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Contagion, Bank Lending Spreads, and Output Fluctuations

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A positive historical shock to external spreads can lead to an increase in domestic spreads and a reduction in the cyclical component of output. Shocks to external spreads immediately after the Mexican peso crisis had a sizable effect on movements in output and domestic interest rate spreads in Argentina.

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Summary findings

Agénor, Aizenman, and Hoffmaister study how contagion affects bank lending spreads and fluctuations in output in Argentina.

They analyze what determines bank lending spreads when verification and enforcement costs for loan contracts are high.

They present estimates of a vector autoregression model that relates bank lending spreads, the cyclical component of output, the real bank lending rate, and the spread in external interest rates.

Using generalized impulse response functions, they show that a positive historical shock to external spreads leads to an increase in domestic spreads and a reduction in the cyclical component of output.

Historical decompositions indicate that shocks to external spreads immediately after the Mexican peso crisis had a sizable effect on movements in output and domestic interest rate spreads in Argentina.

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Contagion, Bank Lending Spreads, and Output Fluctuations

P. R. Agénor, J. Aizenman and A. Hoffmaister*

Abstract

This paper studies the effects of contagion on bank lending spreads and output fluctuations in Argentina. The first part presents the analytical framework, which analyzes the determination of bank lending spreads in the presence of verification and enforcement costs of loan contracts. The second part presents estimates of a vector autoregression model that relates bank lending spreads, the cyclical component of output, the real bank lending rate, and the external interest rate spread. Using generalized impulse response functions, a positive historical shock in external spreads is shown to lead to an increase in domestic spreads and a reduction in the cyclical component of output—as predicted by our analytical framework. Historical decompositions indicate that shocks to external spreads in the immediate aftermath of the Mexican peso crisis had a sizable effect on movements in output and domestic interest rate spreads in Argentina.

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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.

The timing and severity of the economic downturn in Argentina was associated with an adverse external financial 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. Our analysis follows this perspective and models contagion 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. This view is, of course, also consistent with a more general interpretation of external shocks. It is reflected in 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). The real effects of this shock 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 (an important feature of Argentina's financial system), domestic interest rate spreads, and real lending rates.[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 a recent 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 this result as reflecting the fact that most peso borrowers in Argentina cannot arbitrage between domestic and foreign sources of funds, and thus become 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

[1] As documented for instance by Rojas-Suárez and Weisbrod (1995), banks account for between 50 and 90 percent of the financing needs of firms in Latin American countries. Agénor (1998), Edwards and Végh (1997), Greenwald and Stiglitz (1993), and Isard et al. (1996) also develop models which explicitly account for the link between firms' working capital needs and bank credit.

[2] Among recent studies of the determinants of bank spreads are Barajas, Steiner, and Salazar (1998) for Colombia, Demirgüç-Kunt and Huizinga (1998) for a large group of countries. Early studies include Ho and Sanders (1981), and Hanson and de Resende Rocha (1986).

[3] Catão uses, 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 incidence of nonperforming loans is high and refinancing operations are widespread.

analytical framework, which describes 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 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. Generalized variance decompositions are discussed in Section IV. Section V uses generalized 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.

2 The Analytical Framework

The credit channel provides a key transmission mechanism of macroeconomic shocks in developing countries. This channel impacts directly 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 macroeconomic shocks and volatility 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 external cost of credit shocks on domestic output.[4]

[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

We consider an economy where risk-neutral banks provide intermediation services. Agents (producers) demand credit from banks (lenders) to finance their working capital needs. Producers who lack access to the equity market rely 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 is costly; second, enforcing repayment requires costly intervention of the legal system.

Future 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 consist of labor or raw materials) 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 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. We assume that each producer must finance variable input costs prior to the sale of output, and that no one can issue claims on his or her capital stock. Consequently, producer i's variable costs are $(1 + r_L^i)p_m M_i$, where p_m is the relative price of the variable input.

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

the external debt literature, as in Eaton et al. (1986), Bulow and Rogoff (1989), and Helpman (1989).

number of independent producers, allowing 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] A producer will default if

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

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:[6]

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

where $0<c_i<1$.

[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 significant enough.

[6]The cost c_i is paid by banks in order to identify the productivity shock ε_i, and to enforce proper 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.

