On the dynamics of lending and deposit interest rates in emerging markets: a non-linear approach

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Abstract

This paper studies the dynamics of lending and deposit rates in two emerging markets in Latin America: Colombia and Mexico. The dynamics of lending (deposit) interest rates are driven by the exogenous interbank interest rate and deviations from the long-run lending-interbank (deposit-interbank) interest rate relationship. Allowing for different interest rate behavior during periods characterized by large and small values of the spread, the non-linear specification proves superior to the linear one.

Keywords: Interest rates; spreads; emerging markets, non-linear models, regimes

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1. Introduction

The financial sector plays a crucial role in the operation of most economies, as it provides intermediation between borrowers and lenders of funds. To the extent that financial intermediaries are efficient institutions for channeling funds from savers to borrowers, they can affect investment decisions and economic growth.

The theoretical framework of Hannan and Berger (1991), Neumark and Sharpe (1992) and Freixas and Rochet (1997) provides some insights on the determination of deposit and lending interest rates. These are “retail” interest rates affected by the supply of deposits and demand for loans, and (within an imperfectly competitive banking sector) also by an exogenous inter-bank or money market (“wholesale”) rate. The theory suggests that both the lending and deposit interest rates maintain a stable long-run equilibrium relationship with the interbank rate, in the sense that these variables exhibit a systematic co-movement over time. In the short run, however, equilibrium may fail to hold because economies constantly experience shocks and other disturbances, although economic forces do not allow for these short-run deviations from equilibrium to grow indefinitely over time.

An interesting question that arises is that of the type of adjustment back to equilibrium, and in particular the possibility of linear versus non-linear type of adjustments. Using US cross-sectional studies, Hannan and Berger (1991) and Neumark and Sharpe (1992) provide several reasons for asymmetric adjustments of deposit and lending rates such as (i) collusive pricing arrangements and (ii) adverse customer reaction. First, consider the collusion hypothesis. In concentrated markets, banks find it costly to deviate from collusive pricing behavior. The costs are higher (a) for lending rate decreases because of lower payments received by borrowers, and (b) for deposit rate increases because of extra payments that have to be made to depositors. Therefore, the collusion hypothesis points to greater rigidity for lending rate decreases and deposit rate increases. Next, consider the adverse customer reaction hypothesis. Banks may be concerned about negative reactions from customers following increases of lending rates and decreases of deposit rates. If this is the case, lending (deposit) rates should be upwards (downwards) rigid. Amongst the most recent studies, Scholnick (1996a,b) fails to reject the hypothesis of collusion for the emerging markets of Malaysia and Singapore. Frost and Bowden (1999) find asymmetries in the relationship between mortgage and bank bill rates for New Zealand, whereas Lim (2001) finds rigidities both in lending and deposit rate increases for Australia.

Our aim is to test for and model non-linearities in the lending and deposit interest rates in two emerging market economies: Colombia and Mexico. We characterize the behavior of the interest rates using the Smooth Transition Autoregressive model (thereafter STAR; see e.g. Teräsvirta and Anderson, 1992; Granger and Teräsvirta, 1993). In a recent survey of STAR models and their applications, van Dijk *et al*. (2002) point out that this form of non-linear models has mainly been applied to macroeconomic time series and only recently to long and short-term interest rates in developed countries (see e.g. Anderson, 1997 for the US; van Dijk and Franses, 2000 for the Netherlands). The STAR model can be interpreted as a regime-switching model, where the transition from one regime to the other occurs in a smooth way. Assuming that the transition mechanism is controlled by the interest rate spread, we can differentiate between the impact of the spread on lending and deposit interest rates during periods of inefficiency in banking activities and/or periods of financial crises (when the spread is too high), and its impact on lending and deposit rates during periods of increasing banking competition and/or "normal" times (when the spread is low).[1] Furthermore, we can identify threshold levels for the spread rate that mark the transition from one regime to the other, as well as the speed at which this transition takes place.

The outline of the paper is as follows. Section 2 provides a historical background to the

[1] Another reason that may help explain high interest rate spreads is the extent of deposit account dollarization which tends to increase during financial crises. On the positive association between spreads and dollarization see Honohan and Shi (2002).

behavior of interest rates in Latin America. Section 3 introduces the theoretical aspects of non-linear models in the context of the STAR methodology. Section 4 estimates linear and non-linear models for the lending and deposit rates in Latin America. Section 5 presents a discussion of our findings and section 6 provides some concluding remarks.

2. Historical context

For decades, Latin American economies pursued inward-looking development strategies in which government intervention was predominant. Consequently, these economies were characterized by the use of trade barriers and foreign exchange controls to protect indigenous infant industries against foreign competition, and by heavily controlled financial systems that resulted in financial repression. Financial repression is a term that refers to a policy regime in which high reserve requirements are imposed on financial intermediaries as well as ceilings on their deposit and lending interest rates. In addition to these features, there are restrictions on competition in the banking industry and on the composition of bank portfolios. The former takes the form of entry barriers into the banking system and public ownership of financial institutions; the latter consists of the operation of non-price mechanisms of credit allocation in the form of directed lending to specific productive sectors (Agénor and Montiel, 1996).

During the 1980s, Mexico and Colombia saw considerable financial repression. Following the debt crisis of 1982 in Mexico, all Mexican banks were nationalized, and the government imposed high reserve requirements, set ceilings on interest rates, and directed lending to specific "high priority" productive sectors (Saunders and Schumacher, 2000). Colombia entered the 1980s facing the collapse of coffee prices (i.e. the country's main export product and an important determinant of its business cycle), along with a deteriorating situation of government finances. Montenegro (1983) argues that the financial crisis of the early 1980s can be explained, to a great extent, by this economic downturn. The crisis hit strongly poorly capitalized banks as well as small banks, all of which suffered from loan portfolios concentrated on unprofitable firms often belonging to the owners of the banks. This last aspect also reflects a system where the operations of financial intermediaries were not properly supervised and regulated by the authorities.[2]

During the 1990s, Latin American economies adopted policy reforms aimed at providing a transition to a more liberalized domestic financial sector. Policy reforms in Colombia and Mexico included measures to ease entry for new intermediaries, simplify legal reserve requirements and accelerate the privatization of government-controlled firms (see e.g. Barajas *et. al*., 1999; Saunders and Schumacher, 2000). These liberalizing efforts appear to have had a different impact on the market structure of the two countries. Indeed, in 1990, the banking systems in Colombia and Mexico were highly concentrated, as the top three banks held about 56 percent and 66 percent of total assets, respectively. Between 1997 and 2001, these figures stood at an average 29 percent for Colombia and almost twice as much (56 percent) for Mexico.[3]

The liberalization process was tested in two occasions. The first one during the Mexican financial turmoil of 1994-1995 and the second one during the eruption of the severe financial crisis that hit the Asian economies in mid 1997 followed by global economic uncertainties in response to the Russian moratorium in mid 1998. As Brock and Suárez-Rojas (2000) point out, Latin American authorities responded to the financial crises of the 1990s by intensifying their efforts for deeper reforms, rather than turning back to government intervention policies that followed the failure of liberalization measures in the second half of the 1970s.

