External shocks, Bank Lending Spreads, and Output Fluctuations

Argentina in the Aftermath of the Tequila Effect, 1995-96

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Abstract

This paper studies the effects of external shocks on bank lending spreads and output fluctuations in Argentina during the early 1990s. The first part presents the analytical framework. The second presents a VAR model that relates bank lending spreads, the cyclical component of output, the real lending rate, and the external interest rate spread. Impulse response functions show that a positive shock in external spreads leads to higher domestic spreads and lower output. Historical decompositions show that shocks to external spreads in the immediate aftermath of the Mexican peso crisis had a sizable effect on movements in output and domestic spreads.

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1 Introduction

Argentina faced a severe economic downturn in 1995 and early 1996. Output, domestic credit, and stock prices fell dramatically. A massive shift away from peso-denominated deposits was associated with large capital outflows, a sharp drop in official foreign reserves, and a contraction of the monetary base. Unemployment peaked at almost 19 percent in May 1995 and remained high in subsequent months. The liquidity crunch led to a sharp rise in bank lending rates, on both peso- and US dollar-denominated loans. At the same time, the spread between the lending rates on peso- and US dollar-denominated loans widened significantly between February and May 1995 (as shown in Figure 1), reflecting an increase in the perceived risk of a collapse of the currency board regime introduced in 1991 and a subsequent large exchange rate depreciation. The spread between deposit and lending rates, both in pesos and in US dollars, also increased sharply.

[Insert Figure 1 about here]

The economic downturn in Argentina took place in the context of an adverse external shock—an abrupt change in market sentiment regarding the country's economic prospects, triggered by expectations that the currency board regime would collapse. Various observers attributed this phenomenon to a contagion effect triggered by the Mexican peso crisis of December 1994. Indeed, as suggested by the "wake-up call" hypothesis, a crisis in one country can prompt a reassessment of other countries' fundamentals (see Prisker (2001)). Countries with perceived weak fundamentals may therefore be subject to a shift in market sentiment or increased risk aversion, which may translate into higher borrowing spreads.

The purpose of this paper is to provide an analytical framework for understanding the transmission process of external shocks of this type, as well as empirical evidence for Argentina. Specifically, our analysis models external shocks as a temporary increase in the risk premium faced by domestic borrowers on world capital markets—that is, an increase in external interest rate spreads. Our approach is motivated in large part by the sharp increase in interest rate spreads (relative to US rates) on liabilities issued by private—as well as public—borrowers from Argentina in the immediate aftermath of the Mexican peso crisis (Figure 1). As noted earlier, some observers have viewed the timing of this increase as the consequence of a "pure" contagion effect. The real effects of a shock of this nature are analyzed, both analytically and empirically, in a model that incorporates a link between bank credit and the supply side through firms' demand for working capital, domestic interest rate spreads, and real lending rates. As documented for instance by the Inter-American Development Bank (2004), banks account for a sizable fraction of the financing needs of firms in Argentina as well as other Latin American countries.[1]

In general, spreads between lending and deposits rates in most developing countries tend to be relatively large for a variety of reasons—including high required reserve ratios, a limited degree of competition in the financial system, low productive efficiency of financial institutions, and selective credit and interest controls that require these institutions to undertake a substantial amount of concessionary lending. Several studies, in particular, have emphasized the role of market structure.[2] In an empirical study of the determinants of bank spreads in Argentina, for instance, Cãtao (1998) found, using aggregate monthly data for the period June 1993-July 1997, that spreads are positively influenced by the degree of market concentration.[3] He interprets

[1]Agénor (2005), Edwards and Végh (1997), Greenwald and Stiglitz (1993), and Isard et al. (1996) discuss various models which also account explicitly for the link between firms' working capital needs and bank credit.

[2]Among other studies of the determinants of bank spreads are Barajas, Steiner, and Salazar (1999) for Colombia, Afanasieff, Lhacer, and Nakane (2002) for Brazil, and Demirgüç-Kunt and Huizinga (1999) for a large group of countries. Early studies include Ho and Sanders (1981), and Hanson and de Resende Rocha (1986).

[3]Catão used, as we do in our empirical analsis, *ex ante* (or contract) interest rates, rather than *effective* interest rates (obtained from the income statements of commercial banks). As is well known, these two measures can differ markedly in a setting where the

this result as reflecting the fact that peso borrowers in Argentina during that time were unable to arbitrage between domestic and foreign sources of funds, and thus became subject to the monopoly power of local banks. He also found that spreads are also responsive to operating costs and the share of non-performing loans, and to a lesser degree exchange rate risk and the cost of liquidity requirements. Our analysis, by contrast, focuses on the role of external factors, in addition to default risk. In contrast to existing studies, we focus on the role of domestic interest rates in the transmission process of external shocks to output.

The remainder of the paper proceeds as follows. Section II presents the analytical framework, which focuses on the determination of domestic bank lending spreads in the presence of verification and enforcement costs associated with loan contracts. The analysis shows how domestic financial intermediation spreads are related to default probabilities, underlying domestic shocks, and external spreads. Section III estimates a vector autoregression model using monthly data for Argentina (for the period June 1993-June 1998) that relates (in line with the analytical framework) the *ex ante* bank lending spread, the cyclical component of output, the real bank lending rate, the effective reserve requirement ratio, and the external interest rate spread. Variance decompositions are discussed in Section IV. Section V uses impulse response functions to analyze the effects of a contagious shock, defined as an increase in the external spread. Section VI assesses the movements in output and interest rates in Argentina in the immediate aftermath of the Mexican peso crisis of December 1994. Section VII summarizes the main results of the analysis and offers some concluding remarks.

incidence of nonperforming loans is high and refinancing operations are widespread.

2 The Analytical Framework

The importance of the credit channel in the transmission mechanism of macroeconomic shocks in developing countries is now well established analytically and empirically (see Agénor (2004, Chapter 5)). The aspect that we focus on in this study relates to the direct impact of credit on producers who finance their working capital needs via the banking system. Banks engage frequently in costly monitoring and supervision of creditors' performance, to ensure the proper use of credit, and its timely repayment. As the frequency of costly monitoring increases in turbulent times, the credit channel provides a natural way to model the effects of domestic and external shocks on economic activity in developing countries. This section outlines a simplified version of the analytical framework developed by Agénor and Aizenman (1998, 1999), which highlights the impact of productivity and shocks to the external cost of credit on domestic output.[4]

We consider an economy where risk-neutral banks provide intermediation services. There is no equity market, so producers cannot issue claims on their capital stock. Agents (producers) demand credit from banks (lenders) to finance their working capital needs. Producers rely solely on bank credit to finance the cost of variable inputs, which must be paid prior to production and the sale of output. Output is subject to random productivity shocks. The realized productivity shock is revealed to banks only at a cost. In the event of default by any given producer on its bank loans, the creditor seizes a fraction of the realized value of output. Seizing involves two types of costs: first, verifying the net value of output (the state of nature) is costly; second, enforcing repayment (in case of default) requires costly recourse to the legal system.

[4] The Agénor-Aizenman framework combines the costly state verification approach pioneered by Townsend (1979) and the model of limited enforceability of contracts used in the external debt literature, as in Helpman (1989).

Output of producer i is given by

$$y_i = M_i^{\beta}(1 + \delta_0 + \delta_m + \varepsilon_i), \quad 0 < \beta < 1, \ |\varepsilon_i| \leq \Gamma < 1, \tag{1}$$

where M_i denotes the variable input (which may include not only labor, but also raw materials or intermediate products) used by producer i, ε_i is the realized i.i.d. productivity shock, $1 + \delta_0$ is expected productivity, and δ_m is the realized common macroeconomic shock, which is assumed to be distributed binomially:

$$\delta_m = \begin{cases} \nu & \text{probability } 0.5 \\ -\nu & \text{probability } 0.5 \end{cases}.$$

The contractual interest rate on loans made to producer i is r_L^i. Because producers must finance input costs prior to the sale of output, producer i's total variable costs are $(1 + r_L^i)p_m M_i$, where p_m is the relative price of the variable input (taken as given by each producer).