We assume that banks have access to an elastic supply of funds, at a real cost of r^*.[7] Assuming that banks are risk neutral and competitive, the contractual interest rate is determined by an expected break-even condition, derived in Appendix I. 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 in which the aggregate 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 I). Henceforth 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 supplied and where the probability of default is positive, the following proposition can be shown to hold:

Proposition 1 *A higher external cost of credit, r^* raises domestic interest rates and the bank lending spread, and reduces expected output.*

[7] This source of funds may be credit provided by foreign banks, as modeled by Agénor and Aizenman (1998).

[8] Appendix I also derives the producer's expected net income, and indicates that 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.

As discussed in Appendix I, the magnitude of these effects increases with the responsiveness of the domestic interest rate to the cost of funds for banks, $\partial r_L^i / \partial r^*$, and are maximized as we approach the backward-bending portion of the supply of credit facing producers.

3 VAR Estimation and Analysis

We now apply the analytical framework developed above to an analysis of Argentina's experience 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.[10]

Both models are estimated with monthly data from January 1993 through June 1998. In addition to the variables listed above, we considered expanded VAR models with the average effective reserve requirement rate, in an attempt to control for changes in the cost of financial intermediation.[11] Although reserve requirement rates did change significantly during the sample

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

[11] 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 such factors fluctuate relatively little within the sample period.

period, the results obtained from this expanded model were not qualitatively different from the 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 II) was set to three months.

To provide empirical evidence on the analytical framework, we use generalized VAR techniques that are based on reconsidering what impulse response functions (IRFs) and variance decompositions are meant to uncover. Koop, Pesaran, and Potter (1996) and Pesaran and Shin (1998) argue that the notion of IRFs (and variance decompositions) should be re-examined, and proposed to change the focus from "pure" structural shocks identified by orthogonalizing VAR innovations, to an understanding of what a historical innovations suggests about the dynamics in the data. Typically, these historical innovations are not orthogonal but, contrary to standard VAR regression residuals (or innovations), they embody the full information of the contemporaneous correlation of these residuals. This makes them particularly useful for studying the implications of the analytical framework presented above; using standard VAR techniques would require in the present context arbitrary timing assumptions about the contemporaneous interactions between the variables. Although our model does suggest that shocks to external spreads should be considered first in the ordering of the variables in a standard VAR system, it does not lend itself to a natural ordering for the remaining ones, because they are jointly determined. Moreover, as noted below, the historical shocks to the external spread are numerically equivalent to the standard VAR innovations when the external spread is placed first in the ordering. A further consideration is that generalized IRFs (and generalized variance decompositions) are unique, that is, not subject to ordering/compositional effects of standard orthogonal analysis.

To illustrate generalized VAR analysis, consider the moving average representation of the a VAR model:

$$Z_t = C(L)^{-1}\mu_t, \tag{5}$$

where $Z_t' = [ES_t,\ DS_t,\ RL_t,\ \ln(y_t/y_T)]$ and μ_t is distributed (multivariate) normal, that is, $N(0, \Omega)$. This implies that Z_t is also normal with zero mean and covariance matrix $C(L)^{-1}\Omega C(L)^{-1\prime}$. Rather than orthogonalizing the VAR innovations in equation (5), generalized VAR analysis considers the conditional expectation of Z_t given a specific shock to μ_t.

For the sake of argument, consider the generalized (or average) effect on Z_t of the historical shock that is of particular interest in this study, a shock to the ES_t, specifically $\mu_{ES,t}$. This effect is obtained by taking the expectation of equation (5) conditional on the shock $\mu_{ES,t} = v$,

$$GIR(Z_t, \mu_{ES,t} = v) = E[Z_t \mid \mu_{ES,t} = v, \Omega] = C(L)^{-1}E[\mu \mid \mu_{ES,t} = v, \Omega],$$

and given the properties of the multivariate normal distribution:

$$GIR(Z_t, \mu_{ES,t} = v) = C(L)^{-1}\Omega_{ES} \cdot \sigma_{ES,ES}^{-1} \cdot v, \tag{6}$$

where Ω_{ES} is the column of Ω corresponding to ES, and $\sigma_{ES,ES}$ is the variance of the innovation in ES. Note that although v could be any value, it seems natural to set it equal to its historical value: the standard error of the ES shock, $\sigma_{ES,ES}^{1/2}$.