The next section of the paper discusses the theory of regime-switching models in the context of the STAR methodology that will be empirically tested on the behavior of lending and deposit

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[2] In Colombia, banks were also subject to high rates of financial taxation, which provided an additional factor of financial repression. A description of the institutional background in the Colombian financial sector can be found in Barajas *et al*. (1999, 2000).

[3] Similar figures are obtained based on the deposits held by the top three banks. Concentration figures for 1990 were taken from the Financial Structure Database and assembled by Beck *et al*. (2001). Figures for the 1997-2001 period are based on the authors' calculations using data from the Superintendencia Bancaria of Colombia and the Comisión Nacional Bancaria y de Valores of Mexico.

interest rates in the Latin American economies.

3. Specification of STAR models

The STAR model of order k for a univariate time series y_t is written as:

$$y_t = \left(\mu_1 + \sum_{j=1}^{k} \beta_{1,j} y_{t-j}\right)(1 - G(s_{t-d})) + \left(\mu_2 + \sum_{j=1}^{k} \beta_{2,j} y_{t-j}\right) G(s_{t-d}) + \varepsilon_t, t = 1,...,T, \quad (3.1)$$

where μ_1 and μ_2 are intercept terms and $\varepsilon_t \sim iid\ (0, \sigma^2)$. $G(s_{t-d})$ is the transition function, which is assumed to be continuous and bounded between zero and one, and d is the delay parameter. The STAR model (1) can be considered as a regime-switching model which allows for two regimes, $G(s_{t-d}) = 0$ and $G(s_{t-d}) = 1$, respectively, where the transition from one to the other regime occurs in a smooth way. The regime that occurs at time t is determined by the transition variable s_{t-d} and the corresponding value of $G(s_{t-d})$. Different functional forms of $G(s_{t-d})$ allow for different types of regime-switching behavior. In particular, asymmetric adjustment to positive and negative deviations of s_{t-d} relative to a parameter c, can be obtained by setting $G(s_{t-d})$ equal to the 'logistic' function:

$$G(s_{t-d}; \gamma, c) = \{1 + \exp[-\gamma(s_{t-d} - c) / \sigma(s_{t-d})]\}^{-1}, \gamma > 0, \quad (3.2a)$$

where $\sigma(s_{t-d})$ is the sample standard deviation of s_{t-d}. The parameter c is the threshold between the two regimes, in the sense that $G(s_{t-d})$ changes monotonically from 0 to 1 as s_{t-d} increases, while $G(s_{t-d}) = 0.5$ when $s_{t-d} = c$. The parameter γ determines the smoothness of the change in the value of the logistic function and thus the speed of the transition from one regime to the other. When $\gamma \rightarrow 0$ the 'logistic' function equals a constant (i.e. 0.5), and when $\gamma \rightarrow \infty$, the transition from $G(s_{t-d}) = 0$ to $G(s_{t-d}) = 1$ is almost instantaneous at $s_{t-d} = c$.

Another type of regime-switching behavior, which describes asymmetric adjustment to small and large absolute values of s_{t-d}, is obtained by setting $G(s_{t-d})$ equal to the 'exponential' function:

$$G(s_{t-d};\gamma,c) = 1-\exp\{-\gamma(s_{t-d}-c)^2/\sigma^2(s_{t-d})\},\gamma > 0.$$

A possible drawback of the 'exponential' function is that the model becomes linear if either $\gamma \to 0$ or $\gamma \to \infty$. To overcome this drawback, Jansen and Teräsvirta (1996) suggest specifying $G(s_{t-d})$ as the 'quadratic logistic' function:

$$G(s_{t-d};\gamma,c_1,c_2) = \{1+\exp[-\gamma(s_{t-d}-c_1)(s_{t-d}-c_2)/\sigma^2(s_{t-d})]\}^{-1},\gamma > 0. \tag{3.2b}$$

In this case, if $\gamma \to 0$, the model becomes linear, whereas if $\gamma \to \infty$, $G(s_{t-d})$ is equal to 1 for $s_{t-d} < c_1$ and $s_{t-d} > c_2$, and equal to 0 in between.

The estimation of STAR models consists of three steps:

Step 1: Specify a linear autoregressive (AR) model as the base one. The model can be extended to allow for other exogenous variables as additional regressors. This is discussed in the next section.

Step 2: Select the transition variable s_{t-d} and test linearity, for different values of the delay parameter d, against STAR models using the linear model specified in *Step* 1 as the null hypothesis. Linearity testing is based on a third-order Taylor approximation of the transition function around $\gamma = 0$. This results in estimating the auxiliary regression:

$$v_t = \mu_0 + \sum_{j=1}^{k}\phi_{0,j}y_{t-j} + \sum_{j=1}^{k}\phi_{1,j}y_{t-j}s_{t-d} + \sum_{j=1}^{k}\phi_{2,j}y_{t-j}s_{t-d}{}^2 + \sum_{j=1}^{k}\phi_{3,j}y_{t-j}s_{t-d}{}^3 + e_t, \tag{3.3}$$

where v_t are the residuals of the linear model of *Step* 1. The null hypothesis of linearity is

$H_0 : \phi_{1,j} = \phi_{2,j} = \phi_{3,j} = 0$, for $j = 1,\ldots, k$. This is a standard Lagrange Multiplier (LM) type test. To specify the value of the delay parameter d, model (3.3) is estimated for a number of different values of d, say $d = 1,\ldots, D$. In cases where linearity is rejected for more than one values of d, the decision rule is to select d based on the lowest p-value of the linearity test.

Step 3: Proceed by selecting the appropriate form of the transition function $G(s_{t-d})$, that is, select between the 'logistic' function (3.2a) and the 'quadratic logistic' function (3.2b). This is done by running a sequence of LM tests nested within the non-linear model (3.3) of *Step* 2, namely:

$$\begin{aligned} &H_{03} : \phi_{3,j} = 0, \\ &H_{02} : \phi_{2,j} = 0 \mid \phi_{3,j} = 0, \\ &H_{01} : \phi_{1,j} = 0 \mid \phi_{3,j} = \phi_{2,j} = 0. \end{aligned} \tag{3.4}$$

In this case, the decision rule is to select the 'quadratic logistic' function (3.2b) if the p-value associated with the H_{02} hypothesis is the smallest one, otherwise select the 'logistic' function (3.2a).[4] Having done that, proceed by estimating the STAR model (3.1), with the transition function $G(s_{t-d})$ specified based on the sequence of tests in (3.4).