We assume that the bank has information about the input choice of the producer and determines the interest rate such that the expected net repayment equals the cost of credit. Each bank is assumed to deal with a large number of independent producers; this, in turn, allows the bank to diversify the idiosyncratic risk, ε_i. Henceforth we also assume that no default would occur in the good state of the macro shock, but that (at least) some producers will default partially in the bad state of the aggregate shock.[5]

In general, a producer will choose to default if

$$\kappa M_i^{\beta}(1 + \delta_0 - \nu + \varepsilon_i) < (1 + r_L^i)p_m M_i, \tag{2}$$

[5]The key results of our discussion hold even if this assumption is not valid. This assumption is equivalent to

$$\kappa M_i^{\beta}(1 + \delta_0 + \nu - \Gamma) > (1 + r_L^i)p_m M_i > \kappa M_i^{\beta}(1 + \delta_0 - \nu - \Gamma),$$

and will hold if the degree of volatility of the aggregate shock (as measured by ν) is high enough.

where κ is the fraction of realized output that the bank is able to seize in case of default. The left-hand side of equation (2) is the producer's repayment following a default, whereas the right-hand side is the contractual repayment. We denote by ε_i^{max} the highest productivity shock leading to default—that is, the value of ε_i for which (2) holds as an equality:

$$\kappa M_i^{\beta}(1+\delta_0-\nu+\varepsilon_i^{max}) = (1+r_L^i)p_m M_i. \tag{3}$$

If default never occurs, ε_i^{max} is set at the lower end of the support ($\varepsilon_i^{max} = -\Gamma$). In case of default, the bank's net revenue is the producer's repayment minus the state verification and contract enforcement cost, assumed to be proportional to the cost of borrowed funds, c_i:

$$\kappa M_i^{\beta}(1+\delta_0-\nu+\varepsilon_i) - c_i(1+r^*)p_m M_i, \tag{4}$$

where $0 < c_i < 1$.[6]

We also assume that banks have access to an elastic supply of funds, at a real cost of r^*.[7] This source of funds may be credit provided by foreign banks, as modeled by Agénor and Aizenman (1998), but a more general interpretation is to assume simply that domestic banks borrow on world capital markets. We do not explicitly model the (endogenous) spread above and over a "risk-free" interest rate that such lending operations may give rise to (as for instance in Agénor and Aizenman (1998)), but a change in r^* can obviously be interpreted as a change in the *exogenous* component of the external risk premium, that is, a change that is unrelated to domestic market conditions—a "pure contagion" effect, as for instance in Agénor (2005).

[6] The cost c_i is paid by banks in order to identify the productivity shock ε_i, and to enforce adequate payment. The analysis is more involved if some costs are paid *after* obtaining the information about ε_i. In these circumstances, banks will refrain from forcing debt repayment when realized productivity is below an "enforcement threshold." For simplicity of exposition, we refrain from modeling this possibility. We ignore also all other real costs associated with financial intermediation. Adding these considerations would not modify the key results discussed below.

[7] This source of funds may be credit provided by foreign banks, as modeled by Agénor and Aizenman (1998). However, we do not explicitly model the interest rate spread above and over the "risk-free' interest rate that may arise.

Assuming that banks are risk neutral and competitive, the contractual interest rate is determined by an expected break-even condition, derived in Appendix A. As also shown there, the contractual interest rate, r_L^i, is determined by a mark-up rule. r_L^i exceeds the bank's cost of funds, r^*, by the sum of two terms: the first is the expected revenue lost due to partial default in bad states of nature, and the second measures the expected state verification and contract enforcement costs.[8] In the particular case where the idiosyncratic shock follows a uniform distribution, the spread (A2) is characterized by a quadratic equation, which can be combined with (3) to derive a reduced-form solution for the probability of default and for the domestic interest rate.

In general, the domestic interest rate/external cost of credit curve, plotted in the r_L^i-r^* space, is backward-bending, and a given r^* can be associated with two values of r_L^i. This follows from the presence of a trade-off between the interest rate and the frequency of full repayment.[9] The efficient point is associated with the lower interest rate, as more frequent default is associated with a lower expected surplus (see equation (A4) in Appendix A). In what follows we will assume that competitive banks choose the efficient point, and will ignore the backward-bending portion of the r_L^i-r^* curve. For an internal solution where credit is (elastically) supplied and the probability of default is positive, the following proposition can be shown to hold:

Proposition 1. *A higher external cost of credit, r^*, raises the domestic lending rate, lowers the demand for variable inputs, and reduces expected output.*

As discussed in Appendix A, the magnitude of these effects increases with the responsiveness of the domestic interest rate to the cost of funds faced by

[8] Appendix A also derives the producer's expected net income; the optimal level of use of the variable input, M_i, is found by maximizing that expression.

[9] A higher interest rate would increase the probability of default, implying that the net effect of a higher interest rate on the expected repayment is determined by elasticity considerations. This effect is quite standard in this type of models.

domestic banks, $\partial r_L^i/\partial r^*$, and are maximized as we approach the backward-bending portion of the supply of credit that producers face.

Suppose now that the external cost of credit, r^*, can be decomposed into a risk-free rate, r_f^* (as proxied, for instance, by the interest rate on U.S. Treasury bonds), and a risk premium, θ. This premium (which we also refer to as the external spread) is taken as given for the moment. Suppose also that depositors are risk neutral and have free access to offshore accounts (or equivalently, if capital controls exist, they are not effective). Then, if domestic banks are indifferent regarding the source of funds for their lending operations (that is, if domestic deposits and foreign borrowing are perfect substitutes), the domestic deposit rate, r_d, would be set equal to r_f^*.[10] Then, from equation (A5) in Appendix A (which assumes that the idiosyncratic shock is distributed uniformly), we have

$$r_L^i - r_d = \theta + F(\Phi_i), \tag{5}$$

where Φ_i is the probability of default. The following corollary to Proposition 1 therefore holds:

Corollary. *A higher external spread, θ, raises the domestic bank spread, lowers the demand for variable inputs, and reduces expected output.*

The results summarized in Proposition 1 and its corollary are consistent with those obtained with more developed, general equilibrium models, such as those of Agénor (2005) and Neumeyer and Perri (2005). In both of these papers, the cost of credit affects the effective cost of labor, and the external premium is decomposed into an endogenous, domestic-related component and an exogenous component. There are, however, important differences between the models developed in these studies. The Neumeyer-Perri framework

[10] The reason why banks would set r_d equal to r_f^*, instead of r^f, is because the latter is the rate faced only by domestic borrowers (for which there is a risk of default), not depositors. Indeed, if depositors can only obtain r_f^* as an alternative return to their savings, domestic banks would not exceed that rate if they view domestic and foreign sources of funds as perfect substitutes.

is non-monetary in nature, so that working capital needs depend on *real* interest rates. By contrast, in Agénor (2005), external shocks are transmitted through fluctuations in *nominal* interest rates. In addition, the very nature of the Neumeyer-Perri model precludes any account of the role of the banking system in the transmission process of external shocks. Indeed, in their model, where firms borrow from domestic households and foreign investors, domestic lenders always receive back the full value of their loans (plus interest). They play therefore no role in the transmission of external shocks. In Agénor (2005), by contrast, the banking system is explicitly considered. However, this is done in a deterministic setting with no account of credit market imperfections. Our analysis is therefore best seen as complementary to existing studies of the real impact of external shocks induced by "pure contagion" or "wake-up call" effects.