In general, the GIR in equation (6) will differ from the standard impulse responses. However, the GIR in (6) will be numerically equivalent to the Choleski decompositions when the ES is placed first in the ordering, or in the special case when the innovations in μ_t are mutually orthogonal. Aside from these numerical equivalencies, generalized VAR analysis is a conceptually different construct. Generalized VAR analysis is intended to reveal to the analyst how the VAR model behaves following a specific historical shock. Likewise, generalized variance decompositions are intended to provide the "share" of the movements of a specific series associated with historical shocks. Neither generalized IRFs nor generalized variance decompositions intend to uncover the effect of "structural" shocks as in standard VAR analysis, and thus historical shocks are not orthogonal. In this sense, generalized VAR analysis provides "stylized facts" about the VAR model that fully accounts from the dynamics and historical correlations present in the data, that in turn can be compared to the predictions of the analytical framework discussed above. Note that the fact that the generalized shocks are not orthogonal implies that the variance decompositions do not generally add up to 100 percent.

4 Generalized Variance Decompositions

Table 1 presents the generalized variance decompositions (GVDs) for the variables in the system, for both models. At a forecast horizon of less than twelve months, movements in the external spread are primarily associated with their own historical shocks for both models. At longer forecast horizons, historical shocks associated the domestic spread play a more substantial role in both cases; in addition, movements in output play a greater role in model A. These results are consistent with the analytical framework presented above and the extended framework developed by Agênor and Aizenman (1998), in which external spreads have an endogenous component—reflecting the probability of default of domestic producers on their liabilities to domestic banks and the risk of domestic banks defaulting on their foreign loans.[12] More generally, the results obtained with Model A are also consistent with the view that domestic economic conditions affect movements in external spreads through their effect on market sentiment or expectations.

At short horizons, movements of domestic spreads are also greatly influenced by their own historical shocks in both models. At a forecast horizon of

[12]Note also that the analytical framework predicts that shocks to output and domestic spreads are correlated, because the latter variable reflects the probability of default (which is itself related to output shocks). Recall, however, that since generalized VAR analysis focuses on nonorthogonal shocks, it is not valid to add up their shares to obtain a measure of their combined effect.

less than six months, these shocks explain the bulk of the movements in DS. At longer forecast horizons (beyond 6 months), historical shocks to the real lending rate play a greater role in explaining these movements, again in both models. In addition, this is also true for the external spread and movements in output in Model A—with each accounting for about the same share of the movements in the domestic spread at a horizon of 24 months. Again, these results are consistent with the theoretical framework discussed above and our main proposition; shocks to both external spreads and output affect the capacity of domestic firms to repay, thereby raising banks' perceived risk of default.

As is the case with external and domestic spreads, movements in output are mostly explained by their own historical shocks at forecast horizons less than six months in both models, explaining in excess of 75 percent of its movements. At longer forecast horizons, historical shocks for both models associated with the external spread play a substantial secondary role, accounting between 20 to 25 percent of cyclical movements in output at a forecast horizon of 24 months. Shocks to domestic spreads explain a similar portion of movements in output in Model B but account for about half as much as shocks to external spreads in Model A. A rather puzzling fact is that shocks to the real lending rate account for a relatively small proportion of cyclical movements in output in both models.[13] One possible explanation is that our index of output (industrial production) reflects essentially output of traded goods; to the extent that producers of traded goods have a greater access to world capital markets (because of their ability to post collateral in foreign currency terms), one would expect a limited effect of the cost of borrowing on domestic capital markets.

Movements in the real lending rate in both models are greatly influenced by their own historical shocks at all horizons, explaining in excess of 80 per-

[13]We attempted to measure the real lending rate by using various proxies for the expected inflation rate (lagged, current and one-period ahead actual values). This did not change significantly our results.

cent of its movements. Shocks to domestic spreads and to the real lending rate play a secondary and tertiary role respectively in accounting for movements in the real lending rate, accounting respectively about 15 and 10 percent of the movements in the real lending rate and shocks to the cycle.

5 External Spread Shock

Figures 2 and 3 show the generalized impulse responses (GIRs) for the variables of the two models to a positive historical shock in the external spread. As discussed in the introduction, this experiment can be viewed as one way of capturing "pure" (expectations-related) contagion effects, triggered by events taking place elsewhere in the region or the world. Of course, as also noted earlier, a more general interpretation of this experiment is possible; it can be viewed simply as reflecting an adverse external financial shock—related or not to contagion.