4. Empirical results

4.1 The data

We use monthly data on the lending, deposit and inter-bank or money market ("wholesale") interest rates for Colombia and Mexico. The Mexican data set is obtained from the IMF *International Financial Statistics* database and the Colombian dataset comes from the Colombian Banking Association (Asobancaria) and the IMF. Colombian data is from 1989:M1 to 2002:M7

[4] Van Dijk and Franses (2000) motivate the sequence of tests in (3.4) on the grounds that if a 'logistic' alternative is appropriate, then the second order effect in the Taylor expansion should be zero, that is, $\phi_{2,j} = 0$.

and Mexican data is from 1993:M1 to 2002:M7. The sample choice is dictated by the availability of data in the International Monetary Fund's *International Financial Statistics* CD-ROM.

Figure 1 plots the levels of the interest rates for the two emerging market economies. Estimation of linear and non-linear models requires stationarity of the interest rate series. Table 1 reports the Augmented Dickey Fuller (ADF) tests on the levels and the first differences of the interest rates. ADF tests are also reported for the spread between the lending and interbank interest rates and the spread between the deposit and interbank rates. The results suggest that Colombian and Mexican interest rates are non-stationary (i.e. $I(1)$) in levels, whereas the spreads are found to be stationary (i.e. $I(0)$) for both countries. Based on the results of the unit root tests, linear and non-linear models are estimated for the first differences of the interest rates.[5]

In the remaining of the paper we adopt the following notation for the interest rate series in the two emerging markets: *COL_l*, *COL_d* and *COL_ib* refer to the lending, deposit and interbank rates in Colombia. The spread between the lending and interbank rate is denoted by *COL_cvl* whereas *COL_cvd* is the spread between the deposit and interbank rate. *MEX_l*, *MEX_d*, *MEX_ib*, *MEX_cvl* and *MEX_cvd* refer to the corresponding series in Mexico.

4.2 Testing for linearity and STAR model selection

As discussed in section 3, the first step in deriving STAR models involves the estimation of linear interest rate models. The theoretical framework briefly discussed in section 1 suggests that changes in lending (deposit) rates are driven by the exogenous interbank interest rate and deviations from the long-run lending-interbank (deposit-interbank) interest rate relationship. In Table 2, we report unrestricted linear models with lag lengths, k, chosen by the Akaike and Schwarz Bayesian Information Criteria, (AIC and BIC, respectively). The interest rate equations can be interpreted as error correction models; lending and deposit interest rate changes react to

[5] Phillips-Perron tests give similar unit root results and are available by the authors on request.

the disequilibrium errors given by the (lagged) lending-interbank and deposit-interbank rates.[6] The diagnostic tests of the linear models in Table 2 show some evidence of ARCH effects and normality failures. The failure of the diagnostic tests in the linear models provides a further motivation for considering the possibility that the interest rates might be better characterized by a non-linear type of behavior rather than the linear one discussed above.

Having estimated the base linear models, we move on to *Step* 2 of our methodology which involves testing for the existence of non-linear dynamics in the interest rate models for the two Latin American economies selecting the relevant interest rate spread as a possible transition variable s_{t-d}. [7] The empirical results of the LM-type tests for linearity (*Steps* 2 and 3 of section 3) are reported in Table 3. We set d equal to 1 through 6 (although the results are not affected even if we go up to $d = 12$). From Table 3A, the H_0 hypothesis is rejected most strongly both at $d = 1$ and $d = 3$ for the Colombian lending rate equation and at $d = 4$ for the Mexican lending rate equation. Assuming $d = 3$ for Colombia and $d = 4$ for Mexico, the sequence of tests (H_{03}, H_{02}, and H_{01}, respectively) in Table 3B choose the 'logistic' model (3.2a) as the appropriate transition function.[8] The H_0 hypothesis is rejected most strongly at $d = 1$ for the Colombian deposit rate equation and at $d = 5$ for the Mexican deposit rate equation (Table 3A). Given these choices, the sequence of tests in Table 3B favors the 'logistic' model (3.2a) as the appropriate transition function. To sum up, we choose the following transition variables: COL_cvl_{t-3} for the Colombian lending rate model, MEX_cvl_{t-4} for the Mexican lending rate model, COL_cvd_{t-1} for the Colombian deposit rate model and MEX_cvd_{t-5} for the Mexican deposit rate model.

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[6] Notice that the *t*-ratios (but not the coefficient estimates) on the spreads increase by estimating more parsimonious models using the general-to-specific-approach (which deletes one by one the most insignificant variables). A referee warned against estimating parsimonious linear models prior to conducting linearity tests because the tests will then fail to "pick up" various nonlinearities associated with the excluded variables.

[7] It could also be the case that regime switches are driven by macroeconomic variables such as inflation rates. If this is the case, one would expect that a measure of inflation could be the driving transition mechanism between regimes. We considered the use of differenced inflation as a representative indicator of macroeconomic causes of transition, but detailed empirical analysis (available on request) failed to find support for this idea.

[8] Given the choice of $d = 3$ for the Colombian lending rate, the sequence of tests in Table 3B slightly favors the 'quadratic' model over the 'logistic' one, but we could not get sensible estimates for the former. We also estimated logistic and 'quadratic' models assuming $d = 1$, but the results were much less well determined.

4.3 Estimates of the non-linear models

We estimate the STAR model (3.1) using the 'logistic' model (3.2a) by non-linear least squares (NLS). Granger and Teräsvirta (1993) and Teräsvirta (1994) stress particular problems like slow convergence or overestimation associated with estimates of the γ parameter. For this reason, we follow their suggestions in standardizing the exponent of the 'logistic' function (3.2a) by dividing it by the standard deviation of the transition variable, $\sigma(s_{t-d})$ so that γ becomes a scale-free parameter.

Tables 4 and 5 report the NLS estimates of the parsimonious STAR interest rate models. The main parameters of interest in the STAR models are the estimated values of the threshold level, c, and the speed of adjustment, γ. The c estimates reported in Tables 4 to 5 are statistically significant in all models, whereas the estimates of the γ parameter are high for all models indicating that the speed of the transition from $G(s_{t-d}; \gamma, c) = 0$ to $G(s_{t-d}; \gamma, c) = 1$ is rapid at the estimated threshold c. This is evident from Figure 2, which plots the values of the estimated transition function against the spreads. Notice also the high standard error of the γ estimates reported in Tables 4-5. Teräsvirta (1994) and van Dijk *et al*. (2002) point out that this should not be interpreted as evidence of weak non-linearity. Accurate estimation of γ might be difficult as it requires many observations in the immediate neighborhood of the threshold c. Further, large changes in γ have only a small effect on the shape of the transition function implying that high accuracy in estimating γ is not necessary (see the discussion in van Dijk *et al*., 2002).

