# Fixed-Premium Deposit Insurance and International Credit Crunches

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*This article introduces a monopolistically competitive model of foreign lending in which both explicit and implicit fixed-premium deposit insurance increase the degree to which bank participation in relending to problem debtors falls below its globally optimal level. This provides a channel for fixed-premium deposit insurance to inhibit credit extension in bad states, resulting in an increase in the expected default percentage and an increase in the expected burden on the deposit insurance institution.*

While the perverse incentives faced by banks due to fixed-premium deposit insurance have been well-documented,[1] the literature has largely ignored the potential of deposit insurance to distort the organization of the banking industry. This paper introduces a simple model to fill this gap, in which fixed-premium deposit insurance plays a role in determining the structure of bank lending. To the extent that banking organization affects the ability of banks to act in concert, the paper introduces a new channel through which deposit insurance may have an adverse impact on lending outcomes.

The impact of deposit insurance on bank behavior has long been a source of concern to policymakers and researchers. A large literature exists which argues that fixed-premium deposit insurance increases the riskiness of bank lending portfolios (Kareken and Wallace 1978, Kareken 1986, Penati and Protopapadakis 1988, Jaffee 1989, Kane 1989, Duan, et al., 1992). In addition, Penati and Protopapadakis argue that "implicit deposit insurance," where regulators merge rather than close failing banks out of concern for the stability of the financial system, provides an additional subsidy. For example, from 1978 through 1984, only 20 percent of failed U.S. banks were closed. Moreover, these were largely small banks, representing only 0.2 percent of total deposits.

This paper demonstrates that the introduction of fixed-premium deposit insurance, both explicit and implicit, can magnify the degree to which credit extension is sub-optimal by increasing the number of banks participating in the lending package. The analysis is conducted through a monopolistically competitive two-period model of foreign lending, introduced in Section II.

The interesting decision in the two-period model comes at the end of the first period. Banks are confronted with the ability to increase the performance of their outstanding loans by rolling over debt at terms that would not appear to be profitable to unexposed creditors. However, there are positive spillovers associated with new lending, which implies that the disparity between the magnitude of new lending and that which is globally optimal will be increasing in the number of exposed banks. The incentive to avoid this

1. See Santomero (1984) for an extensive survey of this literature.

public good problem limits the number of participating banks in equilibrium. However, fixed-premium deposit insurance mitigates this incentive, increasing the number of participating banks and exacerbating the public good problem. This reduces the expected percentage of debt service, increases the probability of bank failure, and increases the expected burden on the deposit insurance institution. Simulation results below indicate that fixed-premium deposit insurance can have a relatively large impact on default probabilities.

While our qualitative results apply equally well to any situation where externalities among creditors may exist,[2] foreign lending provides a particularly clean example of lending externalities across banks. In domestic lending situations, creditors can partially deal with anticipated future renegotiation difficulties through debt covenants and other contract instruments[3] which are not binding in an international lending context. In addition, rescheduling negotiations in international lending take place under the auspices of the Paris Club, which applies the constraint of equal sharing rules, such that all loans have equal seniority. Finally, episodes of perceived sub-optimal credit extension are well documented in international lending, such as the failure of the Baker Plan to deal with the Latin American debt crisis (Cline 1989).

There is considerable evidence that the deposit insurance subsidy affected bank incentives in foreign lending. Event studies of the impact of the debt crisis in August 1982 on bank equity showed a consistently lower impact on excess returns than would be suggested by the magnitude of the news. For example, while uninsured bond spreads over LIBOR soared from 2 percent in August 1982 to over 7 percent in November (Edwards 1986), there was less than a 2 percent decline in the average annual excess returns of banks exposed to Mexico (Schoder and Vankudre 1986, Bruner and Simms 1987, and Spiegel 1992).[4] Similarly, James (1990) finds that changes in the value of bank stock equity are smaller than exposure-weighted movements in the secondary market prices of sovereign debt would imply.

This paper is organized as follows: Section I reviews the performance of commercial banks under the Baker Plan. Section II then introduces our theoretical model. The impact of deposit insurance, both explicit and implicit, is examined in Section III. Section IV contains simulations concerning the empirical predictions of the effects discussed in the theory. Section V concludes.

## I. THE BAKER PLAN: 1986–1988

The "Baker Plan," named for former Treasury Secretary James Baker, provides one of the best recent examples of collective action problems among international creditors. Subsequent to Mexico's suspension of payment on its external debt, a number of countries experienced difficulties in obtaining financing. The general belief concerning the difficulties faced by these countries was that their problems were ones of "illiquidity" rather than "insolvency." In other words, if financing could be obtained to get countries through a relatively difficult period, they would then be able to service all of their debts. Subsequent to this period "... countries could grow their way out of debt and could expand their exports enough to reduce their relative debt burdens to levels compatible with a return to normal credit market access" (Cline 1989 p. 177).

The Baker Plan called for commercial banks to extend approximately $7 billion annually, or 2.5 percent of total exposure, to fifteen highly indebted developing countries.[5] Cline (1989) claims that banks achieved capital flows of approximately $13 billion over the Baker Plan period, or about two-thirds of their $20 billion target.[6] It was well understood at the time that the anticipated disbursements from commercial banks under the plan were by no means certain. Brainard (1987) suggested that banks needed to understand how the involved government intended to manage the Baker Plan or "... increased official lending will merely substitute for reduced bank credits."

In retrospect, the magnitude of actual flows during the Baker Plan period seems even lower than Cline's estimate. Husain (1989) points out that while the IMF estimates that commercial banks committed $16.4 billion in new money and actually disbursed $15 billion—figures the commercial banks themselves used to support their claims of having come close to the Baker Plan targets—debtor country data show that net new long-term financing to the highly indebted countries amounted to only $4 billion (see Table 1). Moreover, if private nonguaranteed debt is taken into

2. See Bernanke (1991) for a discussion of domestic "credit crunches."

3. See, for example, Berlin and Mester (1992).

4. Beebe (1985) and Cornell and Shapiro (1986) show larger declines over the long run, but their measures do not approach the magnitudes observed in bond-spread movements.

5. The original countries were Algeria, China, Egypt, Greece, Hungary, India, Indonesia, Malaysia, Pakistan, Poland, Portugal, Thailand, and Turkey; Costa Rica and Jamaica were added later.

6. Ironically, public sector capital flows to the Baker nations actually did worse over the plan, falling by $4 billion annually. This occurred because of large decreases in IMF and bilateral lending.

TABLE 1

COMMERCIAL BANK LENDING TO HIGHLY INDEBTED COUNTRIES UNDER THE BAKER PLAN
(In billions of US dollars)

| | 1986 | 1987 | 1988 | 1986–1988 |
|---|---|---|---|---|
| CONCERTED NEW MONEY | | | | |
| Commitments | 8.3 | 2.4 | 5.6 | 16.4 |
| Disbursements | 3.2 | 5.7 | 6.0 | 15.0 |
| CHANGE IN EXCHANGE RATE | | | | |
| Adjusted claims | 3.5 | 0.6 | 2.0 | 6.1 |
| NETDISBURSEMENTS | –0.4 | 2.3 | 2.1 | 4.0 |

Source: Husain (1989)

account, there were net repayments to commercial banks amounting to $2.4 billion. In no case did commercial banks provide more net financing than they received in interest payments. Husain also adds that "U.S. banks have been most active in reducing their developing country exposure. Between mid-1987 and the end of the third quarter of 1988, these banks reduced their claims on all developing countries by more than $20 billion. More than half of this represented a reduction in claims on highly indebted countries" (p. 14).

In summary, the Baker Plan is an example of collective action problems across commercial banks. It was generally perceived that increasing exposure in the aggregate was desirable from the point of view of the exposed banks, but individually each bank had the incentive to "free-ride" on the efforts of the other creditors by not fulfilling its disbursement commitment. The result was that the level of new money extended to the highly indebted nations was sub-optimal from the aggregate creditor perspective. Because commercial banks faced incentive problems in taking collective action, a resolution of the debt crisis required turning to voluntary methods, such as the "market-based menu approach" associated with the Brady Plan (see Diwan and Spiegel 1994).

## II. A MONOPOLISTICALLY COMPETITIVE MODEL OF BANK LENDING

### *Setup*

In this section, we derive a formal model which exhibits collective action problems similar to those encountered under the Baker Plan. There are three types of players in the model, a debtor nation government, a group of identical monopolistically competitive banks, and atomistic depositors. The extensive form of the model has five stages: In the first stage, the magnitude of first-period lending by individual banks, $l_1$, and the number of banks participating in the initial lending package, $n$, are determined. Total first period lending, $L_1$, then equals $nl_1$. Loans are assumed to be short term, coming due at the end of the first period.[7] For simplicity, all first-period lending is assumed to be consumed. First-period output of the debtor nation, $q_1$, is determined in the second stage. In the third stage, banks choose an amount of new lending $l_2$, so that total new lending $L_2 = nl_2$. In the fourth stage, the debtor nation government chooses its percentage of debt service on outstanding first-period loans, $\pi_1$, where $\pi_1 \in [0,1]$. Finally, second period debtor nation output, $q_2$, is determined in the final stage, which simultaneously determines the percentage of debt service on outstanding second period lending, $\pi_2$, where $\pi_2 \in [0,1]$.

$q_1$ and $q_2$ are assumed to be exogenous independently distributed random variables distributed uniformly on the interval $[0,1]$.[8] Total lending in each period is equal to the number of banks participating in the lending package, $n$, times individual bank lending, $l_t$. As we show below, $n$ will be constant across periods. Let $\bar{r}_t$ represent one plus the contractual nominal rate of interest on the loan in period $t$ ($t = 1, 2$). For simplicity, $\bar{r}_t$ is taken as exogenous.[9] Consequently, the outstanding obligation on period $t$ loans at the end of period $t$ will be equal to $\bar{r}_t nl_t$.

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7. Creditors might respond to future anticipated renegotiation problems by lengthening the maturity of their debt contract (Sharpe 1991). In practice, however, creditors responded to the increase in the perceived riskiness of the highly indebted countries by shortening the maturity of their loans. The lack of long-term lending may stem from equal-sharing, since some of the benefits of extending a long-term loan would spill over to short-term creditors.

8. We assume that $q_1$ and $q_2$ are independent for simplicity. This does not drive the results below, but it does increase the parameter space in which relending to a problem debtor is an optimal response. If the two were positively correlated, it would be more "likely" that relending would be throwing good money after bad. Nevertheless, since the point of the exercise is to introduce an example where collective action problems may arise, this assumption is relatively innocuous.

The assumption of exogeneity of these output variables is also made for simplicity and drives none of the results below. However, allowing for "debt overhang" effects, where indebtedness may affect output levels, may also affect the desirability of relending in a more general model.

9. It is well known that allowing banks to choose both $l_t$ and $\bar{r}_t$ would result in a multiplicity of equilibria. Holding $\bar{r}_t$ constant is valid if the debtor is credit constrained, which we assume for a problem international debtor.

*Debtor Decisions*

We assume that output is not storable and consumption takes place at the end of each period. First period consumption, $c_1$, will equal output plus new lending minus service on outstanding debt:

(1a) $$c_1 = q_1 - \pi_1\bar{r}_1 L_1 + L_2 .$$

Similarly, second period consumption is equal to second period output minus service on outstanding second period obligations:

(1b) $$c_2 = q_2 - \pi_2\bar{r}_2 L_2 .$$

The debtor chooses $\pi_t$ $(t = 1, 2)$ to maximize utility. Debtor utility, $u_t$, is an increasing function of current consumption and a decreasing function of an exogenous "default penalty," $P(\pi_t)$:

(2) $$U_t = U[c_t, P(\pi_t)]; \quad (t = 1, 2).$$

where $U_c>0$, $U_{cc}<0$, $U_{ccc} = 0$, $U_P<0$, $U_{PP}<0$, $P_\pi<0$, $P_{\pi\pi}<0$, and $U_{cP} = 0$. $P(\pi_t)$ is decreasing in the percentage of debt service and is intended to represent the discounted costs of default.[10] In the Appendix, we show that maximization of (2) subject to (1a) and (1b) implies that the debtor's first-period decision satisfies the triple:

(3a) $$\pi_1 = \pi_1^* (q_1, L_1, L_2),$$

where $\partial\pi_1/\partial q_1>0$, $\partial\pi_1/\partial L_1<0$, $\partial\pi_1/\partial L_2>0$, and the debtor's second-period decision satisfies:

(3b) $$\pi_2 = \pi_2^* (q_2, L_2),$$

where $\partial\pi_2/\partial q_2>0$, $\partial\pi_2/\partial L_2<0$. Note that the expected percentage of debt service on outstanding first-period loans is increasing in second-period lending. This raises the possibility of profitable rescheduling by exposed creditors, as we show below.

*Deposit Rates*

Define $r_t$ $(t = 1, 2)$ as one plus the risk-free rate of interest. Depositors are risk neutral and atomistic and have the right to remove deposits after each period. They therefore require an expected return equal to $r_t$. Define $q_t^B$ $(t = 1, 2)$ as the probability of bankruptcy by the representative creditor in period $t$. Note that since $q_t$ is distributed uniform on the unit interval, $q_t^B$ also represents the minimum realization of $q_t$ which leaves the creditor solvent. Uninsured depositors will therefore require a nominal rate of interest on uninsured deposits equal to $r_t/(1 - q_t^B)$ in period $t$. Define $\tau$, $\tau \in [0,1]$, as the share of bank deposits that are insured by the deposit insurance institution of the commercial bank, taken as exogenous.[11] Finally, define $\gamma_t$ as one plus the average rate of interest paid by the representative commercial bank on deposits. $\gamma_t$ will satisfy:

(4) $$\gamma_t = r_t[(1 - \tau)/(1 - q_t^B) + \tau]; \quad (t = 1, 2).$$

Note that $\gamma_t$ is increasing in $q_t^B$ and decreasing in $\tau$. To the extent that deposits are uninsured, the average rate paid to depositors will be increasing in the the probability of bankruptcy, $q_t^B$. However, to the extent that deposits are insulated from loss through deposit-insurance, the sensitivity of $\gamma_t$ to $q_t^B$ is diminished.

*Creditors*

There are assumed to be a large number of homogeneous potential creditors who have identical portfolios of non-debtor-nation loans with face values of $a_t$ that pay nominal interest equal to $\rho_t$. Both $a_t$ and $\rho_t$ are assumed to be invariant with respect to the lending decisions towards the debtor nation and deterministic. Given $l_t$ and $\gamma_t$, the creditor finances its lending by issuing $(a_t + l_t)$ in liabilities. The creditor's return in period $t$ satisfies:

(5) $$R_t = \rho_t a_t + \pi_t \bar{r}_t l_t - \gamma_t(a_t + l_t); \quad (t = 1, 2).$$

We assume that banks have limited liability and that regulators close banks in the event that a bank shows negative net worth in either period. We examine the case where regulators merge failing banks under some circumstances below. For simplicity, we assume banks do not retain earnings. Consequently, a bank fails if $R_t$ falls below zero in either period.

*Second Period Lending Decision*

To insure sub-game perfection, we begin with the second period creditor decision. In the second period, creditors choose their individual amount of new lending, $\hat{l}_2$, taking other creditors' new lending as given. Given the po-

10. The default penalty is needed to generate positive lending in equilibrium, but its specification does not drive our results. Default penalties in sovereign lending have been motivated by loss of future access to capital markets (Eaton and Gersovitz 1981), seizure of assets (Bulow and Rogoff 1989), or loss of reputation (Grossman and Van Huyk 1988). Lindert and Morton (1989) show that the ex-post rate of return on sovereign lending has historically been competitive, implying that the perception of a penalty for default exists.

11. Generalizing the model by allowing depositors to increase $\tau$ by "brokering" deposits across a number of banks would actually strengthen the results below by providing an additional incentive for an increase in the number of banks.

tential for sub-optimal levels of credit extension we derive below, only one lender would emerge per nation in the absence of incentives for banks to avoid taking on the entire lending package. However, lending to debtor nations is usually broken up over a large number of banks. To accommodate this empirical fact, we assume that banks are risk-averse and that the riskiness of their asset portfolio is increasing in bank exposure to the debtor nation. Under this assumption, we specify the value function of creditors in period 2, $\Omega_2$, as increasing in the returns on bank operations and decreasing in bank risk, $\sigma_2$, where $\sigma_t$ is an increasing function of $\hat{l}_t$, $\sigma_t = \sigma(l_t)$ $(t = 1, 2)$. In period 2, $\Omega_2$ satisfies:

$$\Omega_2 = \Omega_2[R,\sigma(l_2)]\ , \tag{6}$$

where $R = R_1 + (1/r_2)R_2$ and $\Omega_R>0$, $\Omega_\sigma<0$, $\Omega_{R\sigma} = 0$, $\sigma_{l_2}>0$.