3 VAR Estimation and Analysis

We now dwell on the analytical framework developed above to analyze empirically Argentina's economic downturn in the immediate aftermath of the 1994 Mexican peso crisis. The model's explicit account of the role of external financial shocks in the determination of domestic interest rates and output makes it particularly suitable for that purpose. To implement our framework empirically we use vector autoregression (VAR) techniques and focus on the following variables: the external interest rate spread, ES, the domestic interest rate spread on peso-denominated assets and liabilities, DS, the real lending rate, RL, and two alternative measures of output: deviations of current output from its trend level, $ln(y/y_T)$, and the growth rate of output, $ln(y/y_{-12})$. The trend component y_T is obtained by applying the Hodrick-Prescott filter. We refer in what follows to the model with $ln(y/y_T)$ as Model A, and the one with $ln(y/y_{-12})$ as Model B.[11] Both models are

[11] Appendix B provides precise data definitions. The results of augmented Dickey-Fuller and Phillips-Perron unit root tests are mixed due to the relatively short time spun by the

estimated with monthly data from January 1993 through June 1998.

The link between the analytical framework developed in the previous section and the VAR specification adopted for our empirical investigation is straightforward. We use detrended output, given our focus on short-run fluctuations. The external interest rate spread corresponds, in the model, to the (country-specific) component of the foreign interest rate. As noted earlier, a change in r^* in our analytical framework can be interpreted (for a given risk-free rate) as a change in the risk premium faced by domestic banks on world capital markets, that is, a change in θ. But because, in practice, external spreads may depend on both domestic factors (such as changes in output and domestic credit conditions) and external conditions (such as abrupt changes in market sentiment), we place that variable last in the ordering of the VAR model, in order to "purge" it of its domestic component. In doing so, we are able to capture primarily the exogenous component of the external spread shock (as emphasized in the foregoing discussion) when calculating variance decompositions and impulse response functions. This ordering also allows us to account for the fact that movements in external spreads, despite representing the initial (exogenous) "impulse", may respond subsequently to changes in domestic variables—as for instance in Agénor and Aizenman (1998).

The inclusion of the real lending rate and the domestic bank spread in the VAR is also consistent with our analytical framework. The first variable captures the "level" effect of interest rates on output. which results from the impact of changes in the cost of funds on the optimal demand for variable inputs, as established in Appendix A: the pay-in-advance constraint implies that input costs also reflect borrowing costs. The second variable depends on the external spread (as discussed earlier) or, as we emphasize here, the autonomous component of that variable. As implied by (5), movements in

sample period over which they are done; the series are taken, nonetheless, to be stationary on economic grounds (see Campbell and Perron (1991)).

the domestic spread also depend on changes in the risk premium that banks charge to their borrowers; this premium, in turn, reflects changes in the (perceived) risk of default. To the extent that default risk tends to vary with the state of the business cycle—during recessions, default rates tend to increase, and vice versa—the inclusion of a cyclical measure of output in our VAR allows us to capture it indirectly.

It is worth noting that we do not explicitly introduce credit to firms as an additional variable in the VAR; indeed, in line with our theoretical analysis, we assume that the supply of credit (both at the firm and bank levels) is perfectly elastic at the given rate. Thus, there is no credit rationing. Put differently, credit does *not* represent an independent information variable in our analysis.

Finally, in addition to the variables listed above, we also considered expanded VAR models with the average effective reserve requirement rate, in an attempt to control for changes in the cost of financial intermediation.[12] Although reserve requirement rates did change significantly during the sample period, the results obtained from this expanded model were not qualitatively different from those obtained from the smaller version. Given the relatively short sample size, we opted to present the results based on the more parsimonious versions of the model. The number of lags included in the estimated models (as discussed in Appendix B) was set to three months.

4 Variance Decompositions

Table 1 presents, for both models, the variance decompositions for all the variables included in the system. Following the discussion of the results below, the table shows the share of the variance associated with shocks to

[12] Of course, various other factors (such as changes in taxation of financial services) may affect domestic lending spreads, in addition to reserve requirement rates. Our analysis implicitly takes these factors as given. This assumption is appropriate to the extent that they fluctuated relatively little within the sample period.

ES, and the sum of the shares of the variance associated with shocks to the other variables in the models.[13] As noted before, ES is placed last in the Choleski ordering in an effort to purge this shock of the domestic factors that it could reflect. Qualitatively, the results are robust to the measurement of cyclical output.

[Insert Table 1 about here]

Interestingly, the share of the variance of DS or RL associated with ES shocks is small. This is the case regardless of the specific ordering chosen, or the measurement of cyclical output. The share ranges between 5 to 10 percent after 24 months, and is less than 3 or 4 percent at shorter horizons. At face value these results suggest that on average between January 1993 and June 1998, movements in DS and RL were mostly associated with shocks originating within Argentina. But as discussed later, these results are not entirely corroborated by the historical decompositions.

The share of the variance of the cyclical component of output associated with ES shocks is more substantial. Although the specifics depend on the choice of the cyclical output measure, the share increases with the horizon, and with a horizon of 24 months reaches about 20 to 25 percent. At horizons less than six or nine months, however, the share associated with ES shocks is less than half as much. And for horizons of less than three months, the share is fairly small. These results remain essentially unchanged by the historical decompositions discussed later.

Not surprisingly, the bulk of the variance of ES is associated with its own shocks. This is particularly the case for horizons less than nine months, where its own shocks are associated with more than 80 percent of the variance.

[13] Because the shocks are orthogonal, the sum of the shares reflects the combined shares of the variance associated with shocks to DS, $\ln(y/yT)$ or $\ln(y/y_{t-12})$, and RL. Also, it avoids the thorny issue of identifying the individual shocks of these variables that are not of interest to this study.

Although this share declines somewhat at longer horizons, it remains above 60 percent.

5 External Spread Shock

Figures 2 and 3 show the impulse responses of the variables respectively in models A and B to a positive shock to ES. These impulse responses have been computed by placing ES last in the ordering. As noted earlier, this purges the identified ES shock from the impact of other shocks in the model that are more likely to reflect domestic factors. As discussed in the introduction, this experiment can be viewed as reflecting "pure" contagion or "wake-up call" effects, triggered by events taking place elsewhere. Of course, as also noted before, a more general interpretation of this experiment is possible: it can be viewed as reflecting an adverse external financial shock—related or not to contagion.[14] The figures also display one-standard error bands for each variable.[15]

[Insert Figure 2 about here]
[Insert Figure 3 about here]

As shown in the figures, a one-standard deviation shock to external spreads of roughly 120 basis points leads in the next period to an increase in

[14] In the context of Argentina during the period under consideration, the shock to external spreads that we consider may well also represent an increase in devaluation risk. In principle, accounting for the transmission process of a change in devaluation expectations would require taking into account the fact that firms had large foreign-currency denominated liabilities. But to the extent that adverse balance sheets effects translate into downward movements in the cyclical component of output—because, for instance, the risk premium depends on firms' net worth, as in Bernanke, and Gertler (1989), and Bernanke, Gertler, and Gilchrist (2000)—our empirical framework would indirectly capture them.