From Tables 4 to 5, the non-linear models are preferred to the linear ones based on both the AIC and the BIC. In addition, the non-linear specification captures the ARCH effects that are present in the linear interest rate model for Colombia. Parameter stability tests (not reported but available on request) show that the estimates of our regime switching models are much more stable compared to those of the linear models.

5. Discussion of results

Our research identified the existence of non-linear dynamics in the behavior of the lending and deposit interest rates for two emerging markets in Latin America. Moreover, these interest rates exhibit a regime-switching behavior according to the variation of the interest rate spread. The result confirms the importance of the spread rate as a factor affecting the evolution of the lending and deposit rates. Focusing on the lending/interbank rate differential, the regimes we identify have a plausible economic interpretation. The first regime, i.e. $G(s_{t-d};\ \gamma,\ c) = 1$, is defined by positive values of the interest rate spread relative to the threshold. This may be identified with periods of inefficiency in banking activities which in turn adversely affect domestic savings and investments, or with periods of financial crises which tend to be characterized by large values of the interest rate spread. Conversely, the second regime, i.e. $G(s_{t-d};\ \gamma,\ c) = 0$, is defined by periods during which the interest rate spread is less than a threshold. This may be identified with periods of modernization of the banking system which promotes competition within the banking sector, or with "normal" periods of time.

Figure 3 plots the estimated transition functions against time in order to illustrate the succession of the regimes over the sample period. We report the transition functions governed by the lending/interbank rate differential that are easier to interpret. On the other hand, the transition functions governed by the deposit/interbank rate differential provide strong evidence in favor of intermediate regimes and very weak evidence of the extreme values of 0 and 1; however, these are available on request. The estimated transition function for the Colombian lending rate model classifies most of the sample period into the upper regime, which is consistent with the presence of inefficiencies in the banking system. This finding suggests that the policy reforms adopted since the early 1990s (e.g. easier entry for new intermediaries and privatization of public sector firms) have not improved substantially competition and efficiency in the banking system. This is also discussed in Barajas *et al*. (1999, 2000) and Urrutia (2000) who point out that by international standards, bank intermediation spreads and overhead expenses in Colombia are on average around 2 percent higher than those in the rest of Latin America, and around 6 percent higher than those in industrialized countries.[9]

For Mexico, classification most of the 1993-1997 into the upper regime reflects the profound financial crisis that affected the Mexican economy in the mid 1990s. Other reasons that might explain a high interest rate spread include (i) the extent of deposit account dollarization which tends to increase during periods of financial crises [10] (see e.g. Honohan and Shi, 2002) and (ii) the enforcement of bank capital asset requirements in 1993 which had a negative effect on the supply of bank loans (see Chiuri *et al.*, 2002). Since 1998 however, the estimated transition function for Mexico classifies most of the sample period into the lower regime. This is consistent with the return to more "normal" periods of time.

Our estimates in the left panels of Tables 4-5 for the lending rate equations allow for the behavior of the spread (between the lending and the interbank rate) to vary across regimes. By comparing the coefficients for the Colombian lending/interbank rate spread (i.e. COL_cvl_{t-1}) in the two regimes (see left panel of Table 4) we see that when COL_cvl_{t-3} is below the threshold level of 12 percent, the lending interest rate adjusts relatively fast (i.e. the estimated coefficient equals –0.234). On the other hand, when COL_cvl_{t-3} is above 12 percent, the lending interest rate adjusts much slower (i.e. the estimated coefficient equals –0.159). From the left panel of Table 5, when the Mexican spread (i.e. MEX_cvl_{t-4}) is below 3.7 percent, the lending rate adjusts fast (i.e. the estimated coefficient equals –0.592). On the other hand, no significant adjustment occurs

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[9] For their international comparisons, Barajas *et al*. (1999, 2000) and Urrutia (2000) examine three potential measures of efficiency: (i) the difference between the lending and deposit interest rates, (ii) the accounting value of a bank's net interest revenue as a share of its total assets, and (iii) the accounting value of a bank's overhead expenses as a share of its total assets.

[10] For Mexico, the average share of foreign currency deposits in M2 was 11% during the pre-financial crisis years of 1990 to 1993. The share rose to around 17% during 1994 to 1996, and dropped to the pre-crisis level afterwards. The share of foreign currency deposits in total deposits followed a similar pattern of behavior. These figures are taken from Honohan and Shi (2002), Tables A1 and A2.

above equilibrium. For both countries, these spread asymmetries (below and above equilibrium) are statistically significant based on the F-version of the Wald test on equality of spread effects for the two regimes (see bottom line of Tables 4 and 5). For both countries and both regimes, significant effects from the corresponding interbank interest rates are recorded.

The right panels of Tables 4 and 5 report the deposit rate equations allowing for the behavior of the spread (between the deposit and the interbank rate) to vary across regimes. For Colombia, there is fast error correction effect when the interest rate spread $COL\_cvd_{t\text{-}1}$ falls below –1.44 percent (i.e. the estimated coefficient equals –0.449) and slow otherwise (i.e. the estimated coefficient equals –0.097). These asymmetries are statistical significant (see the bottom line of Table 4). For Mexico, there is some very weak error correction effect when the interest rate spread $MEX\_cvd_{t\text{-}5}$ falls below –7.54 percent (i.e. the estimated coefficient equals –0.015) and no effect otherwise. Significant interbank interest rate effects are recorded for both countries and both regimes.

Overall, the results from Tables 4 and 5 show some similarities for both emerging market economies. We present evidence of greater rigidity for lending rate decreases and deposit rate decreases. Therefore, our results offer no clear support for either the hypothesis of collusive pricing arrangements in the banking sectors of Colombia and Mexico or the adverse customer reaction hypothesis. On the loan side, when the spread between the lending and interbank rate is below equilibrium, banks respond by increasing the lending rate rapidly. Further, when the deposit rate is below its equilibrium with the interbank rate, banks respond by increasing the deposit rate rapidly. A possible economic explanation for these results has to do with the investment opportunities available to domestic depositors abroad. In particular, financial liberalization has given domestic residents the opportunity to rebalance their portfolios internationally, achieving a convergence of domestic deposit rates (adjusted for expectations of exchange rate changes) towards international rates. To respond, it could be the case that domestic banks in Colombia and Mexico attempt to keep domestic depositors at home by making deposit rates inflexible downwards. On the other hand, convergence of domestic and international lending rates is less likely to occur due to information costs associated with monitoring domestic borrowers. As a result, international capital markets do not lend directly to companies, rather, foreign lending is intermediated by domestic banks (see also Brock and Rojas-Suarez, 2000). Domestic banks in Colombia and Mexico are possibly exploiting this by making lending rates inflexible downwards. Notice also the lower interest rate spread effects in the Mexican compared with the Colombian deposit rate equation (see the right panels of Tables 4 and 5). Price rigidity in Mexico results from the high market concentration in the Mexican banking system discussed in section 2 (see also the discussion in Hannan and Berger, 1991).