Consider a representative exposed creditor who extended a loan equal to $l_1$ in the first period. Subsequent to the realization of $q_1$, the creditor's decision problem is to choose the value of $\hat{l}_2$ which maximizes the expected value of $\Omega_2$. The creditor's first-order condition satisfies:[12]

$$\frac{\partial\Omega_2}{\partial R}\left[\frac{\partial\pi_1}{\partial L_2}\bar{r}_1 l_1 + \frac{1}{r_2}\left[\pi_2\bar{r}_2 - \gamma_2 + \frac{\partial\pi_2}{\partial L_2}\bar{r}_2\hat{l}_2 - \frac{\partial\gamma_2}{\partial\hat{l}_2}(a_2+\hat{l}_2)\right]\right] + \frac{\partial\Omega_2}{\partial\sigma}\frac{\partial\sigma}{\partial\hat{l}_2} = 0. \tag{7}$$

The first term in equation (7) reflects the sum of pecuniary benefits in both first- and second-period earnings of an increase in $l_2$ on the margin. The first portion of that term reflects the positive impact on outstanding loans, while the second term reflects the impact on second-period loans. The overall sign of this term is ambiguous because of the second term in equation (7). If banks are sufficiently risk-averse, they will cease lending to the debtor at a point which leaves it profitable for some new entry to occur in the second period. Since this would greatly complicate the model, we rule this out. This requires the parameter restriction that the positive impact on the margin to an exposed creditor exceeds the adverse impact of the increase in risk, i.e.,

$$\frac{\partial\Omega_2}{\partial R}\frac{\partial\pi_1}{\partial L_2}\bar{r}_1 l_1 + \frac{\partial\Omega_2}{\partial\sigma}\frac{\partial\sigma}{\partial\hat{l}_2} > 0.$$

Satisfaction of this restriction leaves the rate of return on second-period lending less than zero and the number of exposed creditors in the second period unchanged at $n$.

The creditor also considers the impact on its deposit rate, $\gamma_2$, when making its new lending decision. We demonstrate in the Appendix that the second-period deposit rate, $\gamma_2$, is a triple:

$$\gamma_2 = \gamma_2(\hat{l}_2,L_2,\tau), \tag{8}$$

where $\partial\gamma_2/\partial\hat{l}_2>0$, $\partial\gamma_2/\partial L_2>0$, and $\partial\gamma_2/\partial\tau<0$. We also conduct the comparative static exercises for (7) and demonstrate that the individual bank's second-period lending is a quadruple:

$$\hat{l}_2 = \hat{l}_2(\hat{l}_1,L_1,L_2,\tau), \tag{9}$$

where $\hat{l}_t$ $(t = 1, 2)$ represents an individual bank's choice of $l_2$ taking the decisions of other banks as given and $\partial\hat{l}_2/\partial l_1>0$, $\partial\hat{l}_2/\partial L_1>0$, $\partial\hat{l}_2/\partial L_2<0$, and $\partial\hat{l}_2/\partial\tau>0$.

Equation (9) is derived as the optimum lending choice for an individual bank taking the lending choices of other banks as given. However, in equilibrium $L_2 = nl_2$. Substituting for $L_2$ in (9), and recalling that $\partial l_2/\partial L_2<0$, we obtain the quadruple:

$$\hat{l}_2 = \hat{l}_2(\hat{l}_1,L_1,n,\tau). \tag{10}$$

where $\partial\hat{l}_2/\partial n<0$.

Second-period lending is increasing in the magnitude of first-period lending because first-period debt service is increasing in second-period lending. Consequently, banks have an incentive to engage in "conciliatory relending" by rolling over a portion of the outstanding debt to decrease the magnitude of first-period default. The greater is a bank's exposure, the greater are the benefits from a unit increase in $\pi_1$ and the greater is the magnitude of relending per bank. An individual bank's second-period lending is decreasing in the total magnitude of outstanding first-period lending, however, because the degree to which a bank benefits from its own relending efforts is decreased by the amount of outstanding claims on the debtor. In equation (10), second period relending is decreasing in $n$, the total number of creditors, because the positive first-period effects of relending diminish as the magnitude of new lending increases, while the negative effects on expected second-period returns are enhanced. Finally, the magnitude of second period lending per bank is increasing in the share of insured deposits. Deposit insurance makes it less costly to engage in conciliatory relending. Consequently, holding all else equal, a bank would be more willing to engage in conciliatory relending the larger is the share of insured deposits.

12. Equation (7) is simplified by noting that on the margin $\partial\pi_t/\partial\hat{l}_2 = \partial\pi_t/\partial L_2$ $(t = 1, 2)$.

### *Comparison with a Globally Optimal Solution*

Finally, we compare the creditor's solution with one which would be globally optimal across creditors. The solution that would maximize global creditor profits would satisfy:

$$(11)\quad \frac{\partial\Omega_2}{\partial R}\left[\frac{\partial\pi_1}{\partial L_2}\bar{r}_1 L_1 + \frac{1}{r_2}\left[\pi_2\bar{r}_2-\gamma_2+\frac{\partial\pi_2}{\partial L_2} - \frac{\partial\gamma_2}{\partial\hat{l}_2}(na_2+L_2)\right]\right] + \frac{\partial\Omega_2}{\partial\sigma}\frac{\partial\sigma}{\partial\hat{l}_2} = 0\,.$$

Rearranging terms, we obtain:

$$(12)\quad \frac{\partial\Omega_2}{\partial R}\left[n\left[\left[\frac{\partial\pi_1}{\partial L_2}\bar{r}_1 L_1+\frac{\partial\pi_2}{\partial L_2}\bar{r}_2 l_2/r_2\right] - \left[\frac{\partial\gamma_2}{\partial\hat{l}_2}(a_2+l_2)\right]/r_2\right] + \left[\pi_2\bar{r}_2-\gamma_2\right]/r_2\right] + \frac{\partial\Omega_2}{\partial\sigma}\frac{\partial\sigma}{\partial\hat{l}_2} = 0.$$

The first bracketed term can be signed as positive by (7), since the other terms in (7) are negative.

Comparing equations (12) and (7), we can see the source of the sub-optimality of lending by the individual bank from the global point of view of creditors. The individual bank's first-order condition accounts for the positive impact of new lending on the rate of return on his outstanding loans. However, he does not consider the positive spillovers to other outstanding creditors. To see this clearly, note that the positive impact on an individual bank's profits from new lending in equation (12) is multiplied by $n$, the number of banks in the lending package, while this number is only multiplied by one in the individual bank's decision in equation (7). Consequently, the individual bank's level of new lending is sub-optimal and the disparity between the individual solution and the globally optimal solution among creditors as a group is increasing in $n$.

However, we should point out that the socially optimal outcome would be the one that would emerge without fixed-premium deposit insurance, rather than one which induced the globally optimal level of second-period lending. The optimal allocation will include some degree of sub-optimal second-period lending because of bank risk-aversion. This outcome could alternatively be achieved by charging banks a variable premium equal to the expected liability of the deposit insurance institution. Nevertheless, while such a policy may be optimal ex ante, it may not be time-consistent ex post since regulators will wish to enhance second-period credit extension.

## III. EQUILIBRIUM UNDER DEPOSIT INSURANCE

### *Equilibrium under Explicit Deposit Insurance*

Given the expected second-period responses derived above, we now compute the first-period values of $\hat{l}_1$ and $n$. We proceed under the assumption that a collective action problem exists among creditors, i.e., that the percentage of first-period debt service is decreasing in $n$. The conditions for this innocuous assumption are derived in the Appendix. Under this assumption, we demonstrate in the Appendix that the first-period deposit rate, $\gamma_1$, is a quadruple:

$$(13)\quad \gamma_1 = \gamma_1(\hat{l}_1, l_1, n, \tau),$$

where $\partial\gamma_1/\partial\hat{l}_1>0$, $\partial\gamma_1/\partial l_1>0$, $\partial\gamma_1/\partial n>0$, and $\partial\gamma_1/\partial\tau<0$.

The bank's cost of funds is increasing in $l_1$ (and $\hat{l}_1$) and $n$ because increases in both raise the stock of outstanding debt and lower the expected percentage of debt service. In addition, increases in $n$ exacerbate the public good problem associated with new lending. However, deposit insurance, by insulating a portion $\tau$ from bankruptcy risk, reduces the sensitivity of depositor interest rates to the probability of creditor bankruptcy. Hence $\partial^2\gamma_1/\partial n\partial\tau<0$, as we show in the appendix.

In the first period, participating creditors choose the value of $\hat{l}_1$ which maximize expected returns subject to limiting their risk exposure, taking the actions of other creditors as given. Similar to our assumption above, we specify the value function of creditors in period 1, $\Omega_1$, as increasing in the returns on bank operations and decreasing in bank risk on both periods, $\sigma_1$ and $\sigma_2$, where $\sigma_t$ ($t$ = 1, 2) is an increasing function of $\hat{l}_t$:

$$(14)\quad \Omega_1 = \Omega_1[R, \sigma(\hat{l}_1), \sigma(\hat{l}_2)]\,.$$

The participating creditor's first-order condition satisfies:

$$(15) \qquad \frac{\partial \Omega_1}{\partial R}\left[\frac{\partial \pi_1}{\partial L_1}\bar{r}_1\hat{l}_1 + \pi_1\bar{r}_1 - \gamma_1 - \frac{\partial \gamma_1}{\partial \hat{l}_1}(a_1+\hat{l}_1) + \frac{1}{r_2}\left[\pi_2\bar{r}_2 - \gamma_2 + \frac{\partial \pi_2}{\partial L_2}\bar{r}_2\hat{l}_2 - \frac{\partial \gamma_2}{\partial \hat{l}_2}(a_2+\hat{l}_2)\right]\frac{\partial \hat{l}_2}{\partial \hat{l}_1}\right] + \frac{\partial \Omega_1}{\partial \sigma}\frac{\partial \sigma}{\partial \hat{l}_1} + \frac{\partial \Omega_2}{\partial \sigma}\frac{\partial \sigma}{\partial \hat{l}_2}\frac{\partial \hat{l}_2}{\partial \hat{l}_1} = 0\,.$$

In addition to the maximization decisions of the individual banks, monopolistic competition among banks with free entry will lead to zero expected "value function" profits for participating banks in each period. Let $\Omega^n$ equal the expected value function for the representative bank if it chooses not to enter. Competition across banks insures that banks will continue to enter until:

$$(16) \qquad \Omega^n \geq \Omega_1.$$

We therefore define a competitive equilibrium satisfying the assumptions above by the solutions $\hat{l}_1^*$ and $n^*$ which represent the maximum level of $n\hat{l}_1$ which satisfies equations (15) and (16) with equality. See Figure 1. Despite the fact that both curves are downward-sloping, the initial equilibrium is the standard one in monopolistic competition models, with one equation representing the individual lenders' profit maximization decision, the "MM curve," and one equation representing a zero-profit condition, the "ZZ curve." Note that both curves are functions of the share of insured deposits $\tau$.

To demonstrate the impact of deposit insurance, we conduct a comparative static exercise on the parameter $\tau$. One can consider the introduction of deposit insurance as a discrete increase in $\tau$ from a zero level, which our analysis approximates. The comparative statics of the model satisfy:

$$(17) \qquad \begin{bmatrix} \partial\Omega_1/\partial\hat{l}_1^2 & \partial^2\Omega_1/\partial\hat{l}_1\partial n \\ \partial\Omega_1/\partial\hat{l}_1 & \partial\Omega_1/\partial n \end{bmatrix}\begin{bmatrix} \partial\hat{l}_1/\partial\tau \\ \partial n/\partial\tau \end{bmatrix} = \begin{bmatrix} -\partial^2\Omega_1/\partial\hat{l}_1\partial\tau \\ -\partial\,\Omega_1/\partial\tau \end{bmatrix}.$$

We sign the terms in the Appendix. As we suggested above, the deposit insurance subsidy affects the equilibrium through two channels: First, the subsidy increases total lending; second, banks have less incentive to organize the lending package in a form conducive to collective action. See Figure 2. An increase in $\tau$ shifts out both curves. The

FIGURE 1

INITIAL EQUILIBRIUM

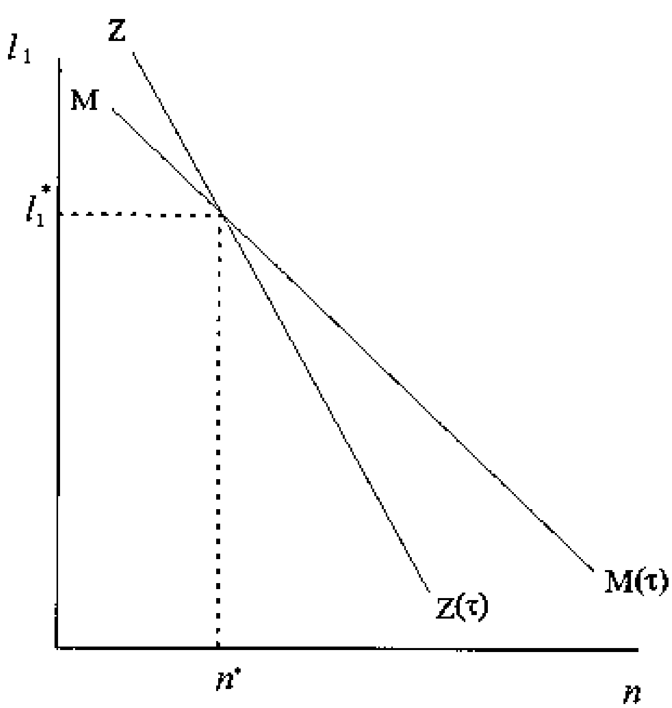
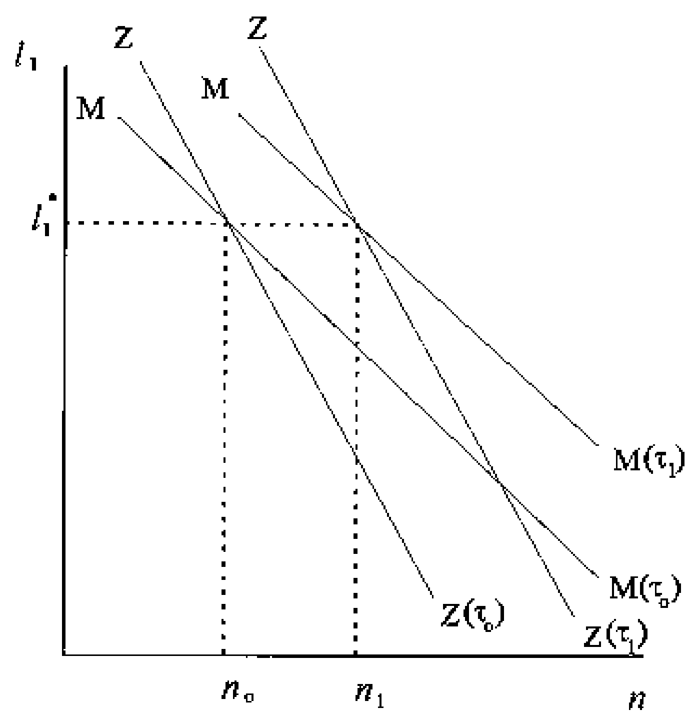



FIGURE 2

COMPARATIVE STATIC RESULTS



zero-profit condition (ZZ) curve shifts out because, holding all else equal, an increase in the share of deposit insurance increases the returns from lending. Given any value of $l_1$, this implies an increase in $n$ to return to zero profits. The MM curve shifts out because holding $n$ constant, the reduction in deposit rates leads each bank to make a larger initial loan, resulting in an increase in $l_1$. However, our comparative static exercise indicates that in equilibrium $\partial n/\partial\tau > 0$ and that $\partial l_1/\partial\tau = 0$. In other words, all of the increase in lending stems from entry rather than increases in individual bank lending. Note that this result magnifies the degree to which second-period lending falls below the global optimum.