[15] In all figures the dotted lines for the IRs show one standard error band in each direction and are based on 1000 Monte Carlo replications. In each replication we sampled the VAR coefficients and the covariance matrix from their posterior distribution. From these replications we calculated the square root of the mean squared deviation from the impulse response in each direction. By construction, these bands contain the impulse response function but are not necessarily symmetric.

the domestic spread by only about 20 basis points in both cases. Whereas the response of the external spread lasts just over a year, the response of the domestic spread lasts for about half as long. The muted response of the domestic spread in response to the external spread shock may result from the fact that the demand for bank loans contracts as a result of a rise in external spreads; this is of course consistent with the predictions of our analytical model, in which the external cost of funds affects the (optimal) demand for variable inputs. An alternative explanation dwells on an extended version of the model presented in Section II to account for two levels of financial intermediation, as described in Agénor and Aizenman (1998). In that paper, the process of financial intermediation is viewed as consisting of two stages: foreign banks provide credit to domestic banks, and domestic banks provide the intermediation services to domestic investors. Each spread is determined by similar considerations—it equals the expected revenue lost due to partial default, and the cost of financial intermediation, at the given level of intermediation. This extended model can explain the finding reported above, if the exogenous shock to the external spread indicates that the likelihood of external default increases by more than the likelihood of internal default. This may be the case if the shock is due to contagion associated with asymmetric information—that is, if Argentina's perceived country risk by foreign lenders increased by more than the riskiness of lending operations in Argentina, as perceived by domestic banks.[16]

Movements in output become significantly negative after 2 months and display a degree of persistence that is similar to that observed for the external spread in both cases.[17] The response of the real lending rate is positive but

[16] As it turns out, Argentina did indeed default on its external debt following the collapse of its curreny board in December 2001; See Perry and Servén (2003).

[17] Note that there is a perverse blip in the output response after one month in Model B, but not in Model A. It is not clear why the measurement of cyclical output in this case makes such a difference. It is possible that the HP filter in Model A may have created a spurious cycle, as discussed by Cogley and Nason (1995). In any event, the output response does not exceed its one-standard error band before two or three months following

imprecisely measured. The initial rise in that variable is consistent with an increase in the domestic spread that is brought about through a rise in the nominal lending rate that exceeds the rise in the nominal deposit rate, with inflation displaying some degree of inertia on impact. Alternatively, it is also consistent with a situation where the fall in the cyclical component of output leads not only to a drop in both domestic rates (with the fall in the nominal deposit rate exceeding the fall in the nominal lending rate) but also to a drop in inflation, associated with a contraction in aggregate demand.

6 The Aftermath of the Peso Crisis: A Historical Decomposition

A useful application of the VAR models estimated above is to assess the movements in output and domestic interest rate spreads in Argentina in the immediate aftermath of the Mexican peso crisis of December 1994. This can be done by using the historical decompositions of these variables for the period immediately following the collapse of the Mexican peso, specifically, from January 1995 to the end of 1996. Table 2 presents these results on a quarterly basis (obtained by averaging over the monthly decompositions) for both models.

[Insert Table 2 about here]

The results for both models indicate that the fall in output in the second quarter of 1995 (by about 3 percent with respect to trend in Model A, and by about 6 percent at an annual rate in Model B) was mostly associated with the adverse effect of higher external spreads—a result that is consistent with our analytical framework. This effect persists until the first quarter of 1996 in both cases.

the shock.

Regarding the domestic spread, the conditional forecasts of the models (based on information available up to December 1994) appear to track the data fairly closely for the period under consideration. The results also suggest that for the first half of 1995, external spread shocks raised the domestic spread by about 0.4 percentage points, compared to about 2 for domestic spread shocks. Note that during the same period, the effect of external spread shocks are larger than domestic shocks. The relatively limited impact of external spread shocks on domestic spreads is consistent with the possibility that credit rationing translates into larger movements in the volume of credit, as opposed to prices. However, in the absence of disaggregated data on credit flows and pools of borrowers (based on their creditworthiness, for instance), it is hard to assess the importance of this effect. Nevertheless, it remains true that during the first part of 1995 (that is, in the immediate aftermath of the Mexican peso), external shocks had important effects on domestic bank lending spreads and (most importantly) economic activity in Argentina.

7 Summary and Conclusions

The purpose of this paper has been to study the effects of external shocks on domestic bank lending spreads and output fluctuations in Argentina. The analytical framework, which was presented in Section II, analyzed the determination of bank lending spreads in the presence of verification and enforcement costs of loan contracts. Section III presented estimates of a vector autoregression system that relates (in line with our analytical framework) the *ex ante* bank lending spread, movements in output (measured as deviations of output from trend, or the growth rate of output), the real bank lending rate, and the external interest rate spread. Variance decompositions, presented in Section IV, showed that at short horizons (less than 6 months) movements of domestic spreads are greatly influenced by domestic shocks. At longer forecast horizons, however, the external spread played a greater role in explaining these movements.

The effects of an external shock, modeled as a shock in external interest rate spreads, were analyzed in Section V using impulse response functions. The results showed that such a shock led to an increase in domestic spreads and a reduction in the cyclical component of output. Both results are consistent with the predictions of our analytical framework. The results also indicated that the response of the domestic spread with respect to the foreign spread is well below one; we argued that this prediction is consistent with an extended version of the model presented here (Agénor and Aizenman (1998)). Finally, Section VI used the VAR models to assess the effects of historical shocks to external spreads on movements in output and domestic interest rate spreads in Argentina in the immediate aftermath of the Mexican peso crisis of December 1994. The results indicated that such shocks played an important role in the behavior of both variables.

The experience of emerging markets in the nineties provides new challenges for economists, requiring us to reassess our understanding of the transmission mechanism from financial markets to real economic activity. The empirical results of our paper are consistent with the notion that financial volatility on world capital markets may have severe domestic consequences in economies where banks and debt contracts are widely used to finance production and (possibly) investment. Further research would help to validate our emphasis on costly financial intermediation and identify the policy implications of our results.

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Appendix A
The Bank Lending Spread
and the Effect of an External Shock

As noted in the text, we assume that banks have access to an elastic supply of funds, at a real cost of r^*. With competitive and risk-neutral banks, the contractual interest rate is determined by the expected break-even condition:[18]

$$(1+r^*)p_m M_i = 0.5 \left\{ (1+r_L^i)p_m M_i + \int_{\varepsilon_i^{max}}^{\Gamma} [(1+r_L^i)p_m M_i] f(\varepsilon)d\varepsilon \right. \tag{A1}$$

$$\left. + \int_{-\Gamma}^{\varepsilon_i^{max}} [\kappa M_i^{\beta}(1+\delta_0-\nu+\varepsilon) - c_i p_m M_i(1+r^*)] f(\varepsilon)d\varepsilon \right\},$$

where $f(\varepsilon)$ is the density function. Using (3) and (A1), the interest rate spread can be shown to be given by

$$r_L^i - r^* = \frac{0.5\int_{-\Gamma}^{\varepsilon_i^{max}} [\kappa M_i^{\beta}(\varepsilon_i^{max}-\varepsilon)] f(\varepsilon)d\varepsilon}{p_m M_i} + \frac{0.5 c_i p_m M_i (1+r^*)\int_{-\Gamma}^{\varepsilon_i^{max}} f(\varepsilon)d\varepsilon}{p_m M_i}. \tag{A2}$$

The contractual interest rate, r_L^i, is determined by a mark-up rule. r_L^i exceeds the bank's cost of funds, r^*, by the sum of two terms: the first is the expected revenue lost due to partial default in bad states of nature, and the second measures the expected state verification and contract enforcement costs.

The producer's expected net income equals

$$(1+\delta_0)M_i^{\beta} - 0.5 \left\{ \begin{array}{c} (1+r_L^i)p_m M_i + \int_{\varepsilon_i^{max}}^{\Gamma} [(1+r_L^i)p_m M_i] f(\varepsilon)d\varepsilon \\ + \int_{-\Gamma}^{\varepsilon_i^{max}} [\kappa M_i^{\beta}(1+\delta_0-\nu+\varepsilon)] f(\varepsilon)d\varepsilon \end{array} \right. . \tag{A3}$$

Using (A1), we can simplify (A3) to

$$(1+\delta_0)M_i^{\beta} - (1+r^*)p_m M_i - 0.5 c_i p_m M_i (1+r^*)\int_{-\Gamma}^{\varepsilon_i^{max}} f(\varepsilon)d\varepsilon. \tag{A4}$$

[18] In what follows we drop the subscript i on ε to simplify notations.