6. Conclusions

In this paper we model the lending and deposit interest rates in two Latin American emerging markets using the smooth transition regime-switching framework. Allowing for different dynamic behavior depending upon large and small interest rate spreads, the non-linear specification seems to work well both in statistical and economic terms. In statistical terms, it captures most of the diagnostic test failures of the linear models. In economic terms, we find evidence of greater rigidity for lending rate decreases and deposit rate decreases. This could be due to the investment opportunities available to domestic depositors abroad following financial liberalization. To respond, it could be the case that domestic banks in Colombia and Mexico attempt to keep domestic depositors at home by making deposit rates inflexible downwards. On the other hand, information costs associated with monitoring domestic borrowers imply that convergence of domestic and international lending rates is less likely to occur. Domestic banks are possibly exploiting this by making lending rates inflexible downwards. Finally, price rigidity in Mexico results from the high market concentration in the Mexican banking system.

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Table 1: Dickey and Fuller unit root tests for the interest rate series

Variable	Lags	τ_C	Order of Integration
COL_d_t	1	-0.885	$\sim I(1)$
COL_l_t	3	-1.136	$\sim I(1)$
COL_ib_t	1	-2.170	$\sim I(1)$
$COL\_cvd_t = COL\_d_t - COL\_ib_t$	1	-6.099 **	$\sim I(0)$
$COL\_cvl_t = COL\_l_t - COL\_ib_t$	1	-5.450 **	$\sim I(0)$
MEX_d_t	1	-2.222	$\sim I(1)$
MEX_l_t	2	-2.562	$\sim I(1)$
MEX_ib_t	2	-2.480	$\sim I(1)$
$MEX\_cvd_t = MEX\_d_t - MEX\_ib_t$	0	-3.400 *	$\sim I(0)$
$MEX\_cvl_t = MEX\_l_t - MEX\_ib_t$	0	-5.544 **	$\sim I(0)$

τ_C indicates that the Dickey-Fuller regression contains a constant.

* indicates that the null hypothesis is rejected at a 5% significance level.

** indicates that the null hypothesis is rejected at a 1% significance level.

Table 2: Estimated linear models

Colombia: Lending rate model			Colombia: Deposit rate model		
$\Delta COL\_l_{t\text{-}1}$	-0.017	(0.079)	$\Delta COL\_d_{t\text{-}1}$	0.340	(0.077)
$\Delta COL\_l_{t\text{-}2}$	-0.014	(0.074)	$\Delta COL\_d_{t\text{-}2}$	-0.200	(0.081)
$\Delta COL\_l_{t\text{-}3}$	0.189	(0.074)	$\Delta COL\_d_{t\text{-}3}$	0.114	(0.080)
$\Delta COL\_l_{t\text{-}4}$	0.063	(0.074)	$\Delta COL\_d_{t\text{-}4}$	-0.122	(0.080)
ΔCOL_ib_{t}	0.146	(0.021)	$\Delta COL\_d_{t\text{-}5}$	0.069	(0.067)
$\Delta COL\_ib_{t\text{-}1}$	0.046	(0.038)	ΔCOL_ib_{t}	0.163	(0.014)
$\Delta COL\_ib_{t\text{-}2}$	0.045	(0.035)	$\Delta COL\_ib_{t\text{-}1}$	0.044	(0.032)
$\Delta COL\_ib_{t\text{-}3}$	0.052	(0.032)	$\Delta COL\_ib_{t\text{-}2}$	0.047	(0.030)
$\Delta COL\_ib_{t\text{-}4}$	0.022	(0.023)	$\Delta COL\_ib_{t\text{-}3}$	0.050	(0.027)
$COL\_cvl_{t\text{-}1}$	-0.117	(0.036)	$\Delta COL\_ib_{t\text{-}4}$	0.082	(0.025)
			$\Delta COL\_ib_{t\text{-}5}$	0.044	(0.017)
			$COL\_cvd_{t\text{-}1}$	-0.100	(0.032)
$\hat{\sigma}_l$	1.414			0.946	
AIC	3.600			2.805	
BIC	3.811			3.058	
AR(12)	1.263	[0.248]		1.332	[0.208]
ARCH(12)	5.056	[0.000]		3.101	[0.001]
NORM(2)	136.380	[0.000]		44.163	[0.000]

Mexico: Lending rate model			Mexico: Deposit rate model		
$\Delta MEX\_l_{t\text{-}1}$	-0.166	(0.132)	$\Delta MEX\_d_{t\text{-}1}$	0.098	(0.093)
$\Delta MEX\_l_{t\text{-}2}$	-0.581	(0.148)	$\Delta MEX\_d_{t\text{-}2}$	-0.166	(0.094)
$\Delta MEX\_l_{t\text{-}3}$	-0.453	(0.149)	$\Delta MEX\_d_{t\text{-}3}$	0.050	(0.071)
$\Delta MEX\_l_{t\text{-}4}$	-0.143	(0.123)	ΔMEX_ib_{t}	0.423	(0.016)
ΔMEX_ib_{t}	0.938	(0.035)	$\Delta MEX\_ib_{t\text{-}1}$	0.118	(0.042)
$\Delta MEX\_ib_{t\text{-}1}$	0.351	(0.131)	$\Delta MEX\_ib_{t\text{-}2}$	0.024	(0.044)
$\Delta MEX\_ib_{t\text{-}2}$	0.472	(0.147)	$\Delta MEX\_ib_{t\text{-}3}$	0.036	(0.038)
$\Delta MEX\_ib_{t\text{-}3}$	0.478	(0.137)	$MEX\_cvd_{t\text{-}1}$	-0.026	(0.018)
$\Delta MEX\_ib_{t\text{-}4}$	0.078	(0.117)			
$MEX\_cvl_{t\text{-}1}$	-0.148	(0.097)			
$\hat{\sigma}_l$	1.517			0.917	
AIC	3.777			2.743	
BIC	4.037			2.962	
AR(12)	1.535	[0.127]		1.306	[0.229]
ARCH(12)	0.595	[0.840]		3.787	[0.000]
NORM(2)	43.524	[0.000]		20.024	[0.000]

Estimates of the constant terms are not reported. Standard errors in (•) and *p*-values in [•]. $\hat{\sigma}_l$: regression standard error. AIC: Akaike information criterion. BIC: Schwarz Bayesian information criterion. AR(12): F-test for up to 12th order serial correlation. ARCH(12): F-test for up to 12th order Autoregressive Conditional Heteroscedasticity. NORM(2): Chi-square test for normality.