### Liability of the Deposit Insurance Institution

Define the expected liability of the deposit insurance institution from exposure to the debtor nation in period $t$ as $\psi_t$. $\psi_t$ satisfies:

$$(18) \qquad \psi_t = \tau q_t^B n\hat{l}_t \quad (t = 1, 2).$$

Differentiating $\psi_1$ with respect to $\tau$ yields:

$$(19) \qquad \partial\psi_1/\partial\tau = q_1^B n\hat{l}_1 + \tau q_1^B \hat{l}_1 (\partial n/\partial\tau) + \tau n\hat{l}_1 (\partial q_1^B/\partial n)(\partial n/\partial\tau) > 0,$$

where:

$$(\partial q_1^B/\partial n)(\partial n/\partial\tau) = -[(\partial\pi_1/\partial n)(\partial n/\partial\tau)\bar{r}_1\hat{l}_1]/(\partial\pi_1/\partial q_1)\bar{r}_1\hat{l}_1 > 0.$$

Equation (19) shows that an increase in $\tau$ unambiguously increases the expected liability of the deposit insurance institution. The first term captures the direct effect: Given the exposure of banks and the expected probability of bankruptcy, an increase in $\tau$ will increase the expected liability of the deposit insurance institution. However, the other two terms are also positive. The second term shows that fixed-premium deposit insurance gives banks an incentive to increase their lending, all of which comes from an increase in $n$. The third term reflects the impact of deposit insurance on the probability of bankruptcy, $q_1^B$.[13] This term is enhanced in our model by the public good problem associated with relending.

### "Implicit Deposit Insurance"

Finally, we examine the implications of extending deposit insurance to insure "implicitly" some uninsured bank deposits. In their discussion of implicit deposit insurance, Penati and Protopapadakis (1988) claim that regulators distinguish between two types of loans: "local loans," whose failure only harms exposed banks, and "system-threatening" loans, whose failure would threaten the stability of the banking system and the solvency of the deposit insurance institution. They respond to a local loan default by closing failing banks, while they respond to systemic loan defaults by merging failing banks with other banks. The salient distinction is that uninsured deposits are carried at par subsequent to a merger, while uninsured deposits in closed banks lose their value.

Assessing the impact of implicit deposit insurance on banking organization requires specification of the criterion used by the bank regulators in identifying "system-threatening loans." Penati and Protopapadakis (1988) took this designation as exogenous. Here, we endogenize the criterion and show that the equilibrium can be affected by the criterion used by regulators to identify system-threatening loans.

Define $p_s$ as the probability that uninsured deposits will be carried at par subsequent to a bank failure, where $0 \le p_s \le 1$. For simplicity, we assume that the risk associated with uncertainty concerning the policy rule is diversifiable, so that it does not affect the value of $\sigma$. Suppose that $p_s$ is an increasing function of total exposure to the debtor, $p_s = p_s(nl_1)$. Define $E(\tau)$ as the expected total share of bank deposits subject to deposit insurance, either explicit or implicit. $E(\tau)$ satisfies:

$$(20) \qquad E(\tau) = \tau + p_s(nl_1)(1-\tau).$$

Equation (20) identifies the link between the equilibrium lending decision and the probability of implicit insurance. Implicit deposit insurance gives banks an incentive to tailor the lending package in a way that enhances the probability that the deposit insurance institution will merge rather than close a failing bank.[14]

To examine the impact of an increase in the importance of implicit deposit insurance under this criterion, we assume that $p_s$ is linear in the magnitude of first-period lending: $p_s = \delta(nl_1)$.[15] We then can examine the implications of an increase in $\delta$ as an example of an increase in the sensi-

13. Note that the impact of changes in $\gamma_1$ on the probability of bankruptcy does not affect $\psi_1$ because they already reflect a liability of the deposit insurance institution.

14. Penati and Protopapadakis (1988) suggested alternatively that the probability that loans are implicitly insured may be increasing in the number of banks involved in the lending package, so that $p_s = p_s(n)$. The qualitative results under this alternative criterion would be identical.

15. Since $0 \le p_s \le 1$, this linear specification must be considered as a local approximation to a non-linear function.

tivity of the probability of a bail-out by the deposit insurance institution to $nl_1$. The comparative statics of the model satisfy:

$$(21)\qquad \begin{bmatrix} \partial\Omega_1/\partial\hat{l}_1^2 & \partial^2\Omega_1/\partial\hat{l}_1\partial n \\ \partial\Omega_1/\partial\hat{l}_1 & \partial\Omega_1/\partial n \end{bmatrix} \begin{bmatrix} \partial\hat{l}_1/\partial\delta \\ \partial n/\partial\delta \end{bmatrix} = \begin{bmatrix} -\partial^2\Omega_1/\partial\hat{l}_1\partial\delta \\ -\partial\Omega_1/\partial\delta \end{bmatrix},$$

where the first matrix has the same signs as above.

We show in the Appendix that the comparative static solutions are $\partial n/\partial\delta>0$ and $\partial\hat{l}_1/\partial\delta = 0$. $n$ is increasing in $\delta$ for two reasons. First, an increase in $\delta$, holding the number of banks in the system constant, represents an increase in the expected share of deposits covered by the deposit insurance institution. Consequently, this directly reduces bank costs and induces additional lending through increases in $n$. In addition, increasing national exposure through an increase in $n$ increases the probability that regulators merge rather than close failing banks, reducing deposit rates. In other words, implicit deposit insurance rewards banks for organizing themselves in a system-threatening manner by increasing the probability of a deposit insurance institution bail-out.

## IV. SIMULATIONS

To examine the potential importance of both explicit and implicit deposit insurance, we use numerical simulations. This requires the assumption of specific functional forms. To make the simulations realistic, we choose parameter values which would be profitable for the banks ex-ante. However, to allow for an analytic solution, we linearize the relationships between the level of lending and the expected percentage of debt service:

$$\pi_1 = 1 - (0.002 \cdot nl_1)$$

and the impact of first-period loans on the second-period returns:

$$E(\pi_2 l_2 - \gamma_2) = -\,0.05[1 - (0.001 \cdot nl_1).$$

The magnitudes of these specifications were chosen to insure an interior solution for the probability of default between 0 and 1. Moreover, we assume that the creditor has a mean-variance value function and that the variance of profits is linear in exposure to the debtor, with $\phi$ representing creditor sensitivity to exposure:

$$\sigma(l_1) = \phi l_1.$$

We assume that the expected probability of bankruptcy is equal to one minus the expected level of debt service. This simplifies $\gamma_t$:

$$\gamma_t = r_f + [(1-\tau)(1-\pi_t)]; \quad (t = 1, 2).$$

The share of explicit deposit insurance is assumed to be roughly equal to $\tau = 0.65$.[16] The specifications of the other exogenous parameters are: $r_2 = 1.10$; $\overline{r}_1 = \overline{r}_2 = 1.20$. Under the "implicit deposit insurance" regime, we assume that the expected percentage of insured deposits is equal to:[17]

$$\tau = 0.65 + (0.01 \cdot nl_1).$$

Given these specifications, simulations were run for a variety of possible values of $\phi$ under four alternative regimes: (1) no deposit insurance, (2) explicit deposit insurance, (3) explicit and implicit deposit insurance, and (4) 100 percent deposit insurance. The results are reported in Table 2 for various values of $\phi$. The introduction of deposit insurance results in an increase in the number of banks in the system, a decrease in the expected percentage of debt service, and an increase in the expected burden on the deposit insurance institution as a percentage of outstanding loans.

Our results imply that the introduction of explicit deposit insurance brings an expected loss to the deposit insurance institution of 2.1 percent of outstanding loans. Moving to 100 percent deposit insurance almost doubles the expected burden to 4 percent of outstanding loans. Note that these expected liabilities were obtained under parameter values for which lending to the debtor nation is profitable ex ante for creditors.[18]

## V. CONCLUSION

In this paper, we examined the implications of fixed-premium deposit insurance in a foreign lending model where rescheduling exhibits positive spillovers across creditors. Our results show that deposit insurance raises the number of banks participating in the lending package through three channels: First, deposit insurance acts as a subsidy on lending; second, deposit insurance weakens the degree to which the market induces banks to organize in a manner that will minimize the public good problem associated with relending to a problem debtor; finally, implicit deposit insurance removes much of the remaining liability side of the bank balance sheet from a private regulating role. Moreover, if

16. This share corresponds to that which existed on average from 1980 to 1985 according to Penati and Protopapadakis (1988).

17. These parameters have been chosen to insure that $0<\tau<1$.

18. The surprising result that the number of banks in the system actually declines with increases in $\phi$ stems from the zero-profit condition. Since increases in $\phi$ make lending less profitable, and individual bank lending remains constant, exit must occur for profits to return to zero.

TABLE 2

SIMULATION RESULTS

| | $\phi$ | $n$ | $l_1$ | $E(\pi_1)$ | $E(\psi_1/nl_1)$ |
|---|---|---|---|---|---|
| (1) NO DEPOSIT INSURANCE | | | | | |
| | 0.00 | 11.01 | 1.01 | 0.98 | — |
| | 0.00 | 10.78 | 1.01 | 0.98 | — |
| | 0.01 | 8.98 | 1.01 | 0.98 | — |
| | 0.02 | 6.74 | 1.00 | 0.99 | — |
| (2) EXPLICIT DEPOSIT INSURANCE ($\tau = 0.65$) | | | | | |
| | 0.00 | 15.56 | 1.01 | 0.97 | 0.02 |
| | 0.00 | 15.24 | 1.01 | 0.97 | 0.02 |
| | 0.01 | 12.70 | 1.01 | 0.97 | 0.02 |
| | 0.02 | 9.52 | 1.00 | 0.98 | 0.01 |
| (3) EXPLICIT PLUS IMPLICIT DEPOSIT INSURANCE | | | | | |
| | 0.00 | 17.52 | 1.01 | 0.97 | 0.03 |
| | 0.00 | 17.12 | 1.01 | 0.97 | 0.03 |
| | 0.01 | 13.94 | 1.01 | 0.97 | 0.02 |
| | 0.02 | 10.19 | 1.01 | 0.98 | 0.02 |
| (4) 100 PERCENT DEPOSIT INSURANCE | | | | | |
| | 0.00 | 20.00 | 1.01 | 0.96 | 0.04 |
| | 0.00 | 19.59 | 1.01 | 0.97 | 0.04 |
| | 0.01 | 16.33 | 1.01 | 0.97 | 0.03 |
| | 0.02 | 12.25 | 1.01 | 0.98 | 0.03 |

the deposit insurance institution's appraisal of the degree of systemic risk in a lending package is endogenous, banks will be rewarded for organizing themselves in a manner that enhances the probability of a bail-out.

Both private creditors and government officials of lending and borrowing countries have argued that the level of loan provision to the highly indebted countries during the debt crisis was sub-optimal from the point of view of the industry as a whole. Previous discussions explain underlending through "herd behavior" followed by banks (Herring and Guttentag 1985). This paper shows that sub-optimally large levels of banking "diffusion," rationally introduced to avoid firm risk and take advantage of fixed-premium deposit insurance, may exacerbate the degree to which credit extensions are sub-optimal, providing an alternative explanation to herd behavior.

## APPENDIX

### *I. Derivation of (3a) and (3b)*

The first-order condition from equation (2) satisfies:

$$-U'\bar{r}_1 L_1 + (\partial U/\partial P)(\partial P/\partial \pi_t) = 0; \quad (t = 1, 2)$$

where $U'$ represents $\partial U/\partial c_t$. Totally differentiating with respect to $\pi_1$ and $q_1$, $L_1$, and $L_2$ yields:

$$d\pi_1/dq_1 = U''\bar{r}_1 L_1/(d^2U/d\pi_1^2) > 0$$

$$d\pi_1/dL_1 = \bar{r}_1(U' - U''\pi_1\bar{r}_1 L_1)/(d^2U/d\pi_1^2) < 0$$

$$d\pi_1/dL_2 = U''\bar{r}_1 L_1/(d^2U/d\pi_1^2) > 0$$

$$d\pi_1/dn = d\pi_1/dL.$$

Taking the first-order condition from (2) and totally differentiating with respect to $\pi_2$ and $q_2$ and $L_2$ yields:

$$d\pi_2/dq_2 = U''\bar{r}_2 L_2/(d^2U/d\pi_2^2) > 0$$

$$d\pi_2/dL_2 = \bar{r}_2(U' - U''\bar{r}_2 L_2)/(d^2U/d\pi_2^2) < 0.$$

### *II. Second-Period Deposit Rates*

By equation (5) and the fact that $q_2$ is distributed uniform on the unit interval, $q_2^B$ satisfies:

$$\pi_2(q_2^B, L_2)\bar{r}_2\hat{l}_2 - \gamma_2(a_2 + \hat{l}_2) + \rho_2 a_2 = 0.$$

Totally differentiating with respect to $q_2^B$ and $\gamma_2$, $L_2$, and $\hat{l}_2$ yields:

$$dq_2^B/d\gamma_2 = (a_2 + \hat{l}_2)/(\partial\pi_2/\partial q_2)\bar{r}_2\hat{l}_2 > 0$$

$$dq_2^B/dL_2 = -(\partial\pi_2/\partial L_2)\bar{r}_2\hat{l}_2/(\partial\pi_2/\partial q_2)\bar{r}_2\hat{l}_2 > 0$$

$$dq_2^B/d\hat{l}_2 = -[(\partial\pi_2/\partial L_2)\bar{r}_2\hat{l}_2 + \pi_2\bar{r}_2 - \gamma_2]/(\partial\pi_2/\partial q_2)\bar{r}_2\hat{l}_2.$$

By equation (7), $dq_2^B/d\hat{l}_2$ is of ambiguous sign because of firm risk-aversion. Intuitively, the ambiguity stems from the possibility that firms are sufficiently risk-averse that additional second-term loans are privately (as opposed to globally among creditors as a whole) profitable. Since we are interested in the case where bank lending falls below its optimum, we rule out this possibility. We proceed under the assumption that the numerator of that expression is negative, i.e., that profits on second-period loans, neglecting their impact on first period debt service, are negative. This leaves the entire expression positive, implying that additional bank lending raises the possibility of future bankruptcy.

Totally differentiating (4) with respect to $\gamma_2$ and $\hat{l}_2$, $L_2$, and $\tau$ and simplifying, we then obtain:

$$d\gamma_2/d\hat{l}_2 = (\partial q_2^B/\partial\hat{l}_2)/\{[(1-q_2^B)^2/r_2(1-\tau)] - (\partial q_2^B/\partial\gamma_2)\} > 0$$

$$d\gamma_2/dL_2 = (\partial q_2^B/\partial L_2)/\{[(1-q_2^B)^2/r_2(1-\tau)] - (\partial q_2^B/\partial\gamma_2)\} > 0$$

$$d\gamma_2/d\tau = -q_2^B/\{[(1-q_2^B)/r_2] - [(1-\tau)/(1-q_2^B)](\partial q_2^B/\partial\gamma_2)\} < 0,$$

since the denominators of all three are positive when returns to depositors are increasing in $\gamma_2$.

## *III. Comparative Statics Concerning Second-Period Lending Decisions*

Totally differentiating the first-order condition from (6) yields:

$$d\hat{l}_2/d\hat{l}_1 = -(\partial^2\Omega_2/\partial\hat{l}_2\partial\hat{l}_1)/(\partial^2\Omega/\partial\hat{l}_2^2).$$

By the second-order condition, the denominator is negative so that:

$$Sign[d\hat{l}_2/d\hat{l}_1] = Sign[\partial^2\Omega/\partial\hat{l}_2\partial\hat{l}_1].$$

By (7):

$$Sign\left[d\hat{l}_2/d\hat{l}_1\right] = Sign\left[\frac{\partial\Omega_2}{\partial R}\left[\frac{\partial\pi_1}{\partial L_2}\bar{r}_1 + \frac{\partial^2\pi_1}{\partial L_2\partial\hat{l}_1}\bar{r}_1\hat{l}_1\right]\right] > 0\ ,$$

where $\partial^2\pi_1/\partial L_2\partial\hat{l}_1 > 0$ from our solution for $\partial\pi_1/\partial L_2$ above.

As above, $Sign[dl_2/dL_1] = Sign[\partial^2\Omega/\partial\hat{l}_2\partial L_1]$. By (7):

$$Sign\left[d\hat{l}_2/dL_1\right] = Sign\left[\frac{\partial\Omega_2}{\partial R}\ \frac{\partial^2\pi_1}{\partial L_2\partial L_1}\bar{r}_1\hat{l}_1\right] < 0\ ,$$

where $\partial^2\pi_1/\partial L_2\partial L_1 < 0$ from our solution for $\partial\pi_2/\partial L_2$ above.