The optimal level of use of the variable input, M_i, is found by maximizing (A4). It can readily be shown to depend negatively on $1 + r^*$.

In the particular case in which the idiosyncratic shock follows a uniform distribution, $-\Gamma \leq \varepsilon < \Gamma$, the spread (A2) is characterized by a quadratic equation, given by

$$r_L^i - r^* = 2\Gamma \frac{\kappa M_i^\beta \Phi_i^2}{p_m M_i} + c_i(1 + r^*)\Phi_i, \tag{A5}$$

where $\Phi_i = (\Gamma + \varepsilon_i^{max})/4\Gamma$ is the probability of default. Combining the above equation with (3) one can infer a reduced form solution for the probability of default and for the domestic interest rate.

To establish the derivations in Proposition I proceeds as follows. Using (3) and (A5), we infer that the probability of default is determined by

$$2\Gamma\kappa M_i^\beta \Phi_i^2 + \left\{ c_i(1 + r^*)p_m M_i - 4\kappa M_i^\beta \Gamma \right\} \Phi_i + (1 + r^*)p_m M_i \tag{A6}$$

$$-\kappa M_i^\beta (1 + \delta_0 - \nu - \Gamma) = 0.$$

This is a quadratic equation, yielding 2 interest rates in the relevant range. Henceforth we assume that competitive forces induces banks to offer the lower interest rate, leading to a probability of default of

$$\Phi_i = \frac{H - \sqrt{Z}}{4\kappa M_i^\beta \Gamma}, \tag{A7}$$

where

$$H = 4\kappa M_i^\beta \Gamma - c_i(1 + r^*)p_m M_i, \quad Z = H^2 - 8\kappa M_i^\beta \Gamma \Lambda,$$

$$\Lambda = (1 + r^*)p_m M_i - \kappa M_i^\beta (1 + \delta_0 - \nu - \Gamma).$$

Using (A6) and (3), we infer that

$$dr_L^i / dr^* = 4\kappa M_i^\beta \Gamma / \sqrt{Z}. \tag{A8}$$

Hence, we operate on the upward-slopping portion of the supply of credit as long as $H > \sqrt{Z}$ and $Z \geq 0$. We approach the backward-bending part of the curve as $Z \to 0$. Henceforth we assume that this condition holds.

The first-order condition determining the demand for the variable input is inferred from (A4) as

$$\frac{d\Pi}{dM_i} = (1+\delta_0)\beta M_i^{\beta-1} - (1+r^*)p_m c_i[\Phi_i + M_i(\frac{\partial \Phi_i}{\partial M_i})] = 0. \qquad \text{(A9)}$$

Applying the implicit function theorem to (A9), and using the second order-condition for profits maximization, we infer that

$$sg[\frac{dM_i}{dr^*}] = -sg[\frac{d^2\Pi/(dxdM_i)}{d^2\Pi/dM_i^2}] = sg[\frac{d^2\Pi}{dr^* dM_i}]. \qquad \text{(A10)}$$

This result implies that, to establish that $dM_i/dr^* < 0$, it suffices to show that $d^2\Pi/(dxdM_i) < 0$. Applying (A9) we infer that

$$\frac{d^2\Pi}{dr^* dM_i} = -\frac{(1+\delta_0)\beta M_i^{\beta-1}}{1+r^*} - (1+r^*)p_m c_i[\frac{\partial \Phi_i}{\partial r^*} + M_i(\frac{\partial^2 \Phi_i}{\partial M_i \partial r^*})]. \qquad \text{(A11)}$$

Applying (A7), and collecting terms, it follows that

$$\frac{\partial \Phi_i}{\partial r^*} = \frac{M_i}{\sqrt{Z}}[1 + \frac{c_i(H-\sqrt{Z})}{4\kappa M_i^\beta \Gamma}] = \frac{M_i}{\sqrt{Z}}(1 + c_i\Phi_i). \qquad \text{(A12)}$$

$$\frac{\partial^2 \Phi_i}{\partial M_i \partial r^*} = \frac{1+(1-\beta)c_i\Phi_i}{\sqrt{Z}} - \frac{M_i(\partial Z/\partial M_i)}{2Z\sqrt{Z}}[1+\frac{c_i H}{4\kappa M_i^\beta \Gamma}] + \frac{c_i}{\sqrt{Z}}[\beta - \frac{(1+r^*)c_i}{4\kappa M_i^\beta \Gamma}]$$

Thus,

$$\frac{\partial \Phi_i}{\partial r^*} + M_i(\frac{\partial^2 \Phi_i}{\partial M_i \partial r^*}) =$$

$$\frac{M_i}{\sqrt{Z}}\left\{2 + (2-\beta)c_i\Phi_i + c_i[\beta - \frac{(1+r^*)c_i}{4\kappa M_i^\beta \Gamma}] - \frac{M_i(\partial Z/\partial M_i)}{2Z}[1 + \frac{c_i H}{4\kappa M_i^\beta \Gamma}]\right\}$$

Using (A7) it can be shown that $M_i(\partial Z/\partial M_i)/2Z < 1$ and $c_i H/4\kappa M_i^{\beta}\Gamma > c_i\Phi_i$. Applying these 2 results to the above equation it can be verified that

$$\frac{\partial \Phi_i}{\partial r^*} + M_i(\frac{\partial^2 \Phi_i}{\partial M_i \partial r^*}) \geq 0,$$

from which we infer that, indeed, $d^2\Pi/dr^* dM_i < 0$. A technical note (available upon request) establishes that lower expected productivity, δ_0, and higher volatility of macroeconomic shocks, ν, raise domestic interest rates and the bank lending spread, and reduces expected output.

Appendix B
Data Sources and VAR Estimation

Data. The data used in this study are at a monthly frequency and cover the period 1993:m6-1998:m6. The variables are measured as follows:[19]

- ES is the external spread of Brady par bonds over U.S. Treasury bills. The series is virtually indistinguishable from spreads on Brady discounted bonds, and its movements are highly correlated with external spread on sovereign bonds (as shown in Figure 1). Data were obtained from Merryll Lynch.

- DS is calculated as the difference between the nominal lending rate on peso-denominated loans and the deposit rate on peso-denominated deposits. The series were obtained from the Fund's *International Financial Statistics* (line 60p and line 60l) and from Catão (1998).

- RL is calculated as the nominal lending rate on peso-denominated loans at a monthly rate minus current monthly inflation, measured by the consumer price index. Raw series were obtained from the Fund's *International Financial Statistics.* (lines 60p and 64)

- $\ln(y/y_T)$ measures deviations of industrial output, y, from trend, y_T. y_T is estimated with the Hodrick-Prescott filter, using a value of $\lambda = 16000$ for the smoothing parameter. $\ln(y/y_{-12})$ is the growth rate of output. The industrial output index was obtained from FIEL.

VAR estimation. To determine the number of lags to include in the VAR models, we started by using standard lag-length tests, that is Akaike Information Criteria (AIC), Hannan-Quinn (HQ), and Schwarz. These tests

[19] The effective reserve requirement rate, which was used in our preliminary experiements, was calculated by subtracting line 14a in the IMF's *International Financial Statistics* from line 14 and dividing by the sum of lines 24 and 25, minus line 14a.

compare the cost of increasing the lag length (a reduction in degrees of freedom) to the benefit of doing so (increased information extraction from the data). Using a maximum lag length of six, all three tests suggested using six lags. This presented a problem due to the size of the sample: using six lags would imply that each of the five equations contains 31 (6*5+1) coefficients to estimate with 66 monthly observations (January 1993-June 1998). Low degrees of freedom would therefore translate into low precision in the estimation.