Table 3: Test for linearity and STAR model selection

Panel A: Linearity tests

Delay	Colombia		Mexico	
	Lending rate model	Deposit rate model	Lending rate model	Deposit rate model
	Transition variable:	Transition variable:	Transition variable:	Transition variable:
d	COL_cvl_{t-d}	COL_cvd_{t-d}	MEX_cvl_{t-d}	MEX_cvd_{t-d}
1	**0.000000**	**0.000002**	0.000005	0.002689
2	0.000001	0.001270	0.000130	0.000009
3	**0.000000**	0.000606	0.072997	0.000004
4	0.000010	0.000022	**0.000004**	0.128496
5	0.000015	0.000003	0.011405	**0.000002**
6	0.000395	0.454903	0.000047	0.000063

Panel B: STAR model selection

Country and model	Delay d	$H_{03}: \phi_{3,j} = 0$	$H_{02}: \phi_{2,j} = 0 \mid \phi_{3,j} = 0$	$H_{01}: \phi_{1,j} = 0 \mid \phi_{3,j} = \phi_{2,j} = 0$	Type of Model
Colombia					
Lending rate model	3	0.006167	**0.000001**	0.055291	LSTAR
Deposit rate model	1	0.034236	0.177982	**0.000000**	LSTAR
Mexico					
Lending rate model	4	**0.000377**	0.007989	0.019665	LSTAR
Deposit rate model	5	0.008685	0.057629	**0.000018**	LSTAR

Notes: The Table reports the *p*-values of the linearity tests developed in section 3. Panel A reports the H_0 test for linearity. Figures in bold denote the minimum probability value of the H_0 test over the interval $1 \leq d \leq 6$. Panel B reports the *p*-values of the nested H_{03}, H_{02} and H_{01} tests for selecting between the 'logistic' model and the 'quadratic logistic' model for the transition function of the STAR models. Figures in bold denote the lowest *p*-value for the three tests. All *p*-values refer to the F-version of the LM test.

Table 4: Estimated non-linear models for Colombia

Lending rate model Transition variable: $COL\_cvl_{t\text{-}3}$ Variable	Coeff.	($s.e.$)	Deposit rate model Transition variable: $COL\_cvd_{t\text{-}1}$ Variable	Coeff.	($s.e.$)
<u>*eq*.1</u>			<u>*eq*.1</u>		
$\Delta COL\_l_{t\text{-}1}$	-0.097	(0.097)	$\Delta COL\_d_{t\text{-}2}$	-0.480	(0.121)
$\Delta COL\_l_{t\text{-}3}$	0.128	(0.087)	ΔCOL_ib_t	0.189	(0.036)
ΔCOL_ib_t	0.169	(0.027)	$\Delta COL\_ib_{t\text{-}1}$	-0.179	(0.065)
$\Delta COL\_ib_{t\text{-}1}$	-0.054	(0.049)	$\Delta COL\_ib_{t\text{-}2}$	-0.078	(0.035)
$\Delta COL\_ib_{t\text{-}2}$	-0.099	(0.039)	$\Delta COL\_ib_{t\text{-}4}$	-0.072	(0.032)
$COL\_cvl_{t\text{-}1}$	-0.234	(0.043)	$COL\_cvd_{t\text{-}1}$	-0.449	(0.095)
<u>*eq*.2</u>			<u>*eq*.2</u>		
$\Delta COL\_l_{t\text{-}1}$	0.290	(0.119)	$\Delta COL\_d_{t\text{-}1}$	0.349	(0.074)
$\Delta COL\_l_{t\text{-}3}$	0.147	(0.127)	$\Delta COL\_d_{t\text{-}2}$	0.076	(0.089)
$\Delta COL\_l_{t\text{-}4}$	0.192	(0.143)	ΔCOL_ib_t	0.157	(0.017)
ΔCOL_ib_t	0.125	(0.028)	$\Delta COL\_ib_{t\text{-}1}$	0.049	(0.021)
$\Delta COL\_ib_{t\text{-}3}$	0.036	(0.039)	$\Delta COL\_ib_{t\text{-}3}$	0.018	(0.016)
$\Delta COL\_ib_{t\text{-}4}$	0.028	(0.031)	$\Delta COL\_ib_{t\text{-}4}$	0.064	(0.018)
$COL\_cvl_{t\text{-}1}$	-0.159	(0.029)	$\Delta COL\_ib_{t\text{-}5}$	0.028	(0.014)
			$COL\_cvd_{t\text{-}1}$	-0.097	(0.030)
γ	60.961	(132.13)	γ	5.947	(3.760)
c	12.092	(0.208)	c	-1.443	(0.603)
$\hat{\sigma}_{nl}$	1.347			0.868	
$\hat{\sigma}^2_{nl} / \hat{\sigma}^2_l$	0.907			0.842	
AIC	3.504			2.637	
BIC	3.776			2.949	
AR(12)	1.081	[0.382]		1.4141	[0.168]
ARCH(12)	2.241	[0.014]		1.7799	[0.060]
NORM(2)	78.474	[0.000]		13.036	[0.002]
WALD	7.411	[0.007]		14.291	[0.000]

The Table reports NLS estimates of the LSTAR model $y_t = (eq.1)(1 - G(\bullet)) + (eq.2)\,G(\bullet)$, where $y_t = \Delta COL\_l_t$, ΔCOL_d_t, and $G(\bullet) = \{1 + \exp[-\gamma(s_{t\text{-}d} - c)/\sigma(s_{t\text{-}d})]\}^{-1}$ is the "logistic" transition function, with $s_{t\text{-}d}$ as the transition variable. Estimates of the constant terms are not reported. Values of 0 and 1 of $G(\bullet)$ are associated with the two alternative regimes. The speed of transition between the two regimes is determined by the parameter γ, and c denotes the threshold between the two regimes. WALD is an F-test on equality of spread effects for the two regimes. Other diagnostic test statistics are described in the notes to Table 2.

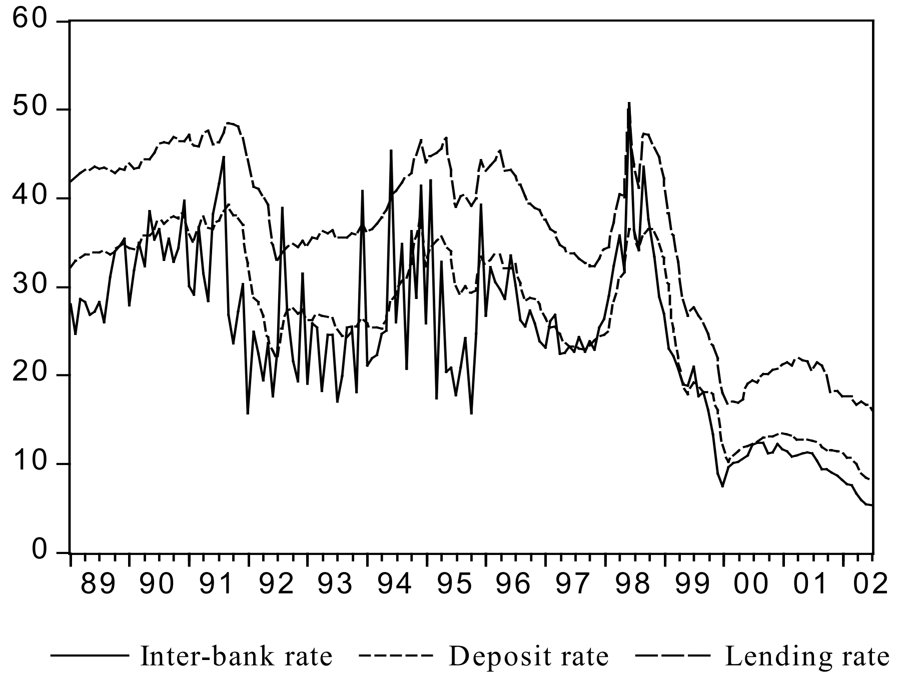
Table 5: Estimated non-linear models for Mexico