Similarly, $Sign[d\hat{l}_2/dL_2] = Sign[\partial^2\Omega/\partial\hat{l}_2\partial L_2]$. By (7):

$$Sign\left[\frac{\partial^2\Omega}{\partial\hat{l}_2\partial L_2}\right] = Sign\left[\frac{\partial\Omega_2}{\partial R}\left[\frac{\partial^2\pi_1}{\partial L_2^2}\bar{r}_1\hat{l}_1 + \frac{1}{r_2}\left[\frac{\partial\pi_2}{\partial L_2}\bar{r}_2 - \gamma_2 + \frac{\partial^2\pi_2}{\partial L_2^2}\bar{r}_2\hat{l}_2 - \frac{\partial^2\gamma_2}{\partial\hat{l}_2\partial L_2}(a_2+\hat{l}_2)\right]\right]\right] < 0.$$

Similarly, $Sign[d\hat{l}_2/d\tau] = Sign[\partial^2\Omega/\partial\hat{l}_2\partial\tau]$. By (7):

$$Sign\left[\frac{\partial^2\Omega}{\partial\hat{l}_2\partial\tau}\right] = Sign\left[-\frac{\partial\Omega_2}{\partial R}\left[\frac{\partial\gamma_2}{\partial\tau} + \frac{\partial^2\gamma_2}{\partial\hat{l}_2\partial\tau}(a_2+\hat{l}_2)\right]/r_2\right] > 0,$$

where $\partial^2\gamma_2/\partial\hat{l}_2\partial\tau < 0$ from our solutions for $d\gamma_2/d\hat{l}_2$ above.

Finally, $Sign[d\hat{l}_2/dn] = Sign[\partial^2\Omega/\partial\hat{l}_2\partial n]$. By (7):

$$Sign\left[\frac{\partial^2\Omega}{\partial\hat{l}_2\partial n}\right] = Sign\left[\frac{\partial\Omega_2}{\partial R}\left[\frac{\partial^2\pi_1}{\partial L_2\partial n}\bar{r}_1 l_1 + \frac{1}{r_2}\left[\frac{\partial\pi_2}{\partial n}\bar{r}_2 - \frac{\partial\gamma_2}{\partial n} - \frac{\partial^2\gamma_2}{\partial\hat{l}_2\partial n}(a_2+\hat{l}_2)\right]\right]\right] < 0\ .$$

## *IV. First-Period Deposit Rates*

By equation (5) and the fact that $q_2$ is distributed uniform on the unit interval, $q_1^B$ satisfies:

$$\pi_1(q_1^B, L_1, L_2)\bar{r}_1\hat{l}_1 - \gamma_1(a_1 + \hat{l}_1) + \rho_1 a_1 = 0.$$

Totally differentiating (5) with respect to $q_1^B$ and $\hat{l}_1$ yields:

$$dq_1^B/d\hat{l}_1 = -[(\partial\pi_1/\partial\hat{l}_1)\bar{r}_1\hat{l}_1 + \pi_1\bar{r}_1 - \gamma_1]/(\partial\pi_1/\partial q_1)\bar{r}_1\hat{l}_1 > 0$$

in the range in which positive lending takes place since the individual bank returns on first-period lending must be positive in the presence of bank risk-aversion. Totally differentiating (4) with respect to $\gamma_1$ and $\hat{l}_1$ then yields:

$$d\gamma_1/d\hat{l}_1 = (\partial q_1^B/\partial\hat{l}_1)/\{[(1-q_1^B)^2/r_1(1-\tau)] - (\partial q_1^B/\partial\gamma_1)\} > 0.$$

As above, the denominator of these terms is positive in the relevant range where returns to depositors are increasing in $\gamma_1$.

Totally differentiating (5) with respect to $q_1^B$ and $l_1$ yields:

$$dq_1^B/dl_1 = -n(\partial\pi_1/\partial L_1)/(\partial\pi_1/\partial q_1) > 0,$$

since $\partial\pi_1/\partial L_1 < 0$ as shown above. Totally differentiating (4) with respect to $\gamma_1$ and $\hat{l}_1$ then yields:

$$d\gamma_1/dl_1 = (\partial q_1^B/\partial L_1)/\{[(1-q_1^B)^2/r_1(1-\tau)] - (\partial q_1^B/\partial\gamma_1)\} > 0.$$

Taking $l_1$ as given, totally differentiating the debtor's first-order condition from (2) (shown above) with respect to $\pi_1$ and $n$ yields:

$$d\pi_1/dn = (d\pi_1/dL_1)(dL_1/dn) + (d\pi_1/dL_2)(dL_2/dn).$$

Substituting, recalling that $L_2 = nl_2$ ,

$$d\pi_1/dn = \bar{r}_1 l_1\,(U' - U''\pi_1\bar{r}_1 L_1)/(d^2U/d\pi_1^2) + U''\bar{r}_1 L_1/(d^2U/d\pi_1^2)[l_2 + n(dl_2/dn)],$$

where $dl_2/dn < 0$ as shown above. Simplifying:

$$d\pi_1/dn = \{l_1U' - U''L_1[\pi_1\bar{r}_1 - l_2 - n(dl_2/dn)]\}\bar{r}_1/(d^2U/d\pi_1^2).$$

This term is of ambiguous sign because it is unclear whether an increase in $n$ results in an increase or decrease in second-period lending, which has a positive impact on first-period debt service. The ambiguity corresponds to the fact that an increase in $n$ results in decreased lending per bank, but more banks in the lending package. We proceed under the assumption that a collective action problem exists, i.e., that an increase in $n$ results in a decrease in first-period debt service. This requires that the above expression be negative. Under this condition, totally differentiating (5) with respect to $q_1^B$ and $n$ yields:

$$dq_1^B/dn = -\, n(\partial\pi_1/\partial n)/(\partial\pi_1/\partial q_1) > 0.$$

Totally differentiating (4) with respect to $\gamma_1$ and $n$ then yields:

$$d\gamma_1/dn = (\partial q_1^B/\partial n)/\{[(1-q_1^B)^2/r_1(1-\tau)] - (\partial q_1^B/\partial\gamma_1)\} > 0.$$

Moreover, note that:

$$\partial^2\gamma_1/\partial n\partial\tau = \frac{-\,(\partial q_1^B/\partial n)[(1-q_1^B)^2/r_1(1-\tau)^2]}{\{[(1-q_1^B)^2/r_1(1-\tau)] - (\partial q_1^B/\partial\gamma_1)\}^2} < 0.$$

Finally, totally differentiating (4) with respect to $\gamma_1$ and $\tau$ yields:

$$d\gamma_1/d\tau = -\, q_1^B/\{[(1-q_1^B)/r_1] - [(1-\tau)/(1-q_1^B)](\partial q_1^B/\partial\gamma_1)\} < 0.$$

## *V. First-Period Equilibrium*

In signing (17), by the first-order condition, $\partial\Omega_1/\partial\hat{l}_1 = 0$. By the second-order condition, $\partial^2\Omega_1/\partial\hat{l}_1^2 < 0$. By (15), $\partial^2\Omega_1/\partial\hat{l}_1\partial n$ satisfies:

$$\begin{aligned}\frac{\partial^2\Omega_1}{\partial\hat{l}_1\partial n} = \frac{\partial\Omega_1}{\partial R}\Bigg[ & \frac{\partial^2\pi_1}{\partial L_1^2}\bar{r}_1\hat{l}_1^2 + \frac{\partial\pi_1}{\partial L_1}\hat{l}_1\pi_1\bar{r}_1 \\ & - \frac{\partial\gamma_1}{\partial L_1}\hat{l}_1 - \frac{\partial^2\gamma_1}{\partial\hat{l}_1\partial n}(a_1+\hat{l}_1) + \Bigg[\frac{\partial\pi_2}{\partial L_2}\bar{r}_2 - \frac{\partial\gamma_2}{\partial L_2} \\ & + \frac{\partial^2\pi_2}{\partial L_2^2}\bar{r}_2\hat{l}_2 - \frac{\partial^2\gamma_2}{\partial\hat{l}_2\partial L_2}(a_2+\hat{l}_2)\Bigg]\frac{\partial L_2}{\partial n}\,\frac{\partial\hat{l}_2}{\partial\hat{l}_1}/r_2 \\ & + \Bigg[\pi_2\bar{r}_2 - \gamma_2 + \frac{\partial\pi_2}{\partial L_2}\bar{r}_2\hat{l}_2 - \frac{\partial\gamma_2}{\partial\hat{l}_2}(a_2+\hat{l}_2)\Bigg] \\ & \bullet\ \frac{\partial^2\hat{l}_2}{\partial\hat{l}_1\partial n}/r_2\Bigg],\end{aligned}$$

which is of ambiguous sign. A sufficient but not necessary condition for the expression to be negative is that $\partial L_2\partial n > 0$. In other words, despite the fact that each bank lends less in the final period, the increase in the number of banks implies that the total level of new lending increases. We proceed by accepting this condition, under which the entire expression can be signed as negative.

By (5) and (14) $\partial\Omega_1/\partial n$ satisfies:

$$\begin{aligned}\partial\Omega_1/\partial n = (\partial\Omega_1/\partial R)\{&[(\partial\pi_1/\partial n)\bar{r}_1\hat{l}_1 \\ &- (\partial\gamma_1/\partial n)(a_1+\hat{l}_1)] + (1/r_2)[(\partial\pi_2/\partial n)\bar{r}_2 l_2 \\ &- (\partial\gamma_2/\partial n)(a_2+\hat{l}_2)]\} < 0.\end{aligned}$$

It follows that the determinant of the system is positive.

$\partial^2\Omega_1/\partial l_1\partial\tau$ satisfies:

$$\begin{aligned}\partial\Omega_1/\partial l_1\partial\tau = \partial\Omega/\partial R\,&[\partial^2 R_1/\partial l_1\partial\tau \\ &+ (1/r_f)\partial^2 R_2/\partial l_1\partial\tau] > 0,\end{aligned}$$

where:

$$\begin{aligned}\partial^2 R_t/\partial l_1\partial\tau &= -\,\partial\gamma_t/\partial\tau - (\partial^2\gamma_t/\partial l_1\partial\tau)(a_t+l_t) \\ &> 0; \\ &(t = 1, 2)\end{aligned}$$

since:

$$\partial^2\gamma_t/\partial l_1\partial\tau = -\,[1/(1-q_t^B)](\partial q_t^B/\partial l_t) < 0.$$

$\partial\Omega_1/\partial\tau$ satisfies:

$$\begin{aligned}\partial\Omega_1/\partial\tau = \partial\Omega/\partial R[&-(\partial\gamma_1/\partial\tau)(a_1+l_1) \\ &- (1/r_2)(\partial\gamma_2/\partial\tau)(a_2+l_2) > 0.\end{aligned}$$

By Cramer's rule:

$$\begin{aligned}\partial n/\partial\tau &= -\,(\partial^2\Omega_1/\partial\hat{l}_1^2)(\partial\Omega_1/\partial\tau)/D > 0 \\ \partial l_1/\partial\tau &= -\,(\partial^2\Omega_1/\partial\hat{l}_1\partial\tau)(\partial\Omega_1/\partial n) \\ &\quad + (\partial^2\Omega_1/\partial\hat{l}_1\partial n)(\partial\Omega_1/\partial\tau)]/D = 0,\end{aligned}$$

where $D$ represents the determinant of the system.

Similarly, for the implicit deposit insurance comparative static exercise in (21):

$$\begin{aligned}\partial^2\Omega_1/\partial\hat{l}_1\partial\delta = \partial\Omega/\partial R[&\partial^2 R_1/\partial\hat{l}_1\partial\delta \\ &+ (1/r_2)\partial^2 R_2/\partial\hat{l}_1\partial\delta] > 0\end{aligned}$$

where:

$$\begin{aligned}\partial^2 R_t/\partial\hat{l}_1\partial\delta &= -\,\partial\gamma_t/\partial\delta - (\partial^2\gamma_t/\partial l_1\partial\delta)(a_t+l_t) \\ &> 0; \\ &(t = 1, 2)\end{aligned}$$

since:

$$\partial\gamma_t/\partial\delta = -[r_t n l_1(1-\tau)/(1-q_t^B) < 0; \quad (t = 1, 2)$$

and:

$$\partial^2\gamma_1/\partial l_1\partial\delta = -[\bar{r}_1 n l_1(1-\tau)/(1-q_1^B)](\partial q_1^B/\partial l_1) < 0$$

$$\partial\Omega_1/\partial\delta = \partial\Omega/\partial R[-(\partial\gamma_1/\partial\delta)(a_1+\hat{l}_1) - (1/r_2)(\partial\gamma_2/\partial\delta)(a_2+\hat{l}_2)] > 0.$$

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# An Analysis of Inefficiencies in Banking: A Stochastic Cost Frontier Approach

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*This paper examines the properties of X-inefficiency and the relations of X-inefficiency with risk-taking and stock returns for U. S. banking firms. After controlling for scale differences, the average small size banking firm is found to be relatively less efficient than the average large firm. Smaller firms also exhibit higher variations in X-inefficiencies than their larger counterparts. While the average X-inefficiency appears to be declining over time, the rank orderings of X-inefficiency are found to be quite persistent. Furthermore, less efficient banking firms are found to be associated with higher risk-taking, and firm-specific X-inefficiencies are significantly correlated with individual stock returns for smaller banking firms.*

The efficiency of banking organizations has been studied extensively in the banking literature. Earlier studies tended to focus on the issues of scale and scope efficiencies. Scale efficiency refers to the relationship between a firm's average cost and output. Detection of a U-shaped average cost curve suggests that there is an optimal scale of production, at which point the production cost would be minimized. Scope efficiency refers to the economies of joint production, where the costs of producing joint products are less than the sum of their stand-alone production costs. Though extensive, the studies of the scale and scope efficiencies of financial institutions to date do not seem to provide conclusive evidence on the economic significance of these types of inefficiencies in U.S. banking firms.

More recently, research on banking efficiency has devoted more attention to the issue of X-inefficiency. X-inefficiency refers to the deviations from the production-efficient frontier which depicts the maximum attainable output for a given level of input. The concept of X-inefficiency was introduced by Leibenstein (1966), who noted that, for a variety of reasons, people and organizations normally work neither as hard nor as effectively as they could. When applied to U.S. banking firms, research to date suggests that X-inefficiencies appear to be large and tend to dominate scale and scope inefficiencies.[1]

Because most of the studies of X-inefficiencies were based on cross-sectional analyses, the time-series properties of X-inefficiencies in U.S. banking firms have not been well-documented. There is little information on how X-inefficiencies in banking may evolve over time in response to market forces and on how the rankings of X-inefficiency of individual banking firms may change over time. These issues are especially interesting given the substantial changes in banking markets and banking regulations that have occurred during the past decade. For instance, if inefficient banking firms have a tendency to remain inefficient, it would be of interest to investigate how they can

1. In their summary of recent research, Berger, Hunter, and Timme (1993) indicated that X-inefficiencies in banking account for approximately 20 percent or more of banking costs, while scale and scope efficiencies—when they can be accurately estimated—are usually found to account for less than 5 percent of costs. See also Berger and Humphrey (1991).

remain economically viable and not be driven out of the banking market. Policymakers would be concerned about whether inefficient banking firms pose additional risks to the banking system and its safety net. Investors would be interested in the relationship between the firm-specific X-inefficiencies and the market valuation of bank stocks.

To examine these issues, we estimate a stochastic cost-efficient frontier à la Aigner, Lovell, and Schmidt (1977) based on a multiproduct translog cost function. Semiannual data for a sample of 254 bank holding companies from 1986 to 1991 are grouped into size-based quartiles to allow for different production technologies for each size class. Separate cost functions are estimated for each size quartile using the method of maximum likelihood. An estimate of X-inefficiency for each sample firm at each sample period is then derived following the method of Jondrow, Lovell, Materov, and Schmidt (1982).

As in the cross-section results reported in earlier studies, we find that X-inefficiencies are quite large. Furthermore, several interesting properties of X-inefficiencies also are detected. First, both the level of X-inefficiencies and their cross-sectional variations are, on average, noticeably smaller for large banking firms than for smaller firms. Second, regardless of firm size, X-inefficiencies appear to have declined gradually between 1986 and 1990, and then edged upward during 1991. Third, despite the decline in X-inefficiencies, the rank orderings of firm-specific X-inefficiencies are highly correlated over time. Specifically, the rank ordering persists for approximately three and one-half years for the sample firms that are in the three smaller size quartiles, and for about one year for the sample firms that are in the largest size quartile.