GIRs and their one-standard error bands are shown for each variable.[14] As indicated earlier, GIRs are obtained as the conditional distribution of each shock in the system and thus provide the dynamic responses of the variables in the VAR model that accounts for all of the historical information in the data sample. As illustrated in equation (6), the GIR to an external spread shock shows the evolution of variables in the model corresponding to "historically correct" shock to external spreads that explicitly accounts for all the contemporaneous movements of the other shocks in the model. As noted in Section III, this is numerically equivalent to the traditional Choleski decomposition when the external spread "moves first," that is, when the external spread shock occurs before other shocks.

[14]In all figures the dotted lines for the GIRs 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.

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 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 first finding is consistent with an extended version of the model presented in Section II to account for two levels of financial intermediation, along the lines of 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. The analysis shows that 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 business in Argentina for domestic lenders.

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.[15] The response of the real lending rate is positive but

[15]Note that on impact in Model A, the response of the cyclical component of output is a perverse blip that is reversed very quickly. As noted in Section III, generalized IRFs embody the full information of the contemporaneous correlation in the VAR innovations and consequently and contrary to standard IRF's, none of these correlation are set to zero. Recall that generalized IRFs are numerically equivalent to traditional Choleski IRFs when the variable of interest is place first in the ordering. In this context, the perverse output blip on "impact" essentially reflects the positive contemporaneous correlation between the VAR innovations in ES and those in $\ln(y/y_T)$ during the sample period (see Table A1) that is "picked up" by the generalized IRFs. Note, however, that this perverse blip is not observed with Model B.

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 in which 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 generalized 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 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.

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—although the effect of shocks to output itself become important after the second quarter of 1995.

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 0.7 percentage points, compared to about 1.3 for domestic spread shocks. Note that during the same period, the effect of external spread shocks are twice as large as those of output shocks, and almost four times as large as shocks to the real lending rate. The relatively limited impact of external spread shocks on the domestic lending rate 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 sizable effects on the behavior of output and domestic bank lending spreads in Argentina.[16]

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 the ex ante bank lending spread, movements in output (measured as deviations of output from trend and the growth rate of output), the real bank lending rate, and the external interest rate spread. Generalized variance decompositions, presented in Section IV, showed in particular, that at short horizons (less than 6 months) movements of domestic spreads are greatly influenced by their own historical shocks. At

[16]Note also that movements in output in this context can be consistent with a demand channel. Again, identifying more precisely the supply-side effects emphasized in this paper would require more disaggregated data.

longer forecast horizons, the external spread and the cyclical component of output played a greater role in explaining these movements. The effects of an external shock, modeled as a positive historical shock in external interest rate spreads, were analyzed in Section V using generalized impulse response functions. The results indicated 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 showed 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 generalized 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 these variables.

The experience of the 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 has adverse consequences in economies where banks and debt contracts are widely used to finance investment. Our results provide tentative support for the predictions of models based upon the notion of costly financial intermediation. Further research is needed to validate these results for other countries, and to identify their policy implications.

Appendix I
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:[17]

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

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

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

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

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

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

[17] 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).

In the particular case in which the aggregate 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

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

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

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

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

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

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

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

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

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, $d^2\Pi/dr^*dM_i < 0$. An Appendix (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 II
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:[18]

- 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 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 calculating standard lag-length tests, that is Akaike Information Criteria (AIC), Hannan-Quinn (HQ), and Schwarz. These

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

tests compare the cost of increasing the lag length (reduced degrees of freedom) to the benefit (increased information extraction from the data). Using a maximum lag length of six, all three tests suggested using six. This presents a problem due to the size of the sample: using the six lags means that each of the five equations would contain 31 (6*5+1) coefficients to estimate with 66 monthly observations (January 1993-June 1998). This translates into unacceptably low degrees of freedom and consequently low precision in the estimation. Rather using the 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 GIRs and GVDs 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.