Lending rate model Transition variable: $MEX\_cvl_{t\text{-}4}$ Variable	 Coeff.	 (*s.e.*)	Deposit rate model Transition variable: $MEX\_cvd_{t\text{-}5}$ Variable	 Coeff.	 (*s.e.*)
<u>*eq*.1</u>			<u>*eq*.1</u>		
$\Delta MEX\_l_{t\text{-}3}$	0.287	(0.103)	$\Delta MEX\_d_{t\text{-}1}$	0.225	(0.127)
$\Delta MEX\_l_{t\text{-}4}$	-0.030	(0.024)	ΔMEX_ib_t	0.521	(0.049)
ΔMEX_ib_t	0.888	(0.031)	$\Delta MEX\_ib_{t\text{-}1}$	-0.086	(0.086)
$\Delta MEX\_ib_{t\text{-}3}$	-0.288	(0.119)	$MEX\_cvd_{t\text{-}1}$	-0.015	(0.014)
$MEX\_cvl_{t\text{-}1}$	-0.592	(0.088)			
<u>*eq*.2</u>			<u>*eq*.2</u>		
$\Delta MEX\_l_{t\text{-}2}$	-0.795	(0.266)	$\Delta MEX\_d_{t\text{-}1}$	-0.143	(0.106)
$\Delta MEX\_l_{t\text{-}3}$	-0.760	(0.182)	$\Delta MEX\_d_{t\text{-}2}$	-0.155	(0.061)
ΔMEX_ib_t	1.176	(0.069)	ΔMEX_ib_t	0.426	(0.018)
$\Delta MEX\_ib_{t\text{-}1}$	0.293	(0.098)	$\Delta MEX\_ib_{t\text{-}1}$	0.307	(0.049)
$\Delta MEX\_ib_{t\text{-}2}$	0.674	(0.197)			
$\Delta MEX\_ib_{t\text{-}3}$	0.778	(0.170)			
$MEX\_cvl_{t\text{-}1}$	0.018	(0.079)			
γ	58.553	(28.094)	γ	8.469	(7.915)
c	3.723	(0.037)	c	-7.545	(0.733)
$\hat{\sigma}_{nl}$	1.161			0.769	
$\hat{\sigma}^2_{nl} / \hat{\sigma}^2_{l}$	0.586			0.703	
AIC	3.226			2.371	
BIC	3.546			2.592	
AR(12)	1.718	[0.078]		1.126	[0.350]
ARCH(12)	1.577	[0.118]		0.576	[0.855]
NORM(2)	11.270	[0.004]		25.027	[0.000]
WALD	30.379	[0.000]		1.291	[0.258]

The Table reports NLS estimates of the LSTAR model $y_t = (eq.1)(1 - G(\bullet)) + (eq.2)\, G(\bullet)$, where $y_t = \Delta MEX\_l_t$, ΔMEX_d_t, and $G(\bullet) = \{1 + \exp[-\gamma(s_{t\text{-}d} - c)/\sigma(s_{t\text{-}d})]\}^{-1}$ is the "logistic" transition function, with $s_{t\text{-}d}$ as the transition variable. Estimates of the constant terms are not reported. Values of 0 and 1 of $G(\bullet)$ are associated with the two alternative regimes. The speed of transition between the two regimes is determined by the parameter γ, and c denotes the threshold between the two regimes. WALD is an F-test on equality of spread effects for the two regimes. Other diagnostic test statistics are described in the notes to Table 2.