The finding that based on rank ordering, inefficient banking firms tend to stay inefficient leads us to investigate how these inefficient firms can be economically viable, if banking markets are truly contestable and efficient. This is especially puzzling given recent changes that suggest increased competition and substantial entry by non-banking firms in financial markets. We hypothesize that many banking markets may be effectively insulated, at least during the time period of this study, which enables inefficient firms to continue to survive by earning economic rents. Perhaps more importantly, with fixed premium deposit insurance during our sample period, inefficient firms may be induced to compensate for their inefficiencies by extracting subsidies from the FDIC through greater risk-taking.[2] Moreover, the managers of inefficient banking firms, who are more likely to be entrenched, may be inclined to take on more risk (Gorton and Rosen 1995). Finally, it is possible that bank regulators may exacerbate this risk-taking incentive by delaying much needed regulatory actions on problem institutions (see, for example, Kane 1992, Kane and Kaufman 1993). Taken together, the hypothesis that inefficient banking firms may be associated with higher risk-taking seems plausible.

We find a strong association between our X-inefficiency estimates and various proxies for bank risk-taking in all four size classes. Specifically, inefficient firms tend to have higher common stock return variance, higher idiosyncratic risk in stock returns, lower capitalization, and higher loan charge-offs. Furthermore, firm-specific X-inefficiencies are found to have explanatory power for banking firms' stock returns, after controlling for the stock market return and changes in the riskless interest rate.

The remainder of this paper is organized as follows: Section I describes the approach we use to estimate firm-specific X-inefficiency. Section II outlines the data used in this study. The properties of the estimated X-inefficiency for our sample banking firms are discussed in Section III. Section IV examines the relationship between X-inefficiency and bank risk-taking. Section V investigates the relationship between X-inefficiency and bank stock returns. Section VI summarizes and concludes this paper.

## I. Measuring X-Inefficiency in Banking

To measure the X-inefficiency of individual banking firms, we use the stochastic efficient frontier methodology of Aigner, Lovell, and Schmidt (1977). In this method, a banking firm's observed total cost is modeled to deviate from the cost-efficient frontier due to random noise and possibly X-inefficiency. For the $n$th firm,

$$\ln TC_n = f(\ln Q_i, \ln P_j) + \varepsilon_n \tag{1}$$

where $TC_n$ is the total cost for firm $n$, $Q_i$ are measures of banking output, and $P_j$ are input prices. In equation (1), $\varepsilon_n$ is a two-component disturbance term of the form:

$$\varepsilon_n = \mu_n + \delta_n\,, \tag{2}$$

where $\mu_n$ represents a random uncontrollable factor and $\delta_n$ is the controllable component of $\varepsilon_n$. In equation (2), $\mu_n$ is independently and identically distributed normal with zero mean and $\sigma_\mu$ standard deviation, i.e., $N(0,\sigma_\mu^2)$. The term $\delta_n$ is distributed independently of $\mu_n$ and has a half-normal

2. The moral hazard of fixed-premium deposit insurance has long been recognized in the banking literature (see for example Merton 1977, Marcus 1984, and Keeley 1990). Furthermore, Marcus and Shaked (1984), Ronn and Verma (1986), and Pennacchi (1987) provide evidence on the mispricing of deposit insurance.

distribution, i.e., $\delta_n$ is the absolute value of a variable that is normally distributed with zero mean and standard deviation $\sigma_\delta$, $N(0,\sigma_\delta^2)$.

The X-inefficiency of firm $n$, defined as $c_n$, can be expressed as the expected value of $\delta_n$ conditional on $\varepsilon_n$ (Jondrow, Lovell, Materov, and Schmidt 1982):

$$(3) \quad c_n = E(\delta_n|\varepsilon_n) = [\sigma\lambda/(1+\lambda^2)][\phi(\varepsilon_n\lambda/\sigma)/\Phi(\varepsilon_n\lambda/\sigma) + \varepsilon_n\lambda/\sigma] ,$$

where $\lambda$ is the ratio of the standard deviation of $\delta_n$ to the standard deviation of $\mu_n$ (i.e., $\sigma_\delta/\sigma_\mu$), $\sigma^2 = \sigma^2_\delta + \sigma^2_\mu$, $\Phi$ is the cumulative standard normal density function, and $\phi$ is the standard normal density function. Estimates of $c_n$ are obtained by evaluating equation (3) at the estimates of $\sigma^2_\delta$ and $\sigma^2_\mu$.

To specify the cost function in equation (1), we employ the following multiproduct translog cost function:

$$(4) \ln TC = \alpha_0 + \Sigma_i\alpha_i\ln Q_i + \Sigma_j\beta_j\ln P_j + {}^1/_2\Sigma_i\Sigma_k\gamma_{ik}\ln Q_i\ln Q_k + {}^1/_2\Sigma_j\Sigma_h\zeta_{jh}\ln P_j\ln P_h + \Sigma_i\Sigma_j\omega_{ij}\ln Q_i\ln P_j ,$$

where $TC$ is total operating costs (including interest costs), $Q_i$ are outputs, and $P_j$ are input prices. Five measures of banking outputs are included: book value of investment securities ($Q1$), book value of real estate loans ($Q2$), book value of commercial and industrial loans ($Q3$), book value of consumer loans ($Q4$), and off-balance sheet commitments and contingencies ($Q5$) which include loan commitments, letters of credit (both commercial and standby), futures and forward contracts, and notional value of outstanding interest rate swaps. Three input prices are utilized: the unit price of capital ($P1$) measured as total occupancy expenses divided by fixed plant and equipment, the unit cost of funds ($P2$) defined as total interest expenses divided by total deposits, borrowed funds, and subordinated notes and debentures, and the unit price of labor ($P3$), defined as total wages and salaries divided by the number of full-time equivalent employees. The linear homogeneity restrictions,

$$\Sigma_j\beta_j = 1, \quad \Sigma_h\zeta_{jh} = 0, \forall j, \quad \Sigma_j\omega_{ij} = 0, \forall i,$$

are imposed by normalizing the total cost and the input prices by the price of labor. To allow the cost function to vary across size classes, the sample banking firms are first sorted into size-based quartiles according to average total assets between 1986 and 1991. Assuming the cost function to be stationary over time, pooled time-series cross-section observations are used to estimate the stochastic cost frontier separately for each size-based quartile by the method of maximum likelihood. Estimates of $c_n$, which represent the measure of firm-specific X-inefficiency, are then computed for each sample firm in each sample period.

## II. DATA

Semiannual bank holding company data from 1986 through 1991 are obtained from the Federal Reserve FR Y-9C Bank Holding Company Reports. Since only bank holding companies with total consolidated assets of $150 million or more or with more than one subsidiary bank are required to file the FR Y-9C Report, our sample consists mainly of larger banking organizations. Daily stock price data for our sample bank holding companies are obtained from the Center for Research in Security Prices (CRSP) at the University of Chicago.

Our sample consists of 254 bank holding companies, of which 174 had complete time-series data from 1986 through 1991. The average total assets of the 174 sample firms with a complete time series of observations are used to sort these firms into size-based quartiles. The remaining 80 sample firms with an incomplete time series of observations are then classified into respective size classes using the quartile break points established by the 174 firms at matching time periods. This classification method ensures that the sample firms stay in the same size class throughout the study period, which is necessary to study the time-series properties of X-inefficiency.[3]

Table 1 reports the summary statistics of banking outputs, input prices, total assets, and total costs for the 254 sample banking firms. Both firm size and the cost function variables are highly skewed, indicating the desirability of grouping firms into size classes. In addition, off-balance sheet activities tend to be concentrated in the larger firms in the sample, further suggesting that the cost functions of large banking firms may be different from those of smaller firms.

## III. PROPERTIES OF X-INEFFICIENCY IN BANKING

Table 2 reports summary statistics of the estimates of $c_n$ in equation (3). These firm-specific X-inefficiency estimates are derived from the stochastic cost frontier estimated separately for banking firms in each size-based quartile. Consistent with earlier studies, we find that substantial inefficiencies exist in banking, averaging between 10 to 20 percent of total costs. However, after controlling for scale

3. Potential misclassification due to intertemporal size changes of individual firms does not seem to be a major concern. If the sample firms had been permitted to move freely from size class to size class intertemporally, there would have been 69 instances of firms moving up to the next size class (of which 51 are within 10 percent of the quartile break points), and 77 instances of firms moving down to the next size class (of which 72 are within 10 percent of the quartile break points).

differences, both the mean and the median estimates of inefficiency decrease monotonically from Quartile 1 to Quartile 4. This suggests that, on average, smaller bank holding companies deviate more from their respective cost-efficient frontier than do larger bank holding companies. Relatively speaking, smaller banking firms appear to be less efficient than their larger counterparts. Moreover, both the intra-quartile range and the standard deviation of inefficiency decrease with firm size. Hence, not only are smaller firms relatively less efficient than larger firms, but their variations in X-inefficiencies also seem to be higher than their larger counterparts. Interestingly, Table 2 also

TABLE 1

DATA SUMMARY FOR 254 BANK HOLDING COMPANIES, BASED ON SEMIANNUAL DATA FROM 1986 TO 1991

| | 25TH PERCENTILE | MEDIAN | MEAN | 75TH PERCENTILE |
|---|---|---|---|---|
| Total assets[a] | 1,198,481 | 2,779,545 | 9,814,536 | 8,110,207 |
| Commercial and industrial loans[a] | 164,143 | 434,074 | 1,657,808 | 1,435,509 |
| Real estate loans[a] | 306,258 | 689,684 | 2,136,602 | 1,857,829 |
| Consumer loans[a] | 139,356 | 345,852 | 1,178,900 | 957,541 |
| Investment securities[a] | 266,438 | 613,962 | 1,407,576 | 1,480,544 |
| Commitments & contingencies[a,e] | 71,486 | 307,048 | 17,684,563 | 1,984,561 |
| Total costs[a] | 50,644 | 121,354 | 462,233 | 346,316 |
| Price of labor[b] | 12.41 | 14.02 | 14.85 | 16.08 |
| Price of physical capital[c] | 0.126 | 0.166 | 0.180 | 0.219 |
| Price of funds[d] | 0.025 | 0.027 | 0.028 | 0.030 |
| Number of observations | | 2,733 | | |

[a] in thousands of dollars.
[b] in thousands of dollars per full-time equivalent employee.
[c] in thousands of dollars per thousands of dollars of fixed assets.
[d] in thousands of dollars per thousands of dollars of deposits and borrowed funds.
[e] includes loan commitments, letters of credit, futures and forward contracts, and notional value of outstanding interest rate swaps.

TABLE 2

SUMMARY STATISTICS OF X-INEFFICIENCY

| | QUARTILE 1 | QUARTILE 2 | QUARTILE 3 | QUARTILE 4 |
|---|---|---|---|---|
| Mean | 0.1855 | 0.1446 | 0.1211 | 0.0808 |
| Median | 0.1483 | 0.1166 | 0.1003 | 0.0704 |
| Minimum | 0.0146 | 0.0197 | 0.0159 | 0.0208 |
| Maximum | 0.9460 | 0.6144 | 0.4708 | 0.3212 |
| Std. Deviation | 0.1454 | 0.0977 | 0.0819 | 0.0417 |
| Skewness | 1.6447 | 1.4156 | 1.2244 | 1.4741 |
| Kurtosis | 3.1797 | 2.4199 | 1.4317 | 3.0111 |
| *N* | 774 | 657 | 643 | 659 |

Note: Quartile 1 (4) contains the smallest (largest) firms.

shows that the X-inefficiency estimates are positively skewed and that they are more fat-tailed for firms in Quartiles 1 and 4.

Figure 1 depicts the 10th and 90th percentile of the X-inefficiency estimates at each semiannual subperiod for the 174 firms that have complete time-series of inefficiency estimates. In addition to confirming that controllable firm-specific inefficiency tends to be relatively larger and to have higher variation among smaller banking firms, Figure 1 indicates that the median X-inefficiency estimate exhibits a gradual decline from 1986 to mid-1990, and then turns up slightly during the last three quarters of the sampling period. The decline in inefficiency from 1986 through 1990 suggests that the market and regulatory changes in banking during the 1980s may have forced banking firms to respond to increased competition in banking by operating more efficiently. While the slight increase in inefficiency since 1990 is somewhat puzzling, the observed pattern may be related to regulatory developments that occurred during this period. First, the increase in inefficiency may be partially driven by the steep rise in deposit insurance premiums, from 8.33 cents per $100 of domestic deposits in 1989 to 23 cents per $100 of domestic deposits in 1992. This structural change in banking costs may not be fully reflected by $\mu_n$ in equation (2) and may spill over into $\delta_n$, resulting in higher estimated inefficiencies. Second, the increase in capital requirements as a result of the 1988 Basle Capital Accord may lead to spurious estimates of X-inefficiency.[4] It is possible that banking firms may have responded to the risk-weighted capital requirement by rebalancing their product mix, for example, by shifting from loans to investment securities.[5] While the shift in product mix may be an efficient way to address the new capital constraint, this shift can result in higher observed inefficiency if, for example, the factors of loan production cannot be quickly adjusted to the new product mix.

The final property of X-inefficiency to be investigated in this section is the issue of persistence. Specifically, we are interested in examining the temporal relationship of the cross-sectional rankings of individual firms' inefficiency estimates. Table 3 reports the Spearman rank correlations of the estimated inefficiencies for firms which have a complete time series of data between June 1986 and eleven subsequent time periods. In Quartiles 1, 2, and 3, the rank orderings of X-inefficiency are significantly correlated over time at the 1 percent level for seven subperiods, suggest-

4. The Accord requires that the minimum standard ratio of capital to weighted risk assets be 8 percent, of which the core capital element must be at least 4 percent to be effective at the end of 1992.