Instead of using six lags as suggested by the tests, we use three lags based on two considerations. First, it is the smallest lag length where the reduced-form innovations are white noise judging by Ljung-Box Q tests for serial correlation (up to order 12). This ensures that the white noise assumption implicit in the estimation procedure is not violated. Second, and more importantly, the impulse responses and the variance decompositions using three lags are qualitatively the same as those using six lags. Thus, using the shorter lags does not affect the main qualitatively results presented in the paper. Table A1 presents a summary of the estimated VAR equations that underlie the empirical results in the paper.

[Insert Table A1 about here]

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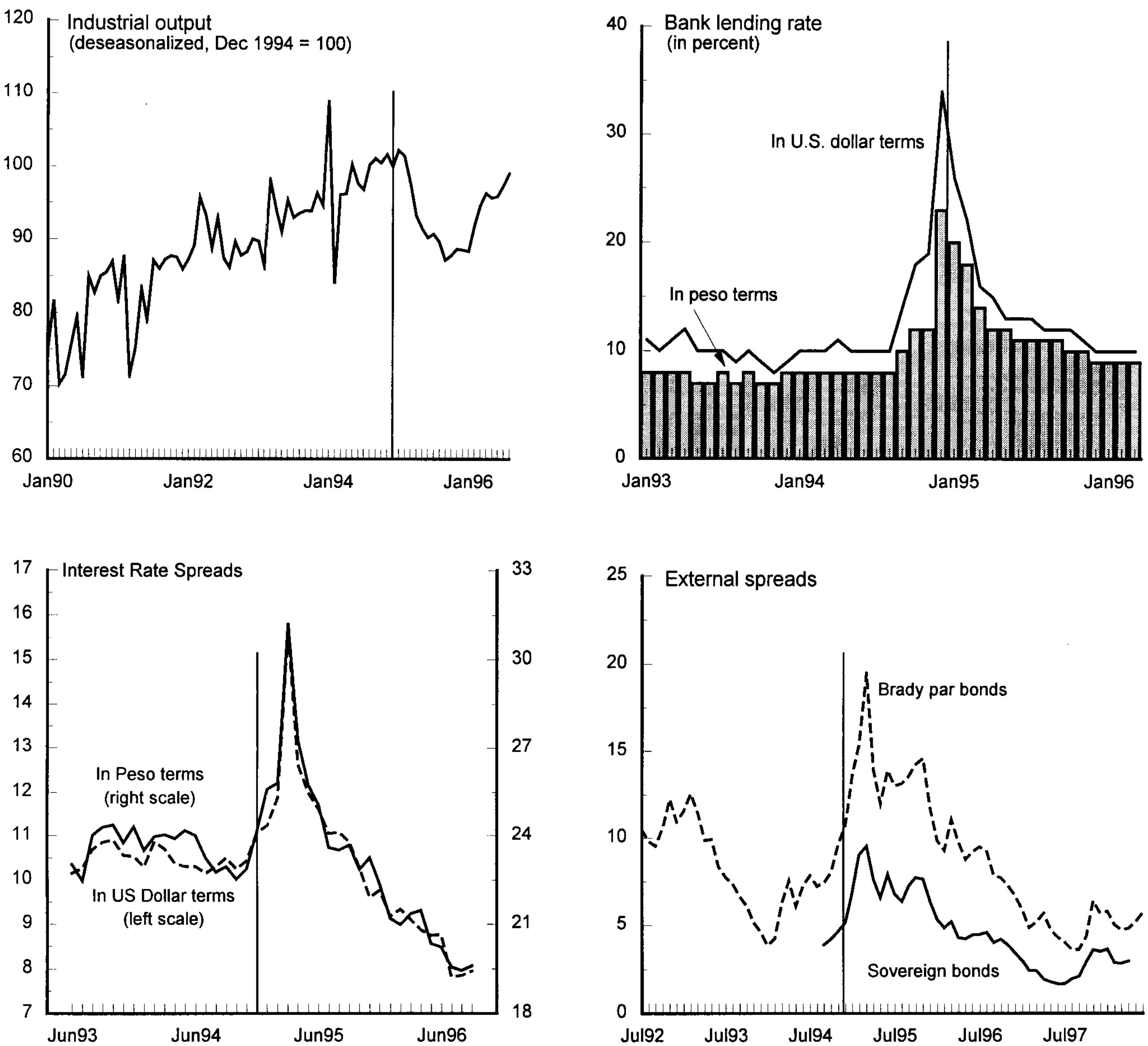
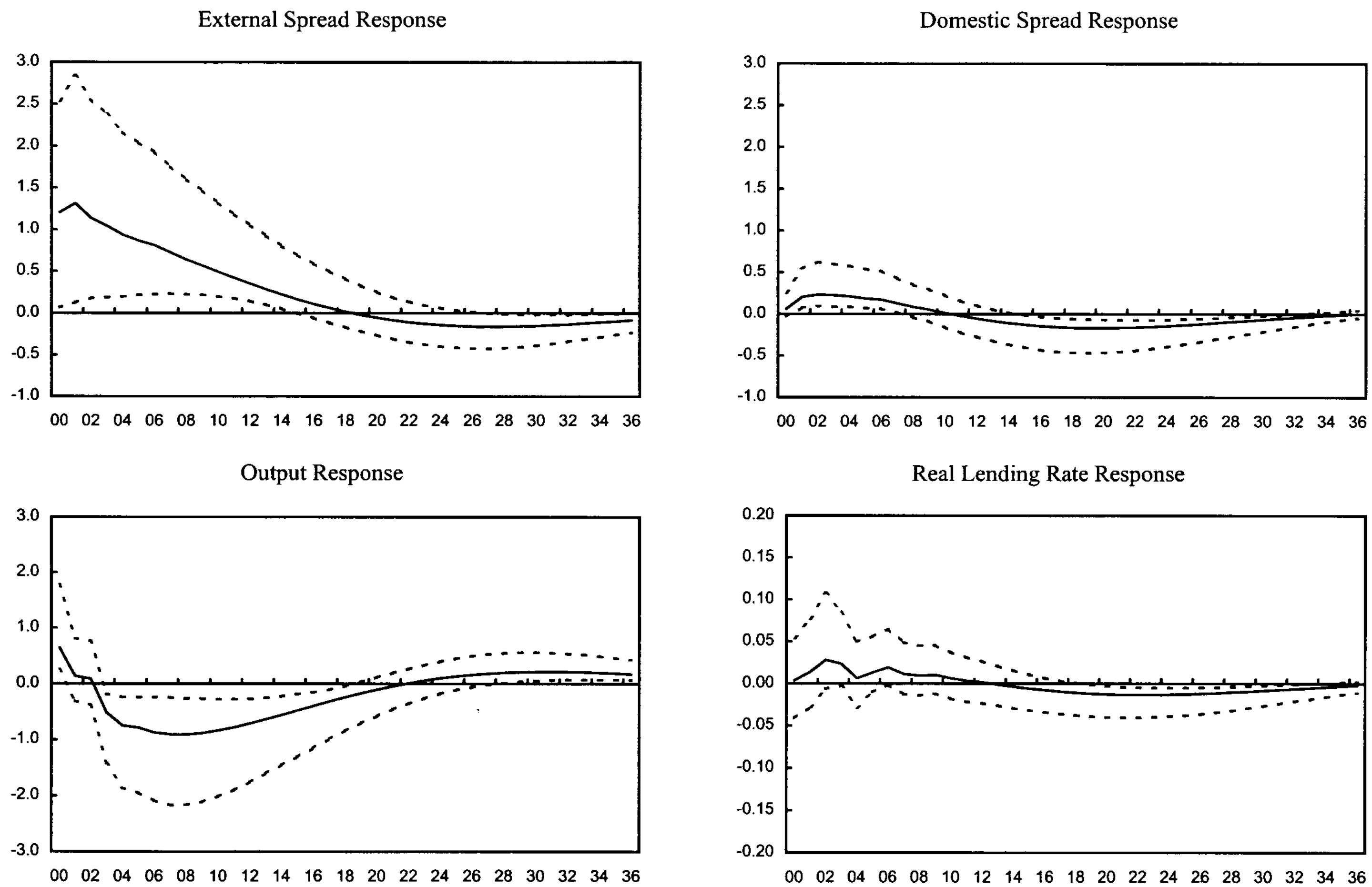
Figure 1
Argentina: Output and Interest Rates 1/



Sources: FIEL; International Monetary Fund, Bloomberg, Inc., and Merryll Lynch.