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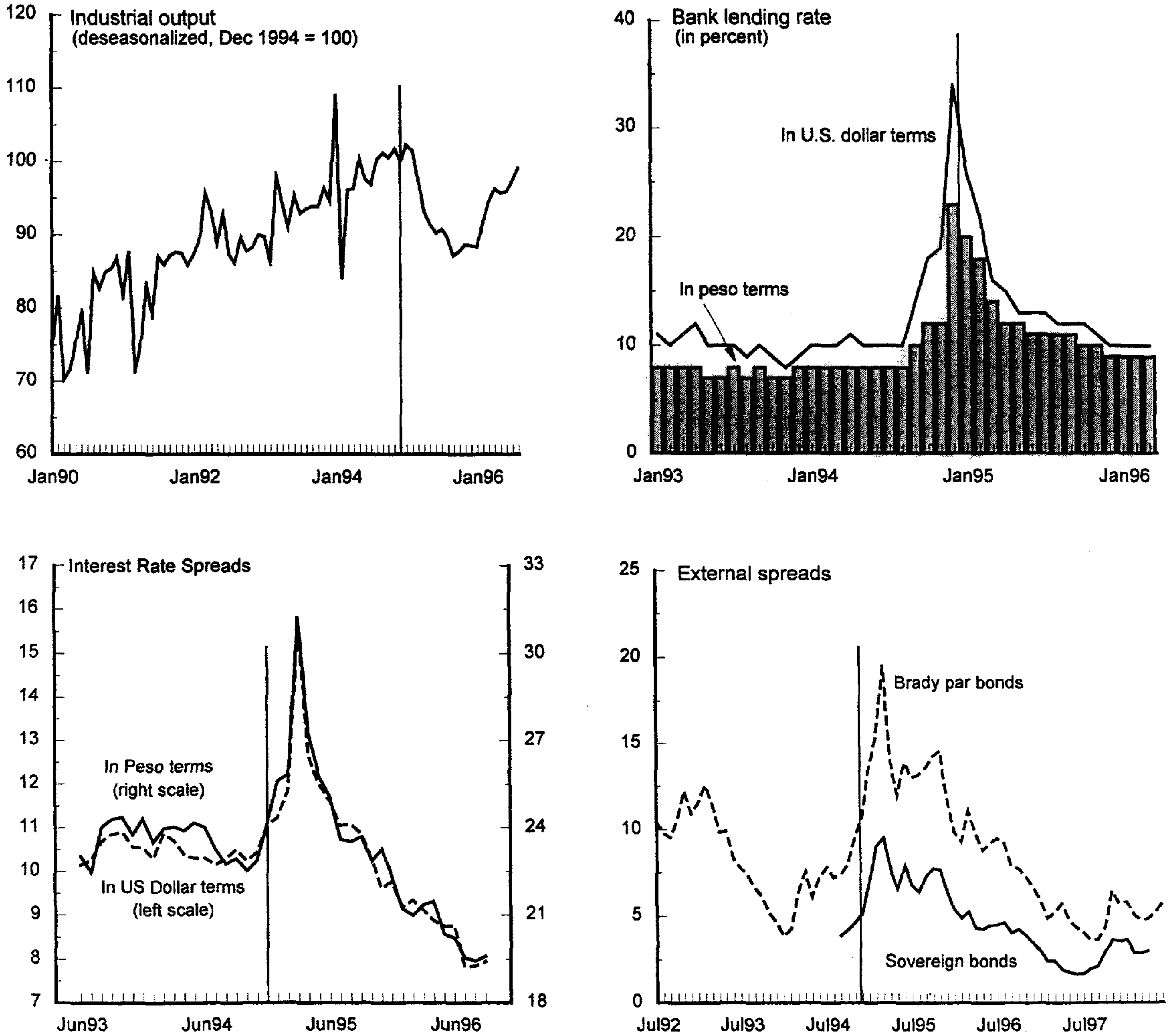
Figure 1
Argentina: Output and Interest Rates 1/