5. Some banking observers further attribute this portfolio shift to the so-called credit crunch in 1990.

FIGURE 1A

QUARTILE 1 FIRMS

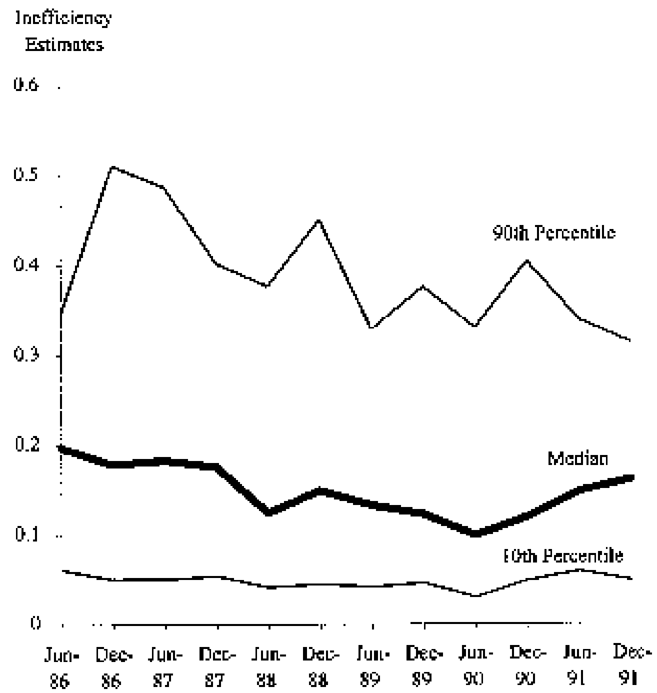



FIGURE 1B

QUARTILE 2 FIRMS

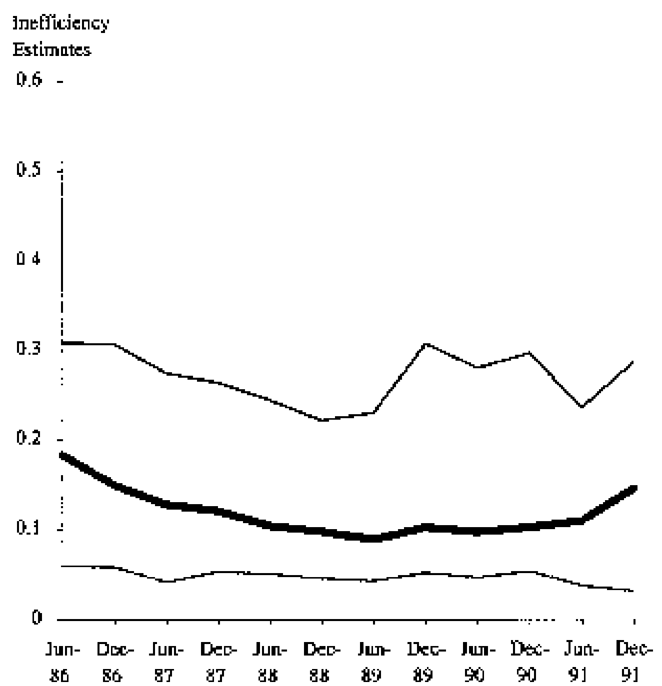


FIGURE 1C

Quartile 3 firms

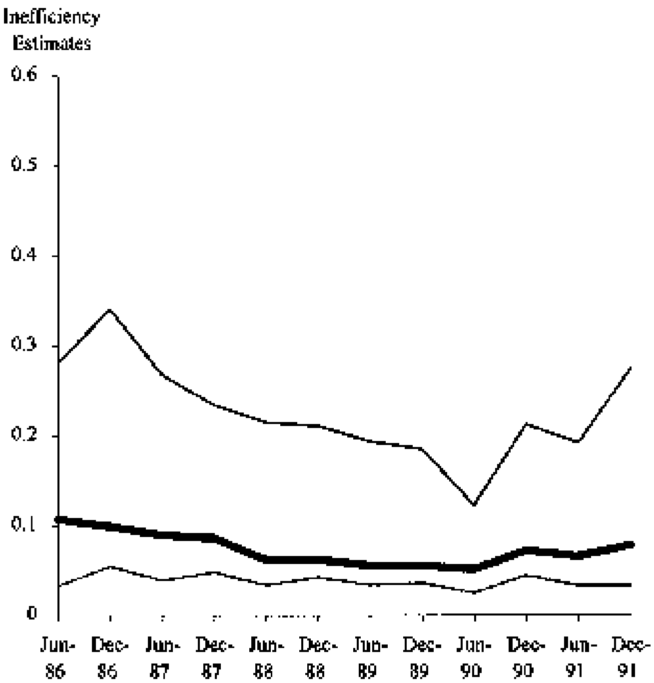
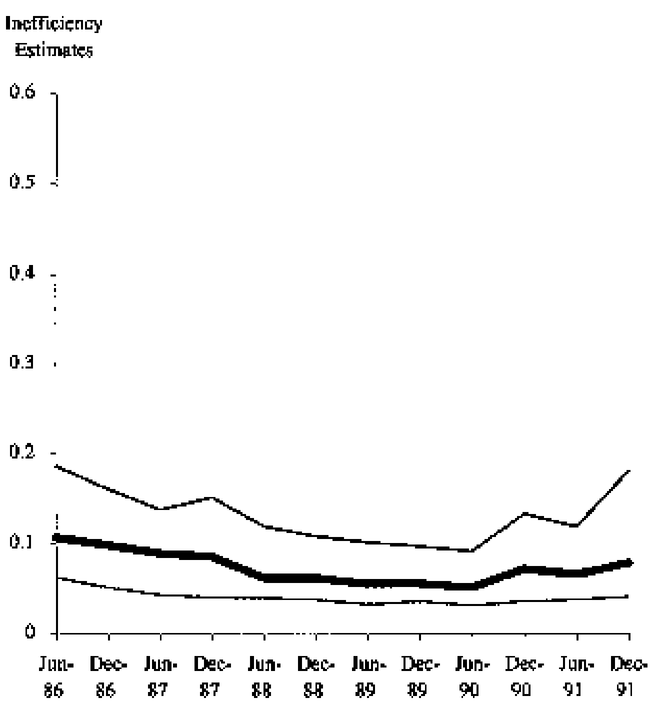



FIGURE 1D

Quartile 4 firms




ing that the ranking of firm-specific inefficiency persists for up to three and one-half years. For the largest firms in Quartile 4, the rank orderings of X-inefficiency are significantly correlated at the 1 percent level for only two subperiods, indicating that the ranking of X-inefficiency is relatively short-lived for large banking firms. Qualitatively similar results are obtained when different reference periods are used.

The findings in Table 3 again imply that the properties of the controllable firm-specific X-inefficiency for the very large banking firms are quite different from those of the smaller ones. The very large banking firms, on average, seem to operate closer to their respective efficient frontiers, and their firm-specific X-inefficiency appears to be transitory. In contrast, the smaller firms, on average, tend to operate further away from their respective frontiers, and their firm-specific X-inefficiency appears to be more permanent.

## IV. X-Inefficiency and Bank Risk-Taking

The apparent persistence of X-inefficiency, at least among the smaller banking firms, prompts us to investigate how inefficient firms can remain economically viable, especially if financial markets are efficient. Specifically, do inefficient firms do anything differently to compensate for being off the efficient frontier? In this paper, we investigate one plausible linkage between controllable X-inefficiency and firm behavior, namely, bank risk-taking. With fixed premium deposit insurance, the moral hazard hypothesis postulates that a bank insured by the FDIC may be able to increase the option value of deposit insurance by increasing bank risk. Theoretically, deposit insurance can be modeled as a put option written by the FDIC to the bank (Merton 1977). For simplicity, assuming all bank debts are insured at face value, in the event of insolvency, an insured bank can put the bank's assets to the FDIC at the face value of its debts, and the value of this put option increases with the bank's asset risk. However, not all banks engage in risk-maximizing behavior. The valuable bank charter, which will be lost upon failure, limits bank risk-taking (Marcus 1984 and Keeley 1990). To the extent that an inefficient banking organization may have a lower charter value to be preserved, it may be more prone to risk-taking than an efficient banking firm. Thus, it would be interesting to find out whether inefficient firms are associated with a higher level of risk.

We use five measures of bank risk, of which three are market-based and two are accounting-based. The three market measures of risk are: (i) standard deviation of daily stock returns, which reflects the total systematic and nonsystematic risks of the banking firm's common stock; (ii)

## TABLE 3

SPEARMAN RANK CORRELATION COEFFICIENTS OF INEFFICIENCY ESTIMATES AT JUNE 1986 AND SUBSEQUENT TIME PERIODS

| TIME PERIOD | QUARTILE 1 | QUARTILE 2 | QUARTILE 3 | QUARTILE 4 |
|---|---|---|---|---|
| Dec. 86 | 0.7809*** | 0.7862*** | 0.8003*** | 0.6951*** |
| June 87 | 0.7792*** | 0.7171*** | 0.6727*** | 0.4737*** |
| Dec. 87 | 0.7377*** | 0.6192*** | 0.4665*** | 0.2987* |
| June 88 | 0.6070*** | 0.5326*** | 0.4684*** | 0.3580** |
| Dec. 88 | 0.6077*** | 0.4769*** | 0.4644*** | 0.3082** |
| June 89 | 0.6226*** | 0.5240*** | 0.3959*** | 0.2971* |
| Dec. 89 | 0.4276*** | 0.6890*** | 0.4186*** | 0.5158*** |
| June 90 | 0.3582** | 0.5353*** | 0.1356 | 0.3703** |
| Dec. 90 | 0.2576* | 0.3882*** | 0.2486 | 0.2153 |
| June 91 | 0.3248** | 0.2530* | 0.1750 | 0.1871 |
| Dec. 91 | 0.2611* | 0.2547* | 0.1128 | 0.1718 |
| *N* | 43 | 44 | 44 | 43 |

***, **, * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively.

standard deviation of the residuals from the Market Model,[6] which captures the non-systematic, idiosyncratic risk of the firm's stock; and (iii) the ratio of market value of equities to book value of total assets, which measures the banking firm's capitalization. The two accounting measures of risk are (i) the ratio of book value equity to total assets and (ii) the ratio of loan charge-offs to total loans, which measure respectively the firm's book value capitalization and exposure to credit risk.[7] The moral hazard hypothesis predicts that inefficiency is positively related to the total risks and the idiosyncratic risk of stock returns, negatively related to capitalization, and positively related to loan charge-offs.

Panels A and B of Table 4 report the Pearson correlation coefficients between the estimated X-inefficiency and the five risk measures. Regarding stock returns, X-inefficiency is found to be positively correlated with both the total risks and the idiosyncratic risk of the banking firm's stock at the 1% significance level, regardless of firm size. On the association between inefficiency and capitalization, X-inefficiency is found to be negatively correlated with market value capitalization for firms in Quartiles 1, 2, and 3 at the 1 percent significance level and negatively correlated with book value capitalization for firms in all size classes at the 1 percent significance level. Finally, on the relation between inefficiency and credit risk, X-inefficiency is found to be positively correlated with loan charge-offs at the 1 percent significance level for firms in Quartiles 1, 2, and 3, and at the 5 percent significance level for firms in Quartile 4.

However, since the volatility of stock returns is positively related to capitalization, *ceteris paribus*, the bivariate relations between inefficiency and stock return volatility in panel A may be confounded by the effect of capitalization. To control for the leverage effect, standard deviations of daily stock returns are regressed against the inefficiency estimate and the ratio of market value equity to book value total assets. The OLS estimation results, reported in panel C of Table 4, indicate that even after controlling for the leverage effect, inefficiency has a significantly positive effect on stock return volatility. Similar results are obtained when the dependent variable is replaced by the standard deviation of the Market Model residual, reported in panel D of Table 4. The relations between inefficiency and risks embedded in stock returns seem robust.

6. In the Market Model, daily individual stock returns are regressed against the CRSP value-weighted market portfolio returns and an intercept term.

7. A caveat with respect to the ratio of loan charge-offs to total loans is that it also may capture managerial quality, which is correlated with inefficiency.

## TABLE 4

### Relations between X-Inefficiency and Firm Risk for 254 Bank Holding Companies from 1986 to 1991

Panel A: Pearson correlation coefficient between inefficiency and market measure of risk

| | Standard Deviation of Daily Stock Returns | Standard Deviation of Residuals from Market Model | Market Value Equity to Book Value Assets | N |
|---|---|---|---|---|
| Quartile 1 | 0.3605*** | 0.3637*** | –0.3333*** | 636 |
| Quartile 2 | 0.2906*** | 0.2961*** | –0.3636*** | 596 |
| Quartile 3 | 0.1786*** | 0.1791*** | –0.2589*** | 550 |
| Quartile 4 | 0.1493*** | 0.1462*** | –0.0676 | 554 |

Panel B: Pearson correlation coefficient between inefficiency and accounting measure of risk

| | Ratio of Loan Charge-offs to Total Loans | Book Value Equity to Asset Ratio | N |
|---|---|---|---|
| Quartile 1 | 0.5288*** | –0.5355*** | 774 |
| Quartile 2 | 0.4708*** | –0.3469*** | 657 |
| Quartile 3 | 0.3162*** | –0.3388*** | 643 |
| Quartile 4 | 0.0782** | –0.2531*** | 659 |

Panel C: OLS regression of standard deviation of stock returns on inefficiency and capitalization

| | Inefficiency | Market Value Equity to Total Assets | N |
|---|---|---|---|
| Quartile 1 | 0.058*** | –0.130*** | 636 |
| | (0.008) | (0.022) | |
| Quartile 2 | 0.026*** | –0.118*** | 596 |
| | (0.006) | (0.013) | |
| Quartile 3 | 0.013** | –0.107*** | 550 |
| | (0.006) | (0.012) | |
| Quartile 4 | 0.033*** | –0.125*** | 554 |
| | (0.010) | (0.013) | |

Panel D: OLS regression of standard deviation of Market Model residuals on inefficiency and capitalization

| | Inefficiency | Market Value Equity to Total Assets | N |
|---|---|---|---|
| Quartile 1 | 0.059*** | –0.130*** | 636 |
| | (0.008) | (0.022) | |
| Quartile 2 | 0.025*** | –0.117*** | 596 |
| | (0.006) | (0.013) | |
| Quartile 3 | 0.012** | –0.101*** | 550 |
| | (0.006) | (0.012) | |
| Quartile 4 | 0.026*** | –0.105*** | 554 |
| | (0.008) | (0.011) | |

***, ** indicate significance at the 1 percent and 5 percent levels, respectively. Standard errors are in parentheses.

Taken together, the findings provide strong evidence that X-inefficiency is associated with bank risk-taking and thus are consistent with the moral hazard hypothesis. Inefficient banking firms tend to have higher stock return variances, higher idiosyncratic risk in stock returns, lower capitalization, and higher loan losses. While the results in Table 4 reflect association, and not necessary causation, X-inefficiency seems to have important implications for risk management and bank safety, which should concern bank management as well as bank regulators.

## V. X-INEFFICIENCY AND STOCK MARKET VALUATION

This section further explores the relationship between X-inefficiency and bank stock returns. Previous research has shown that bank stock returns are sensitive to changes in interest rates, in addition to the market return, based on the two-index model (see, for example, Flannery and James (1984), Kane and Unal (1990), and Kwan (1991)). Both Flannery and James (1984) and Kwan (1991) also found that the sensitivity of bank stock returns to interest rate changes is related to the individual bank's assets and liabilities maturity profile, indicating that certain firm-specific factors have explanatory power for bank stock returns. In a similar spirit, it would be interesting to test whether another firm-specific factor, namely, operating efficiency, also provides explanatory power for bank stock returns.

To test the effect of operating efficiency on bank stock performance, the two-index model is modified to include the X-inefficiency estimate, in addition to the market return and changes in long-term interest rates.[8]

$$R_{jt} = \beta_0 + \beta_1 R_{mt} + \beta_2 R_{it} + \beta_3 \mathit{Inefficiency}_{jt} + \varepsilon_{jt} \tag{5}$$

where

$R_{jt}$ = return on firm $j$'s stocks for the semiannual period ending at time $t$,

$R_{mt}$ = return on the CRSP value-weighted market portfolio for the semiannual period ending at time $t$,

$R_{it}$ = relative change in 30-years constant maturity Treasury yield ($y$) from time $t$–1 to time $t$, i.e., $(y_t - y_{t-1})/y_{t-1}$,

$\mathit{Inefficiency}_{jt}$ = firm $j$'s estimated X-inefficiency for the semiannual period ending at time $t$, $\beta$'s are regression coefficients, and $\varepsilon_{jt}$ is the disturbance term.

Equation (5) is estimated by OLS using pooled time-series cross-section observations separately for each size class and the results are reported in Table 5. Consistent with prior studies, the coefficients of the CRSP market portfolio return are significantly positive and are close to unity. Moreover, the coefficients of the relative change in

8. Using short-term interest rates provides qualitatively similar results.

TABLE 5

OLS REGRESSION RESULTS OF BANK STOCK RETURNS ON THE CRSP MARKET RETURN, RELATIVE CHANGE IN THE LONG-TERM TREASURY YIELD, AND X-INEFFICIENCY

| | Coefficient Estimate | | | | |
|---|---|---|---|---|---|
| | Market Return | Treasury Yield Change | $\text{Inefficiency}_{jt}$ | $N$ | Adj. $R^2$ |
| Quartile 1 | 1.0233<br>(12.597)*** | –0.5684<br>(–5.115)*** | –0.3718<br>(–5.034)*** | 569 | 0.30 |
| Quartile 2 | 1.0706<br>(13.368)*** | –0.6259<br>(–5.672)*** | –0.4349<br>(–4.311)*** | 543 | 0.33 |
| Quartile 3 | 1.1278<br>(16.136)*** | –0.6608<br>(–7.024)*** | –0.1337<br>(–1.280) | 505 | 0.43 |
| Quartile 4 | 1.3554<br>(17.433)*** | –0.4728<br>(–4.437)*** | –0.3148<br>(–1.365) | 512 | 0.42 |

*** indicates significance at the 1 percent level; $t$-statistics are in parentheses.

long-term bond yield are significantly negative, indicating that increases in interest rates have a negative effect on bank stock returns. The level of firm-specific X-inefficiency is significantly negatively related to bank stock returns for firms in Quartiles 1 and 2, suggesting that inefficiency has a negative effect on stock returns. Although it has the expected negative sign, the coefficient of X-inefficiency is insignificant for the larger firm quartiles. However, the fact that the X-inefficiency is both smaller and has less cross-sectional variation among larger firms may make it more difficult to detect a statistically significant relationship between X-inefficiency and stock returns for these firms. On balance, inefficient banking firms seem to be associated with poor stock return performance, *ex post.*

## VI. Summary and Conclusion

Our findings provide further empirical evidence that substantial X-inefficiencies seem to exist in banking. In addition, several interesting properties of X-inefficiency are detected. After controlling for scale differences, smaller banking firms on average are found to be relatively less efficient than larger banking firms. Moreover, smaller banking firms tend to exhibit larger variations in X-inefficiencies than larger firms. While the findings suggest that the average large banking firm operates closer to its respective efficient frontier than the average small banking firm, the sources of these cross-sectional variations in X-inefficiencies can be answered only by future research.

Furthermore, the average X-inefficiency appears to decline over the period 1986 to mid-1990, apparently responding to the increased competition in banking wrought by market and regulatory changes. Although the average X-inefficiency seems to be falling, the rank orderings of firm-specific X-inefficiency are strongly correlated over time. The persistence of X-inefficiency rankings suggests that relatively efficient (inefficient) banking firms tend to stay relatively efficient (inefficient) over a fairly long period.