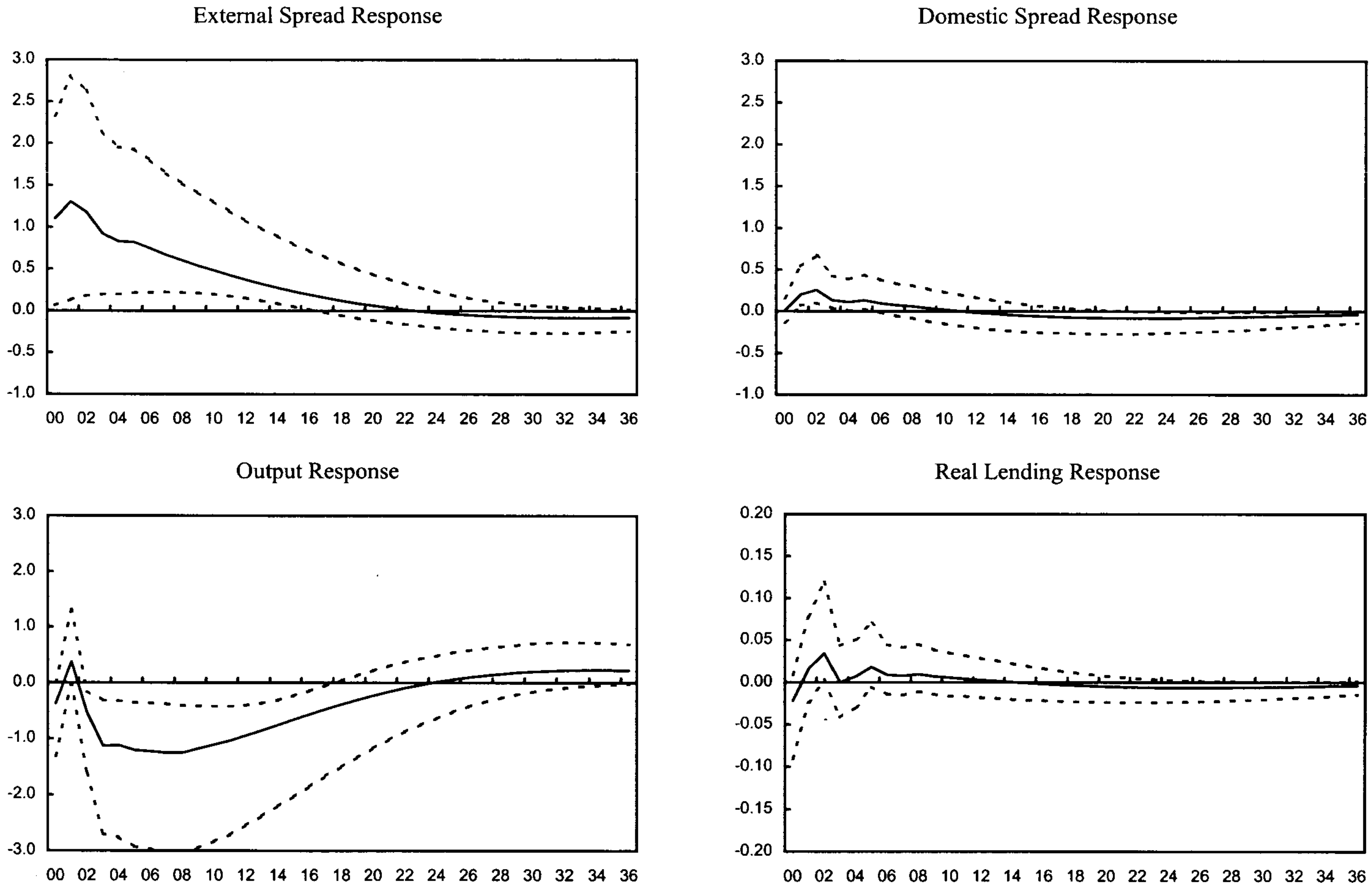
1/ The vertical line corresponds to the Mexican peso crisis (December 20, 1994).

Figure 2. Generalized Impulse Responses, Model A.
(Historical Shock to the External Spread)



Note: The impulse responses were obtained from a VAR model with four variables: the external spread, the domestic spread, output (deviation from trend), and the real lending rate; all variables are measured in percentage points except output which is measured as the percentage deviation from trend output. The shock to the external spread equals the standard deviation of its VAR innovation, 120 basis points. The VAR model is estimated with three lags using monthly data from 1993:M1 through 1998:M6. One standard error band in each direction are based on 1,000 Monte Carlo replications. See appendix for details.

Figure 3. Generalized Impulse Responses, Model B.
(Historical Shock to the External Spread)



Note: The impulse responses were obtained from a VAR model with four variables: the external spread, the domestic spread, output ($\log(y_t/y_{t-12})$), and the real lending rate; all variables are measured in percentage points. The shock to the external spread equals the standard deviation of its VAR innovation, 120 basis points. The VAR model is estimated with three lags using monthly data from 1993:M1 through 1998:M6. One standard error bands in each direction are based on 1,000 Monte Carlo replications. See appendix for details.

Table 1. Generalized Variance Decompositions

Months	Model A: External Spread (ES) — ES	DS	$\ln(y/y_T)$	RL	Model B: External Spread (ES) — ES	DS	$\ln(y_t/y_{t-12})$	RL
	Percentage of the variance associated with historical shocks to:				Percentage of the variance associated with historical shocks to:			
1	100.0	0.4	2.9	0.0	100.0	0.0	0.8	0.4
2	99.5	0.4	5.2	0.1	98.6	0.3	0.5	0.2
3	95.9	1.5	3.8	1.8	92.8	2.0	4.5	3.1
6	92.8	5.0	6.1	1.3	89.2	6.7	3.6	2.2
9	87.4	8.2	8.4	1.1	83.3	11.9	2.9	2.1
12	81.4	10.3	10.3	1.6	76.9	16.0	2.5	2.6
24	70.5	11.2	12.2	4.7	61.8	20.8	1.9	6.7

Months	Model A: Cyclical Component of Output ($\ln(y/y_T)$) — ES	DS	$\ln(y/y_T)$	RL	Model B: Output Growth ($\ln(y_t/y_{t-12})$) — ES	DS	$\ln(y_t/y_{t-12})$	RL
	Percentage of the variance associated with historical shocks to:				Percentage of the variance associated with historical shocks to:			
1	2.9	0.4	100.0	5.7	0.8	0.0	100.0	9.2
2	2.8	3.0	94.1	6.1	1.4	0.0	96.1	9.9
3	2.6	5.1	87.6	10.0	2.7	2.9	92.4	9.4
6	9.7	5.9	78.2	9.5	15.5	11.2	72.6	7.2
9	19.1	8.0	68.3	8.5	23.3	19.8	55.3	5.8
12	24.8	10.3	61.4	7.5	25.6	25.4	43.6	5.8
24	25.2	12.5	53.0	8.6	20.9	30.1	28.0	10.7