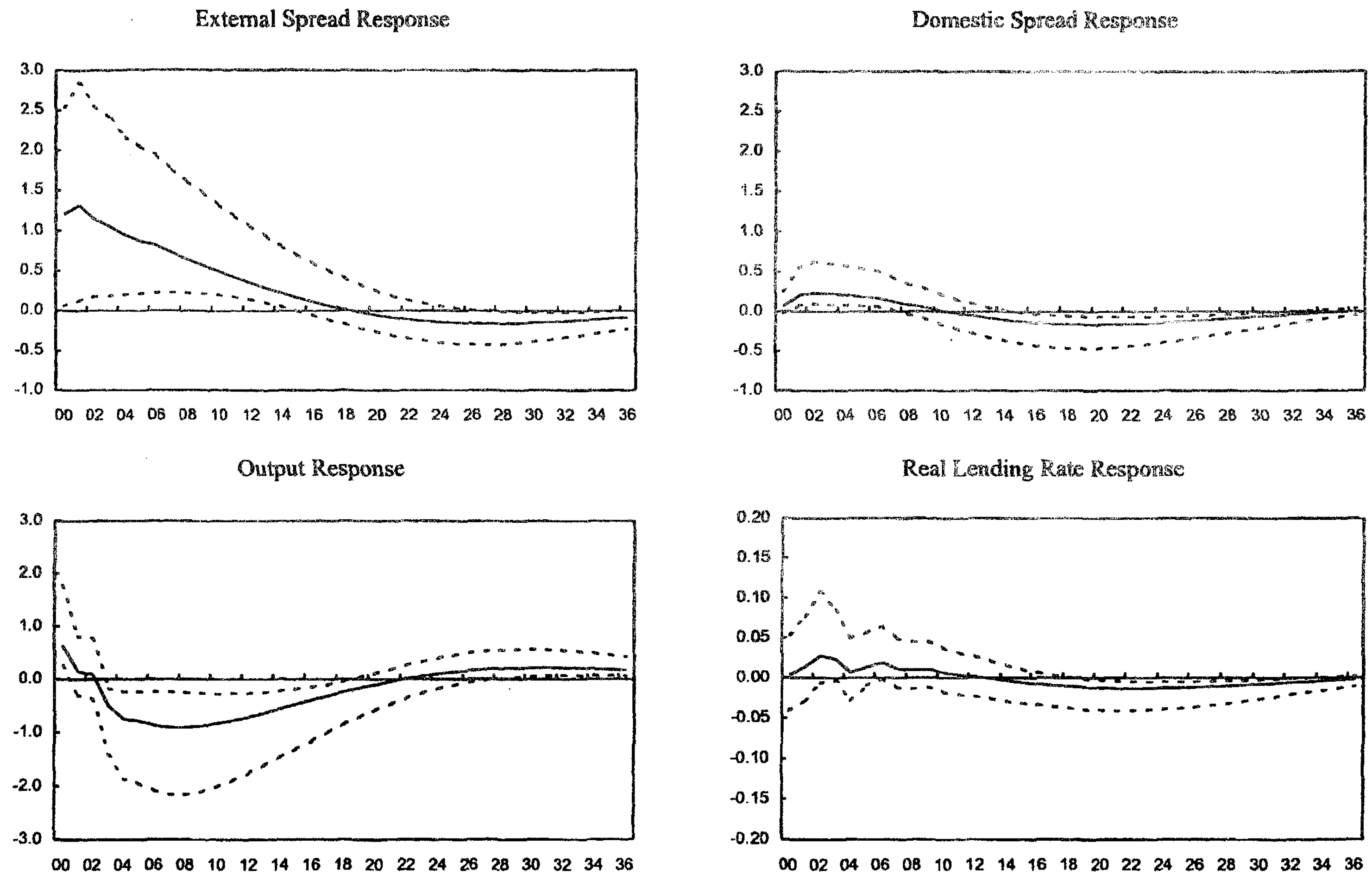


Sources: FIEL; International Monetary Fund, Bloomberg, Inc., and Merryll Lynch.