The persistence of firm-specific X-inefficiency leads us to investigate how the inefficient firms compensate for their inefficiency in the banking industry in order to avoid being driven out of the banking market. A strong correlation between firm-specific X-inefficiency and bank risk-taking is detected. Specifically, inefficient banking firms exhibit higher stock return variances, greater idiosyncratic risk in stock returns, lower capitalization, and higher loan charge-offs. The findings are consistent with the moral hazard hypothesis that inefficient banking firms may be able to extract larger deposit insurance subsidies from the FDIC to offset part of their operating inefficiencies. Hence, operating inefficiencies should concern not only bank management but also bank regulators.

Finally, for the smaller banking firms which exhibit large cross-sectional variations in X-inefficiencies, bank stock returns are found to be significantly negatively related to firm-specific X-inefficiency, after controlling for the market return and changes in risk-free interest rates. However, X-inefficiency appears to provide little explanatory power for the stock returns of larger banking firms, which tend to be more clustered together inside their respective efficient frontiers. The detection of a significant statistical relationship between X-inefficiency and *ex post* bank stock returns lays the groundwork for a more important research question: whether and how operating risk is priced in bank stocks.

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# Commodity Prices and Inflation

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*This study examines the empirical relationship between changes in commodity prices and inflation by looking at the performance of non-oil commodity prices as stand-alone indicators of inflation and in conjunction with other leading indicators of inflation. The results indicate that the empirical link between commodity prices and inflation has changed dramatically over time. Commodity prices were relatively strong and statistically robust leading indicators of overall inflation during the 1970s and early 1980s, but they have been poor stand-alone indicators of inflation since the early 1980s. When considered in conjunction with other likely indicators of inflation, non-oil commodity prices have had a somewhat more statistically robust relationship with inflation in recent years,though the added information content in commodity prices regarding inflation is limited.*

Commodity prices rose sharply from mid-1993 into 1995, more than 20 percent according to the Commodity Research Bureau index for all commodities. This burst in commodity prices raised concerns that overall inflation, which had been running at the lowest rate in years, would soon be on the rise. Despite the run-up in commodity prices, however, overall inflation remained relatively stable.

The role of commodity prices as precursors of inflation has been addressed extensively in the literature, with varying results. A long list of studies has shown that changes in the Commodity Research Bureau index and other commodity price indexes led aggregate inflation in the 1970s and the first part of the 1980s.[1] At the same time, studies by Garner (1995) and Bloomberg and Harris (1995) find that some commodity prices have not been reliable leading indicators of inflation since about the mid-1980s.[2]

This study examines the empirical relationship between changes in commodity prices and inflation by looking at the performance of non-oil commodity prices as stand-alone indicators of inflation and in conjunction with other leading indicators of inflation. The results indicate that the empirical link between commodity prices and inflation has changed dramatically over time, largely because of the changes in the extent to which movements in commodity prices reflect idiosyncratic shocks. Commodity prices were relatively strong and statistically robust leading indicators of overall inflation during the 1970s and early 1980s, a period dominated by relatively high inflation in commodity prices and in overall prices. However, commodity prices have been poor stand-alone indicators of inflation since the early 1980s, a period during which overall inflation has been relatively low and stable while commodity

1. For evidence on the short-run relationships see, for example, Cody and Mills (1991), Hafer (1983), Garner (1985), Defina (1988), Webb (1989), Furlong (1989), Kugler (1991), and Trivedi and Hall (1995).

2. Garner (1995) finds evidence of a decline in the statistical significance of several leading indicator variables in explaining inflation. For commodity prices, the study finds that lagged changes in commodity prices Granger cause inflation for the entire period 1973 to 1994, but were not significant for the period 1983 to 1994. Bloomberg and Harris (1995) look at samples split at 1987 and find similar results.

prices have been more volatile and generally declining relative to the overall price level.

When considered in conjunction with other likely indicators of inflation, non-oil commodity prices have had a somewhat more statistically robust relationship with inflation in recent years, though the added information content in commodity prices regarding inflation still is limited; for example, from 1993 through 1995 inflation may have been low relative to expectations, but shocks to commodity prices contribute little to explaining the puzzle.

The next section provides background for the statistical analysis by discussing the possible links between commodity prices and inflation and by illustrating the general patterns in the behavior of commodity prices and overall prices. Section II presents the empirical analysis relating to the bivariate relationship between commodity prices and overall prices. Section III presents the multivariate analysis results. Conclusions are presented in Section IV.

## I. LINKS BETWEEN COMMODITY PRICES AND INFLATION

Commodity prices are argued to be leading indicators of inflation through two basic channels. One is that they respond more quickly to general economic shocks, such as an increase in demand. The second is that some changes in commodity prices reflect idiosyncratic shocks, such as a flood that decimates the supply of certain agricultural products, which are subsequently passed through to overall prices. Depending on the type of the shock, the observed link between commodity prices and inflation would be expected to be different. Moreover, changes over time in the mix of shocks in the economy could affect the stability of a bivariate link between commodity prices and inflation.

The strongest case for commodity prices as indicators of future inflation is that they are quick to respond to economy-wide shocks to demand. Commodity prices generally are set in highly competitive auction markets and consequently tend to be more flexible than prices overall. As a result, movements in commodity prices would be expected to lead and be positively related to changes in aggregate price inflation in response to aggregate demand shocks.[3] In addition, to the extent that demand shocks are not sector-specific, the levels of commodity prices and overall prices also would be linked.

Any commodity, however, also is subject to idiosyncratic shocks. This complicates the empirical relation between commodity prices and inflation. In the case of a direct shock to the supply of a commodity, movements in the price of the commodity could be positively related to overall prices. The observed effect would depend on the relative importance of the commodity being shocked and the flexibility of other prices. Poor weather conditions, for example, could reduce the supply of agricultural commodities and push up their prices. The higher prices would eventually be reflected in the price of the related final food products bought by consumers. To the extent that the shock affects aggregate supply and that the stickiness in the prices of other consumer goods limits their adjustment, the net effect would be higher overall prices. The rise in the prices of the affected agricultural commodities would be larger than the effect on overall prices, which means the relationship of the level of prices of the affected commodities to overall prices would be affected.

One complication, however, is that a shift in relative demand for a commodity might dampen an otherwise positive correlation between the change in the price of a commodity and overall inflation. Take, for example, the case in which an increase in aggregate demand coincides with an increase in demand for manufactured goods or services relative to agricultural products. While this could lead to a rise in overall prices, prices of agricultural commodities might fall. In the short run, changes in commodity prices would not be positively related to inflation, and the levels of prices of the affected commodities and overall prices would drift apart.

These examples do not exhaust the possible permutations of shocks affecting commodity price and inflation; however, they do indicate that the relationship between the movements in commodity prices and inflation depends on what is driving commodity price changes. Given the alternative links between commodity prices and overall prices, two characteristics of empirical patterns are of interest. The first is whether commodity prices and overall prices are tied together in the long run. The second is the nature of the short-run relationship between changes in commodity prices and inflation.

### *Empirical Patterns*

As background for the more formal statistical analysis, this section gives a graphical overview of how commodity prices and overall prices have been related. The series used are the Commodity Research Bureau index for all commodities (CRB), its index for raw materials (CRBRAW), and the Consumer Price Index (CPI) as a measure of aggregate

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3. In fact, in theoretical models, such as in Boughton and Branson (1988), the relative flexibility of commodity prices results in their overshooting in order to bring markets into equilibrium in response to monetary shocks. Frankel (1986) also shows commodity prices can be expected to overshoot in response to monetary shocks.

prices.[4] These two commodity price indexes are examples of indexes that previous studies have found to be statistically significant leading indicators of aggregate inflation.

Figure 1 suggests a lack of a long-run relationship between the two commodity price indexes and the CPI. The figure plots the CRB and CRBRAW along with the CPI, each indexed to 100 in 1947. Over the period shown, the commodity series and the CPI drift apart. The drift is particularly pronounced starting in the 1980s. During the 1980s and 1990s, the CRB and CRBRAW indexes exhibit little if any trend, while the CPI continues to rise. The figure indicates, then, that over the past several years commodity price indexes have been influenced substantially by relative price movements.

To illustrate the short-run relationship between commodity prices and inflation, Figure 2 plots the 12-month percent changes in the CPI against the CRB and CRBRAW indexes. Peaks and troughs in commodity price inflation tend to precede turning points in CPI inflation. The pattern is the most regular in the 1970s and early 1980s. It appears that since the mid-1980s or so, the relation of CPI inflation to commodity price inflation has been looser. In the case of the CRB index, the 1987 peak in commodity price inflation preceded the next peak in CPI inflation by four years; this compares with an average lead of about nine months in the period prior to the mid-1980s. Moreover, commodity price inflation generally was rising from late 1991 on, but CPI inflation still had not picked up noticeably by late 1995.

## II. BIVARIATE VARs

To investigate the nature and consistency of the bivariate relationship between commodity prices and overall prices more formally, we use vector autoregression (VAR) models. The VARs include one of the Commodity Research Bureau commodity price indexes along with the CPI.

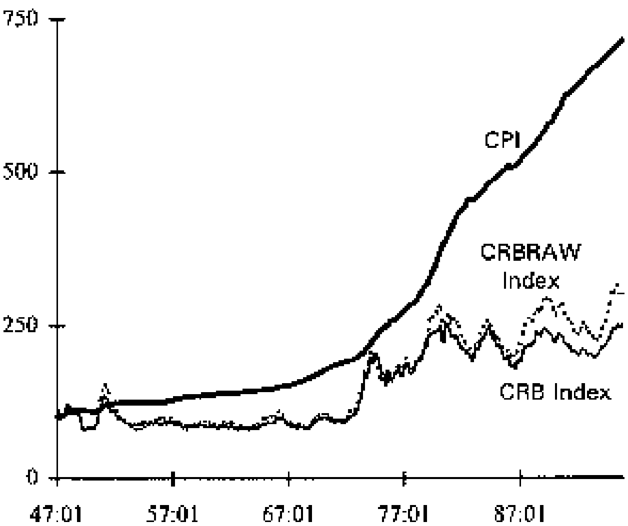
Integration tests on the log of the commodity price indexes indicate that these series have unit roots.[5] The integration tests involving the CPI are more problematic. Using the Augmented Dickey-Fuller (ADF) test for the change in the log CPI, we would reject the null hypothesis that the

4. The CRB index includes: cereals, meat, sugar, oils and seed oils, coca, cotton, rubber, hides, jutes, pint cloth, burlap, tallow, rosin, copper, iron ore, tin, zinc, lead. The CRBRAW index excluded the food and metals included in the CRB index. Neither index included energy or petroleum products.

5. Based on data for the period 1947:01 to 1995:12 and using an Augmented Dickey-Fuller test, we find that both CRB and CRBRAW are stationary in log-first differences.

FIGURE 1

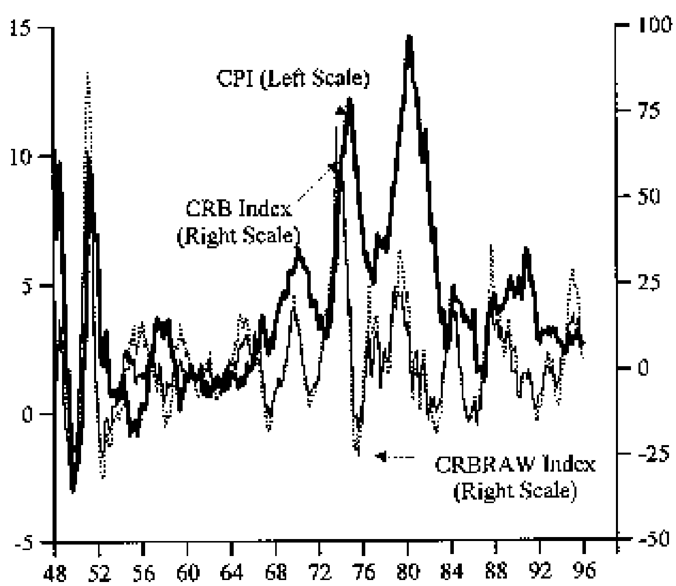
LONG-RUN LINK BETWEEN COMMODITY PRICES AND THE CPI




Note: Index, 1947:01=100

FIGURE 2

SHORT-RUN LINK BETWEEN COMMODITY PRICES AND THE CPI




Note: 12-month percent change

series is first-difference stationary. However, additional analysis suggests that first-differencing probably induces stationarity. For example, tests on the second difference indicate overdifferencing.[6] Therefore, we proceed assuming that the log-first difference of the CPI is a stationary series.

The lack of a long-term link between the two commodity price indexes and the CPI suggested by Figure 1 is confirmed by the results from ADF and Johansen tests for cointegration using monthly data for the logs of the three indexes for the period 1955:01 to 1995:12. The ADF tests indicate no bivariate cointegration. The results for the Johansen test vary depending on lag length used; cointegration is rejected at the 1 percent level for more than seven lags and at the 5 percent level for nine or more lags. Given these results, the analysis in this section assumes that the commodity price indexes and the CPI are not cointegrated.

To examine the short-run relationship between commodity prices and the CPI, then, we estimate bivariate VAR models with variables measured in log-first differences. The equations include 12 lags each for a commodity price index and the CPI. The commodity price indexes are ordered first, though the ordering has no effect on the conclusions regarding the relationship of commodity prices to overall inflation.[7]

### *Rolling Regressions*

We first look at the stability of the relationship between CPI inflation and changes in the commodity price indexes using a rolling regression approach. This involves identifying how the sum of the coefficients on the lagged commodity price index terms in the CPI equation vary as the sample length changes. Figure 3A shows the results for the lagged CRB terms when starting with the sample from 1960:01 to 1995:12 and dropping successive observations from the beginning of the sample. After dropping an observation, the equation is reestimated to get another value for the sum of the commodity coefficients and an $F$-statistic for the joint significance of the commodity index terms. The values plotted for a given date are the statistics estimated when the sample begins at that date.[8]

The results in the top left panel show that the sum of the coefficients on the lagged CRB terms begins to decline as the observations in the early 1970s are dropped from the sample. The decline continues through the early part of the 1980s. The results in the bottom left panel show relatively high marginal significance levels through the late 1970s. With the exception of a small spike, the marginal level of significance of the commodity terms is consistently under 5 percent (the dashed line) until the middle of 1979. After that, the marginal level of significance based on the $F$-test deteriorates: Though it goes below 5 percent for a brief period in the late 1980s, its value is generally in the 10 to 60 percent range. Hence, when the observations from the 1970s are removed from the sample, the commodity terms are no longer jointly statistically significant at conventional levels in the CPI equation.

The two right-hand panels present results from the reverse experiment. We begin with a relatively small number of observations from the beginning of the sample and show how the sum of the commodity coefficients and their level of significance change when the sample is extended. The figures plotted for a given date are the statistics estimated when the sample ends at that date. The top right panel shows that the sum of the coefficients for the lagged CRB terms increases and then drops sharply as observations for the 1960s are added. The sum of the coefficients rises through the first half of the 1970s, then dips, rebounds, and dips again in the early 1980s. The bottom right panel shows that the marginal level of significance based on the $F$-test improves when data for the early 1970s are included in the sample, but it falls below 5 percent only after observations for 1973 are included in the sample.

