi*.Thus,

[6] In such case, however, Boschi (2005) suggests that contagion, defined as a significant

Figure 4: Estimation errors diagnostic analysis

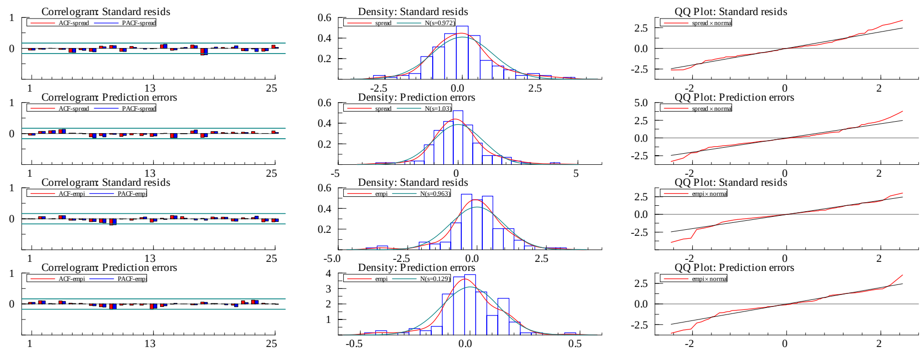
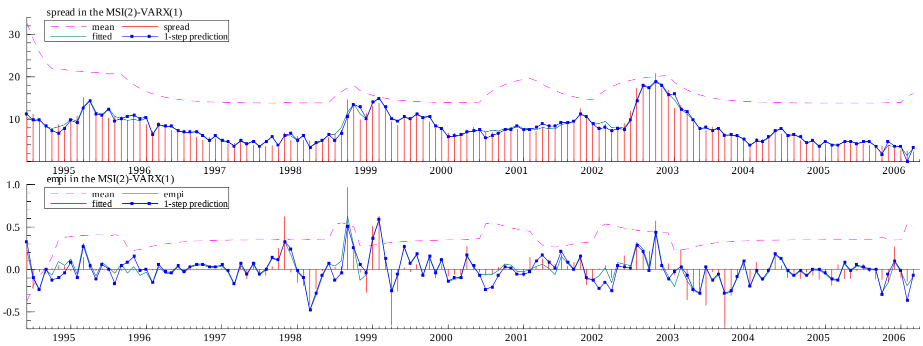


Figure 5: MSI(2)-VAR(1) model: fitted and actual values of Brazil's *spreads* and *empi*



the estimated MS-VAR model derived as an empirical specification of the theoretical framework discussed in Section 2 fit the data remarkably well. This supports the hypothesis that the relationship between sovereign bond spreads and their fundamental economic determinants is non-linear due to the effect of market expectations on currency devaluation. This helps to understand why the previous literature failed to recover a linear relationship between spreads and fundamentals.

5 Conclusion

In this paper, I show how the model of Boschi and Goenka (2007) can be used to analyse the determinants of sovereign debt interest rates spreads in emerging markets. Earlier studies highlight that over crisis periods the relationship between sovereign spreads and macroeconomic fundamentals breaks down, suggesting market inefficiency. The model of Section 2, by contrast, shows that a non-linear relationship between emerging markets spreads and fundamentals can be the outcome of investors' rational behaviour as they take into account the probability of domestic currency devaluations. The paper tests this hypothesis by developing an empirical version of the theoretical framework based on a Markov-switching regime multivariate model of spreads and an index of market pressure on foreign exchange. The empirical illustration uses Brazilian data.

The results show that the model fits the data very well, and in this respect it outperforms a linear version of the multivariate model, as well as univariate specifications, both linear and non-linear, of sovereign spreads.

6 Appendix

6.1 Derivation of equations (15) and (16)

Following Jeanne and Masson (2000), we assume that ϕ_t, i.e. the expected value of economic fundamentals given by equation (11), is a linear index aggregating the macroeconomic variables that are most relevant for the policy maker's choice of maintaining or not the fixed peg, plus a shock, i.e.:

$$\phi_t = \boldsymbol{\alpha}'\mathbf{x}_t + \eta_t \tag{23}$$

where $\boldsymbol{\alpha} = (\alpha_1, ..., \alpha_K)'$ is a vector of coefficients, $\mathbf{x}_t = (x_{1t}, ..., x_{Kt})'$ is a vector of relevant economic fundamentals, and η_t is an i.i.d. shock representing other exogenous determinants of the policy maker's behaviour.

In order to linearize equation (14), I assume that the fluctuations of the risk premium, the expected value of fundamentals, and the thresholds level of fundamental triggering devaluations are small, i.e:

$$\rho_t = \overline{\rho} + \delta\rho_t \tag{24}$$

increase in cross-country correlation of relevant financial markets' indices, did not occur.

$$\phi_t = \overline{\phi} + \delta\phi_t \tag{25}$$

$$\phi^*_s = \phi^* + \delta\phi^*_s \tag{26}$$

where $\delta\rho_t$, $\delta\phi_t$, and $\delta\phi^*_s$ are of the first order. Than, a first order Taylor expansion of equation (14) at the points $\overline{\rho}$, $\overline{\phi}$, and ϕ^* gives the equation:

$$\pi_t = F[\overline{\rho}, \overline{\phi}, \phi^*] + F_1[\overline{\rho}, \overline{\phi}, \phi^*]\delta\rho_t + F_2[\overline{\rho}, \overline{\phi}, \phi^*]\delta\phi_t + F_3[\overline{\rho}, \overline{\phi}, \phi^*]\sum_{s=1}^{n}\theta(s_t, s)\delta\phi^*_s \tag{27}$$

where F_i, for $i = 1, 2, 3$ is the partial derivative of function F with respect to its i-th argument.