Months	Model A: Domestic Spread (DS) — ES	DS	$\ln(y/y_T)$	RL	Model B: Domestic Spread (DS) — ES	DS	$\ln(y_t/y_{t-12})$	RL
	Percentage of the variance associated with historical shocks to:				Percentage of the variance associated with historical shocks to:			
1	0.4	100.0	0.4	13.0	0.0	100.0	0.0	10.5
2	4.1	93.9	3.6	11.3	3.5	92.3	1.8	8.6
3	6.5	76.2	4.0	14.0	6.3	75.9	2.6	13.6
6	9.0	60.9	8.6	12.4	5.9	67.6	2.2	13.3
9	8.4	51.2	11.0	14.5	5.2	61.2	1.7	15.6
12	7.4	45.9	11.9	16.8	4.5	56.9	1.5	17.8
24	11.6	40.6	11.3	18.7	5.0	50.9	1.4	21.0

Months	Model A: Real Lending Rate (RL) — ES	DS	$\ln(y/y_T)$	RL	Model B: Real Lending Rate (RL) — ES	DS	$\ln(y_t/y_{t-12})$	RL
	Percentage of the variance associated with historical shocks to:				Percentage of the variance associated with historical shocks to:			
1	0.0	13.0	5.7	100.0	0.4	10.5	9.2	100.0
2	0.1	12.6	8.7	98.6	0.5	10.6	10.5	99.2
3	0.7	14.0	9.0	93.2	1.3	12.6	13.9	91.2
6	1.1	16.0	9.7	90.4	1.4	15.1	14.6	88.3
9	1.4	17.0	10.2	87.8	1.5	16.7	14.2	85.9
12	1.4	17.2	10.4	86.1	1.5	17.4	13.9	84.2
24	2.0	16.9	10.4	84.2	1.6	17.7	13.3	82.1

Note: These decompositions are based on the generalized VAR analysis following Koop, Pesaran and Potter (1996) who propose to consider non-orthogonal historical shocks. Consequently the variance decompositions do not add up to 100 percent. The variance decompositions are obtained from VAR models comprised by the following variables: ES, DS, ln(y/yT) in Model A and ln(yt/yt-12) in Model B, and RL.
The model is estimated with three lags using monthly data from 1993:M1 through 1998:M6; see Appendix B for details.

Table 2. Generalized Historical Decompositions

Quarter	Model A: Cyclical Component of Output ($\ln(y/y_T)$)						Model B: Output Growth ($\ln(y_t/y_{t-12})$)					
	Actual	Model projection	Associated with historical shocks to: ES	DS	$\ln(y/y_T)$	RL	Actual	Model projection	Associated with historical shocks to: ES	DS	$\ln(y_t/y_{t-12})$	RL
1995:Q1	0.064	0.015	0.016	0.003	0.036	0.011	0.043	0.028	-0.007	-0.004	0.011	0.004
1995:Q2	-0.031	-0.004	-0.035	-0.004	0.016	-0.009	-0.063	0.017	-0.048	-0.028	0.003	0.000
1995:Q3	-0.089	-0.010	-0.030	-0.013	-0.032	-0.010	-0.109	0.020	-0.042	-0.053	-0.025	-0.005
1995:Q4	-0.104	-0.009	-0.024	-0.017	-0.053	-0.004	-0.128	0.030	-0.034	-0.065	-0.055	0.004
1996:Q1	-0.029	-0.004	-0.026	-0.018	0.021	0.004	-0.075	0.043	-0.031	-0.067	-0.028	0.019
1996:Q2	0.003	0.003	-0.010	-0.025	0.035	-0.007	0.058	0.056	-0.005	-0.068	0.059	0.011
1996:Q3	-0.020	0.010	-0.007	-0.031	-0.012	0.003	0.101	0.068	0.002	-0.064	0.065	0.016
1996:Q4	-0.031	0.016	-0.001	-0.020	-0.041	0.015	0.114	0.078	0.000	-0.041	0.047	0.033

Quarter	Model A: Domestic Spread (DS)						Model B: Domestic Spread (DS)					
	Actual	Model projection	Associated with historical shocks to: ES	DS	$\ln(y/y_T)$	RL	Actual	Model projection	Associated with historical shocks to: ES	DS	$\ln(y_t/y_{t-12})$	RL
1995:Q1	0.190	0.167	0.006	0.014	0.001	0.004	0.190	0.163	0.006	0.016	0.002	0.004
1995:Q2	0.186	0.165	0.008	0.012	0.005	0.001	0.186	0.159	0.005	0.023	-0.001	0.000
1995:Q3	0.172	0.161	0.003	0.010	0.005	-0.005	0.172	0.155	0.002	0.018	-0.001	-0.005
1995:Q4	0.162	0.156	0.002	0.013	-0.001	-0.006	0.162	0.152	0.002	0.015	0.000	-0.006
1996:Q1	0.150	0.152	-0.004	0.016	-0.005	-0.005	0.150	0.149	-0.005	0.014	0.001	-0.003
1996:Q2	0.145	0.150	-0.005	0.012	0.000	-0.010	0.145	0.147	-0.003	0.015	0.000	-0.009
1996:Q3	0.133	0.148	-0.005	-0.004	0.003	-0.011	0.133	0.146	-0.002	-0.002	-0.002	-0.010
1996:Q4	0.132	0.148	-0.005	-0.004	0.000	-0.006	0.132	0.145	-0.002	-0.005	-0.002	-0.006

Note: These historical decompositions are calculated by averaging the monthly decompositions using generalized VAR analysis. The VAR model projections are obtained as dynamic forecasts of the models conditional on information up to December 1994. Since historical shocks are not orthogonal, the model projections and the effects associated with each historical shock do not add up to the actual series.

Table A1. VAR Estimates, Monthly Observations from January 1993 to June 1998.

Model A	ES	DS	$\ln(y/y_T)$	RL
Coefficient of Determination (R^2)	0.883	0.788	0.524	0.326
Adjusted R^2	0.852	0.731	0.397	0.146
Sum of Squared Errors	84.094	54.854	832.493	7.059
Standard Error of Estimate	1.367	1.104	4.301	0.396
Significance of Lagged Regressors:				
External Spread	64.582 *	0.810	1.494	0.111
Domestic Spread	1.474	30.049 *	1.316	1.325
Output	0.707	1.676	2.804 *	0.505
Real Lending Rate	2.148	3.596 *	1.214	3.105 *
Correlation with the VAR innovations of:				
External Spread	1.450	0.062	**0.171**	0.011
Domestic Spread		0.946	0.059	0.361
Output			14.353	0.239
Real Lending Rate				0.122
Tests for Serial Correlation:				
Breusch-Godfrey	64.89	52.62	9.95	8.84
Ljung-Box Q	91.93	97.12	54.63	56.71

Model B	ES	DS	$\ln(y_t/y_{t-12})$	RL
Coefficient of Determination (R^2)	0.902	0.776	0.714	0.339
Adjusted R^2	0.876	0.717	0.637	0.163
Sum of Squared Errors	0.007	0.006	0.108	0.001
Standard Error of Estimate	0.013	0.011	0.049	0.004
Significance of Lagged Regressors:				
External Spread	61.986 *	1.001	1.843	0.133
Domestic Spread	1.693	33.823 *	1.708	1.422
Output	3.710 *	0.837	5.458 *	0.819
Real Lending Rate	1.487	3.116 *	0.721	2.839 *
Correlation with the VAR innovations of:				
External Spread	1.217	0.009	-0.087	-0.064
Domestic Spread		0.996	-0.016	0.323
Output			18.555	0.304
Real Lending Rate				0.119
Tests for Serial Correlation:				
Breusch-Godfrey	35.59	38.15	1.34 *	18.35
Ljung-Box Q	89.21	97.03	34.00	11.58

Note: The VAR models are estimated with three lags. The significance tests are F-tests for the joint significance of all of the lags of the corresponding variable; these tests have respectively three and 53 degrees of freedom in the numerator and the denominator. The tests for serial correlation test for serial correlations of up to order 12. An asterisk (*) denotes significant rejection of the respective null hypothesis at the five percent significance level.