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

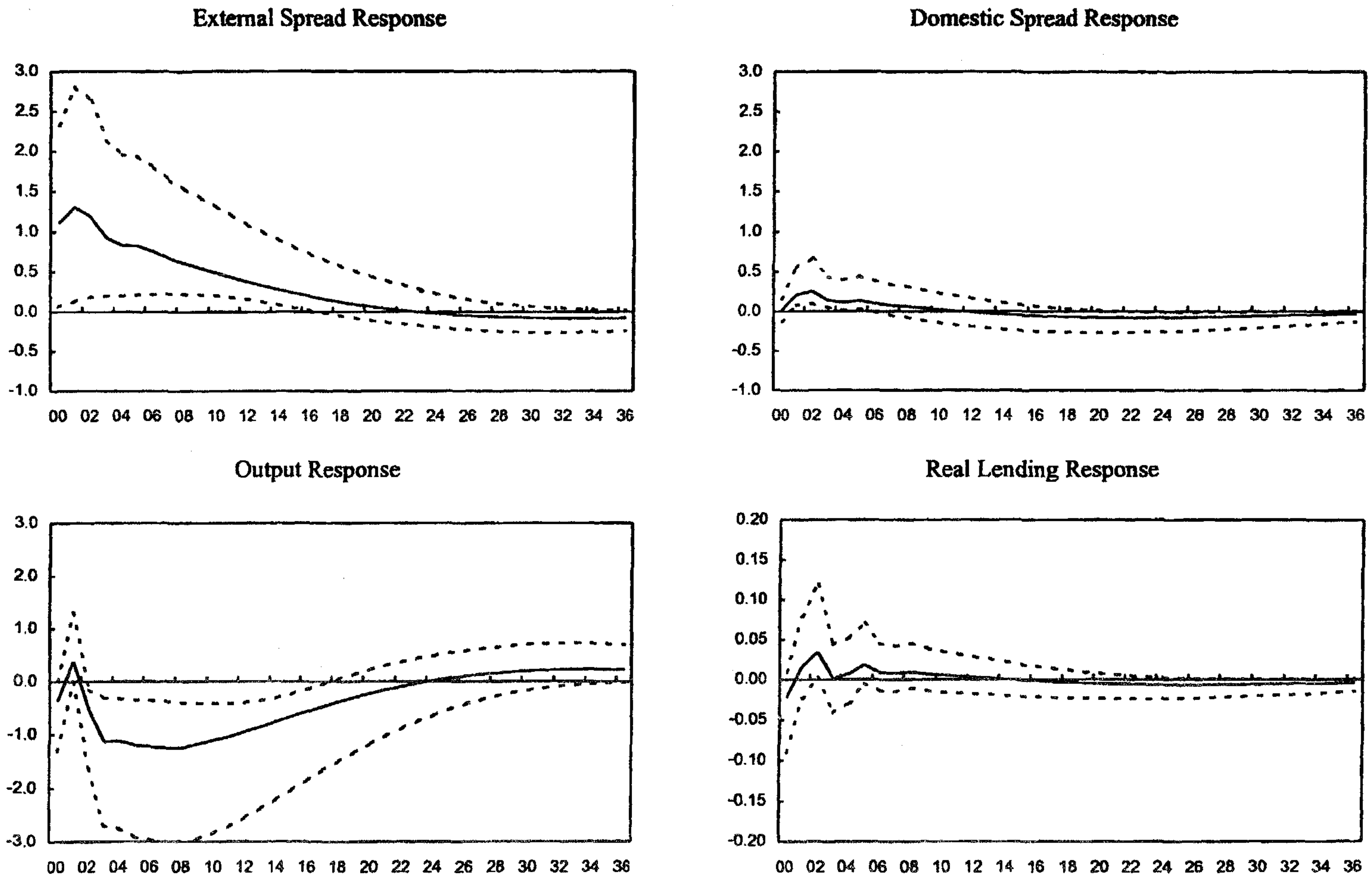
Figure 2. Generalized Impulse Responses, Model A.

(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

	Model A				Model B			
	External Spread (ES)				**External Spread (ES)**			
Months	Percentage of the variance associated with historical shocks to				Percentage of the variance associated with historical shocks to			
	ES	DS	$\ln(y/y_T)$	RL	ES	DS	$\ln(y_t/y_{t-12})$	RL
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
	Cyclical Component of Output ($\ln(y/y_T)$)				**Output Growth ($\ln(y_t/y_{t-12})$)**			
	Percentage of the variance associated with historical shocks to				Percentage of the variance associated with historical shocks to			
	ES	DS	$\ln(y/y_T)$	RL	ES	DS	$\ln(y_t/y_{t-12})$	RL
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
	Domestic Spread (DS)				**Domestic Spread (DS)**			
	Percentage of the variance associated with historical shocks to				Percentage of the variance associated with historical shocks to			
	ES	DS	$\ln(y/y_T)$	RL	ES	DS	$\ln(y_t/y_{t-12})$	RL
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
	Real Lending Rate (RL)				**Real Lending Rate (RL)**			
	Percentage of the variance associated with historical shocks to				Percentage of the variance associated with historical shocks to			
	ES	DS	$\ln(y/y_T)$	RL	ES	DS	$\ln(y_t/y_{t-12})$	RL
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 II for details.

Table 2. Generalized Historical Decompositions

Quarter	Model A						Model B					
	Cyclical Component of Output ($\ln(y/y_T)$)						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
	Domestic Spread (DS)						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.

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