Then, reminding that from equation (25)

$$\delta = 1 - \frac{\overline{\phi}}{\phi_t}$$

equation (27) can be manipulated to yield equation (15):

$$\pi_t = \gamma_{s_t} + \varphi_1\rho_t + \boldsymbol{\psi}'_1\mathbf{x}_t + v_t$$

where:

$$\gamma_{s_t} = F[\overline{\rho}, \overline{\phi}, \phi^*] - F_2[\overline{\rho}, \overline{\phi}, \phi^*]\overline{\phi} + F_3[\overline{\rho}, \overline{\phi}, \phi^*]\sum_{s=1}^{n}\theta(s_t, s)\delta\phi^*_s,$$

$$\varphi_1 = F_1[\overline{\rho}, \overline{\phi}, \phi^*]\delta,$$

$$\boldsymbol{\psi}'_1 = F_2[\overline{\rho}, \overline{\phi}, \phi^*]\boldsymbol{\alpha}',$$

and

$$v_t = F_2[\overline{\rho}, \overline{\phi}, \phi^*]\eta_t.$$

In order to obtain equation (16), consider that the model discussed in Section 2 implies the following general expression for the risk premium:

$$\rho_t = G(\pi_t, \phi_t, \phi^*_s). \tag{28}$$

Assume that

$$\pi_t = \overline{\pi} + \delta\pi_t \tag{29}$$

where $\delta\pi_t$ is of the first order. Then, by taking a first order Taylor approximation at the points $\overline{\pi}$, $\overline{\phi}$, and ϕ^*, I obtain:

$$\rho_t = G(\overline{\pi}, \overline{\phi}, \phi^*) + G_1(\overline{\pi}, \overline{\phi}, \phi^*)\delta\pi_t + G_2(\overline{\pi}, \overline{\phi}, \phi^*)\delta\phi_t + G_3(\overline{\pi}, \overline{\phi}, \phi^*)\sum_{s=1}^{n}\theta(s_t, s)\delta\phi^*_s \tag{30}$$

where G_i, for $i = 1, 2, 3$ is the partial derivative of function G with respect to its i-th argument. By setting $\overline{\pi} = 0$, i.e. the initial probability of devaluation is expected to be null, and rearranging equation (30), I obtain equation (16):

$$\rho_t = \mu_{s_t} + \varphi_2\pi_t + \boldsymbol{\psi}'_2\mathbf{x}_t + e_t$$

where:

$$\mu_{s_t} = G(\overline{\pi}, \overline{\phi}, \phi^*) - G_2(\overline{\pi}, \overline{\phi}, \phi^*)\overline{\phi} + G_3(\overline{\pi}, \overline{\phi}, \phi^*)\sum_{s=1}^{n}\theta(s_t, s)\delta\phi^*_s,$$

$$\varphi_2 = G_1(\overline{\pi}, \overline{\phi}, \phi^*)\delta,$$
$$\boldsymbol{\psi}_2' = G_2(\overline{\pi}, \overline{\phi}, \phi^*)\boldsymbol{\alpha}',$$

and

$$e_t = G_2(\overline{\pi}, \overline{\phi}, \phi^*)\eta_t.$$

6.2 Data sources and variables construction

All data, except the spreads, the real effective exchange rate, and the stock market index, are obtained from the International Financial Statistics database published by the IMF. Codes are provided below.

Spreads (*spreads*) The source of the EMBI+ time series is DATASTREAM.

Exchange Market Pressure Index (*empi*) See Section 4.1.

Reserves (RES and res) RES is the Total Reserves minus Gold (code: .1L.DZF). The series res is obtained by dividing RES by the volume of GDP in millions of US dollars (code: 99B..ZF - quarterly data are interpolated using the rates of growth of industrial production).

Nominal exchange rate (e) Series of National Currency per US Dollar (code: ..RF.ZF...).

Interest rate differential (id) It is obtained by subtracting the US Federal Funds rate (code: 60B..ZF) to the Brazilian Money Market rate (60B..ZF).

Unemployment rate (u) Code: 67R..ZF. Missing values are linearly interpolated from annual data.

Government deficit (Δd) The local currency-denominated Deficit or Surplus (80...ZF) is first converted to US dollars using the series e and then divided by the volume of GDP.

Trade balance (tb) The series is given by the difference between Exports (70..DZF) and Imports (71.VDZF) divided by the volume of GDP.

Real effective exchange rate ($reer$) The original series, an index (2000=100) constructed by JP Morgan, is detrended using the Hodrick and Prescott (HP) filter.

Risk-free real interest rate (r^{*f}) Let R^{US} be the US Federal Funds rate (60B..ZF) and P^{US} the US Consumer price index (64...ZF). Then the series r^{*f} is given by the following formula:

$$r^{*f} = (1/12) \cdot \ln(1 + R^{US}/100) - \Delta \ln(P^{US})$$

Rate of inflation (Δp) Let P be the Brazilian Consumer price index (64...ZF), then $\Delta p = \Delta \ln(P)$.

Current account balance (ca) The series code is 78ALDZF. The annual data are linearly interpolated and then divided by the volume of GDP.

Money aggregate (m) Sum of the series Money (34...ZF) and Quasy money (35...ZF).

Bank lending and deposit rate spread (ld) The series is given by the ratio of the bank lending (60P..ZF) and the deposit rate (60L..ZF).

Domestic credit ($cred$) The series is the Brazilian Domestic credit (52...ZF).

Stock market index (eq) The series is the Brazil Bovespa index provided by DATASTREAM.

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