Figure 1: Interest rate series

Colombia



Mexico

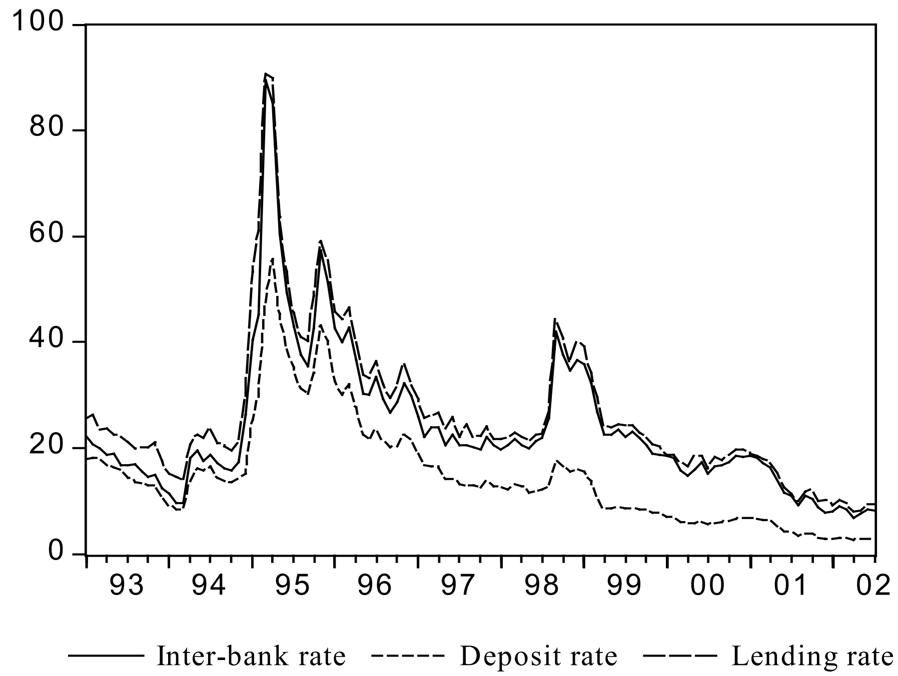
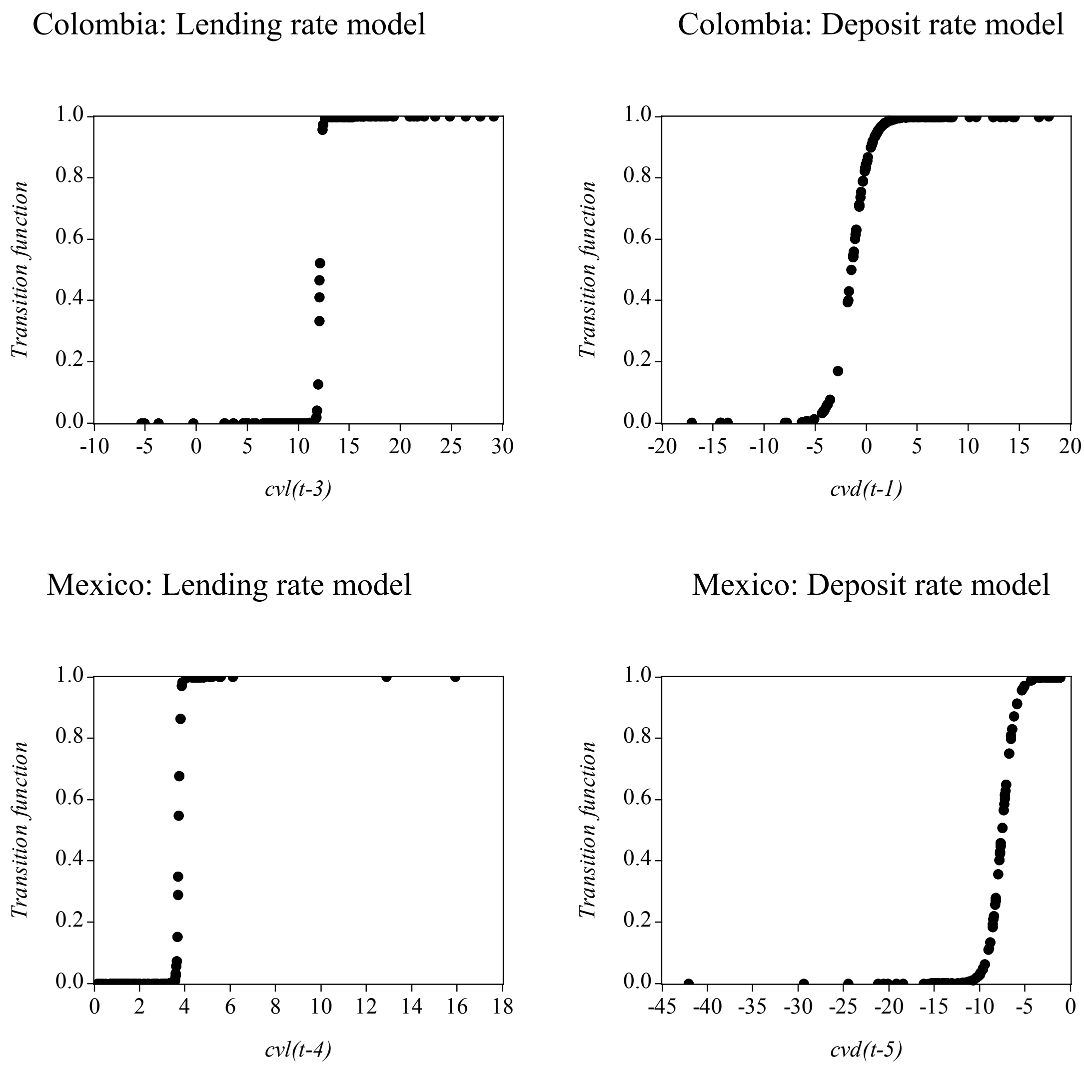


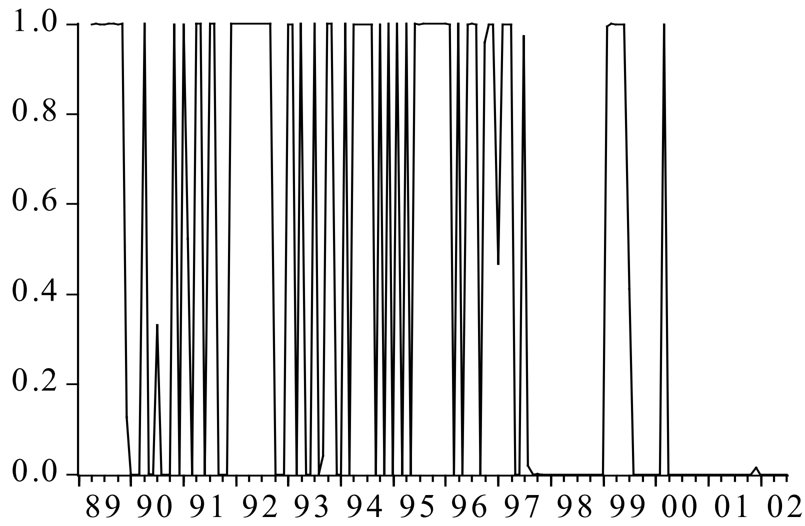
Figure 2: Estimated transition functions versus transition variables



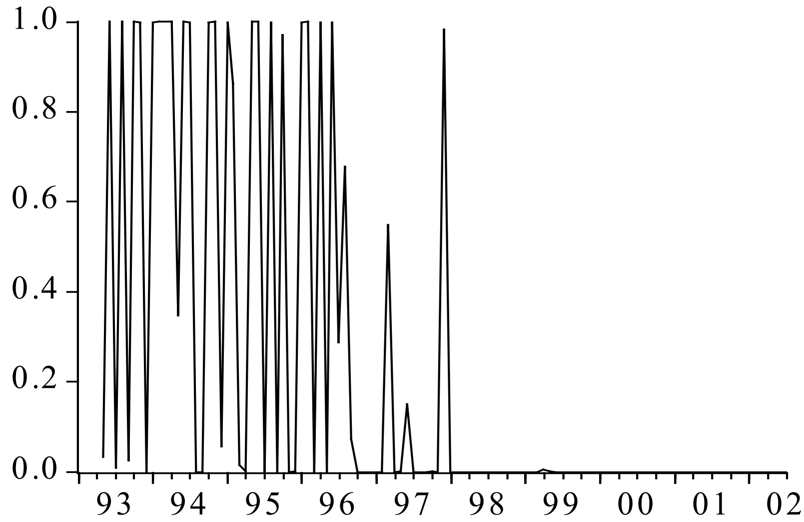
Estimated transition functions from the corresponding STAR models presented in Tables 4 and 5 for Colombia and Mexico, respectively.

Figure 3: Estimated transition functions for the lending rate models against time

Colombia



Mexico



Estimated transition functions from the corresponding STAR lending rate models reported in the left panels of Tables 4 and 5 for Colombia and Mexico, respectively. Extreme values of 0 and 1 of the transition functions are associated with the two alternative regimes.