As Figure 3B shows, the results for the rolling regressions for the CPI equation when CRBRAW is included in the bivariate system are very similar to those for CRB. For the non-oil commodity indexes then, the empirical relationship with inflation is stronger and more robust for samples that include the 1970s. Moreover, a shift in the bivariate relationship appears to have occurred in the early to mid-1980s. Therefore, we consider below two subperiods, one from 1973 to 1983 and the second from 1984 to 1995.[9]

### *Subperiods*

Table 1 and the related Figures 4A and 4B show results from the bivariate VARs for the subperiods and serve as the basis for comparison with the results from the multivariate models presented in the next section. Table 1A reports the results for Granger causality tests for the CPI equations. As

6. Miller (1991) finds a similar result for the implicit price deflator.

7. All of the variance decompositions and impulse responses reported below are derived using a Choleski Factorization.

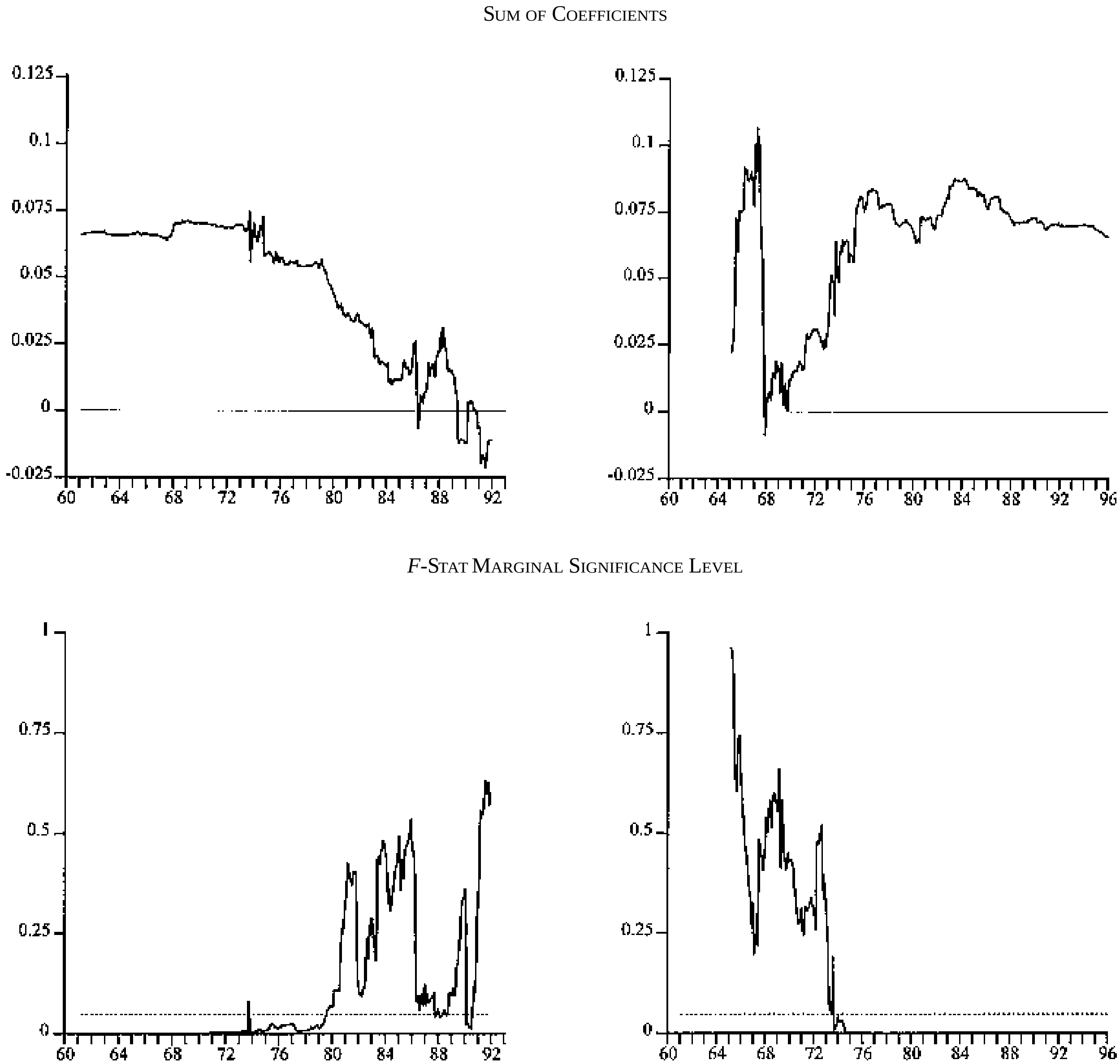
8. Note that because we are dropping observations as we move from left to right in the graph, there are fewer degrees of freedom in the denominator as we move to the right; the last significance level plotted in the graph is from an $F$- (12,24) test.

9. The first sample was started in 1973 to facilitate the inclusion of the foreign exchange value of dollar in the multivariate analysis presented in the next section. Also see Bryden and Carlson (1994).

FIGURE 3A

CRB IN CPI EQUATION: ROLLING REGRESSION RESULTS

| START DATE ROLLS, END DATE FIXED AT 95:11 | START DATE FIXED AT 61:02, END DATE ROLLS |
| --- | --- |



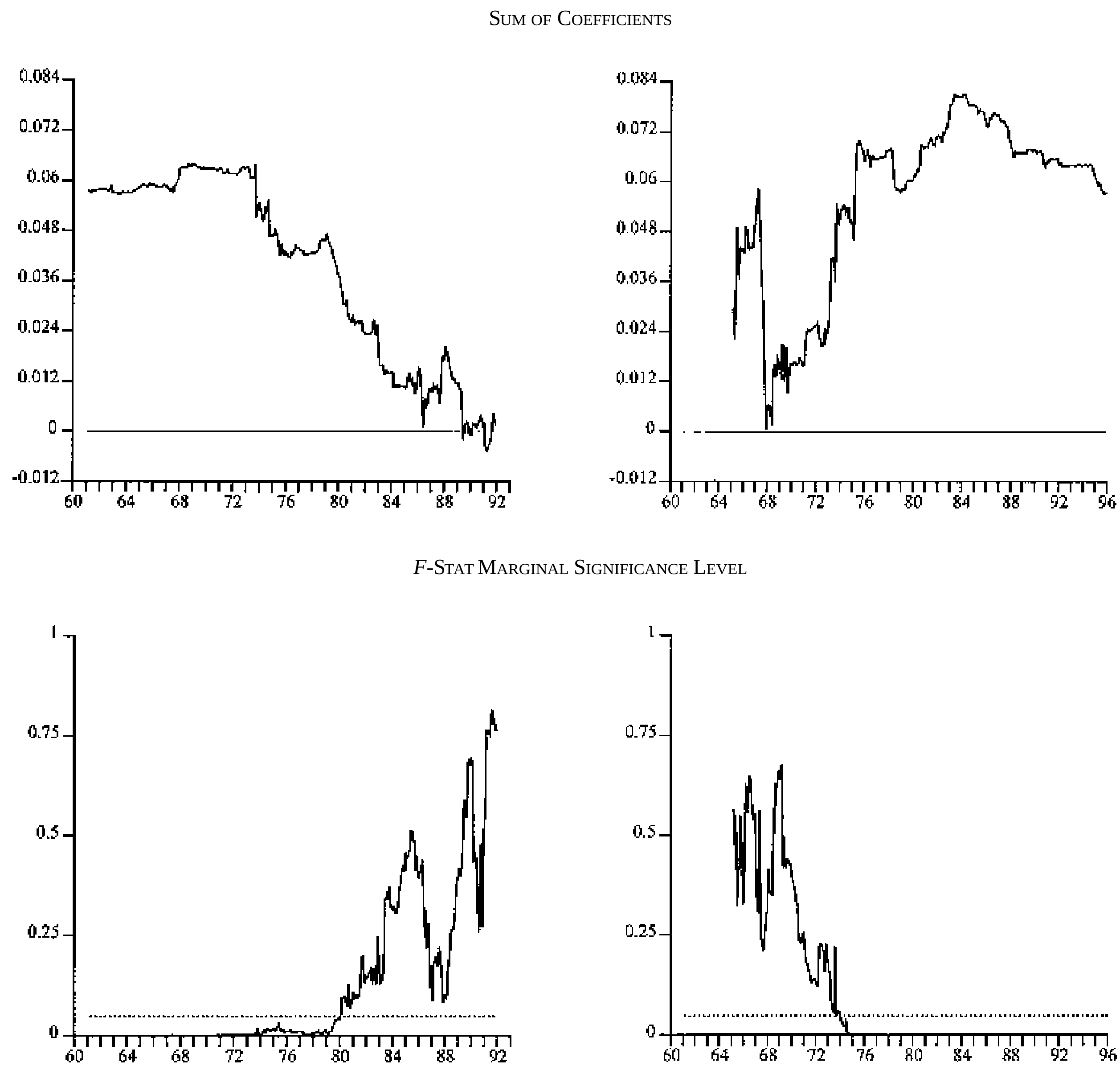

would be anticipated from Figure 3, the *F*-statistics indicate that the lagged coefficients for the commodity price terms are jointly significant only for the first period. Moreover, the overall explanatory power of the equation is almost three times larger for the first period than for the second. These results point to a change in the usefulness of commodity price indexes as stand-alone leading indicators of inflation.

Evidence on the variance decompositions also confirms a large difference in the relative importance of movements in commodity prices in explaining overall inflation in the two subperiods. Table 1B reports the forecast errors for CPI inflation over three horizons along with the share of that error accounted for by shocks to the commodity indexes and to the CPI. For the 24- and 36-month horizon, the share of the forecast error in CPI inflation attributed to the

FIGURE 3B

CRBRAW IN CPI EQUATION: ROLLING REGRESSION RESULTS

| START DATE ROLLS, END DATE FIXED AT 95:11 | START DATE FIXED AT 61:02, END DATE ROLLS |
|---|---|




commodity price indexes is more than three times larger in the first period than in the second period.

The differences in the variance decomposition results for the time periods reflect the combined effects of differences in the size and frequency of shocks as well as the magnitude of the response of the CPI to a given shock to the commodity price indexes. To some extent, the greater variance decomposition shares for the commodity price indexes for 1973–1983 may reflect the relatively greater volatility of commodity prices. Shocks to the commodity price indexes were 1.7 times greater in the first period than in the second, while the comparable figure for the CPI is 1.5 times. However, the results in Figure 3 suggest that the average response of the CPI to a given size shock to commodity prices also differs between the two periods.

The impulse responses for the CPI to shocks to the commodity price indexes for the bivariate VARs illustrate that this is the case. In Figures 4A and 4B, the responses are

## TABLE 1A

### CPI EQUATION RESULTS

| SPECIFICATION | CRB, CPI | | CRBRAW, CPI | |
|---|---|---|---|---|
| PERIOD | 1973–1983 | 1984–1995 | 1973–1983 | 1984–1995 |
| *F*-STAT: CRB (MSL) | 1.99<br>(.032) | 1.01<br>(.448) | 2.74<br>(.003) | 1.15<br>(.326) |
| *F*-STAT: CPI (MSL) | 6.08<br>(.000) | 3.41<br>(.000) | 6.66<br>(.000) | 3.29<br>(.000) |
| $R^2$ | .44 | .16 | .47 | .17 |

## TABLE 1B

### VARIANCE DECOMPOSITIONS: TWO-VARIABLE VAR

| | 1973–1983 | | | 1984–1995 | | |
|---|---|---|---|---|---|---|
| FORECAST HORIZON | STANDARD ERROR | SHARE OF ERROR DUE TO: | | STANDARD ERROR | SHARE OF ERROR DUE TO: | |
| | | CRB | CPI | | CRB | CPI |
| 12 Months | .00296 | 26.4 | 73.6 | .00183 | 11.8 | 88.2 |
| 24 Months | .00325 | 37.1 | 62.9 | .00184 | 11.7 | 88.3 |
| 36 Months | .00332 | 36.5 | 63.5 | .00184 | 11.7 | 88.3 |
| | | CRBRAW | CPI | | CRBRAW | CPI |
| 12 Months | .00305 | 32.0 | 68.0 | .00184 | 14.1 | 85.9 |
| 24 Months | .00335 | 42.1 | 57.9 | .00185 | 14.3 | 85.7 |
| 36 Months | .00339 | 41.9 | 58.1 | .00185 | 14.3 | 85.7 |

derived using the same size shock for each of the time periods. The shocks are the average shocks to the log changes in the CRB or the CRBRAW indexes over the entire sample period, which equal about .018. The figures show the average responses and the upper and lower two standard deviation bands (the bands are four standard deviations wide).[10] The upper panels of Figures 4A and 4B show the response of CPI inflation. The bottom panels show the implied response (and error bands) for the log level of CPI. Multiplying the response in the lower panels by 100 gives the compounded percent change in prices for each forecast horizon.

The figures indicate that the responses of overall prices to the shocks to commodity prices changed significantly between the two periods. For shocks to CRB and CRBRAW, the response of CPI inflation in the first period is two standard errors above the zero axis out to about the two-year horizon. With a few exceptions, the response of inflation in the post-1983 period is not significantly above zero.

The bottom panels of Figures 4A and 4B provide a perspective on the relative size of the cumulative response, with the increase in CPI being eight times greater in the first period than in the second. The results in the bottom panels also indicate that the more pronounced responses

10. Bands for the impulse responses were calculated from the results of 1,000 impulse responses, with each response generated using a covariance matrix of residuals altered by a random draw from a standard normal distribution. We then computed the variance from the first and second moments and set the band width equal to two standard errors above and below the average response.

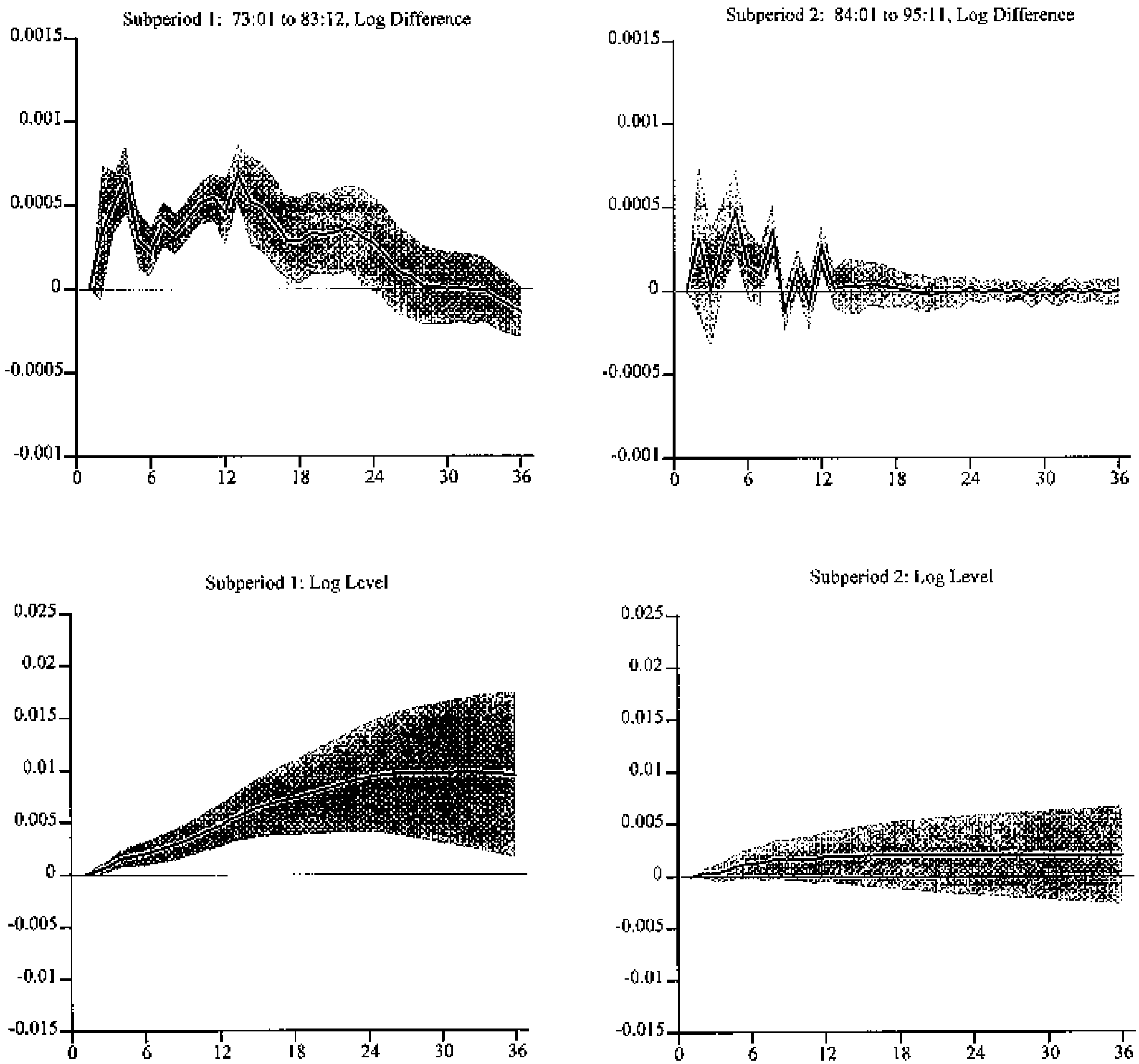
in the first subperiod are significantly greater than those estimated for the more recent years. The average responses of CPI for the 1973–1983 period are above the upper two-standard deviation bands for the 1984–1995 period. The average responses for the second period, in turn, are below the lower bands for the 1973–1983 period.

## *Out-of-Sample Forecasts*

The implications of the change in the relationship between the commodity price indexes and inflation in the bivariate models can be illustrated more concretely by relating it to the recent behavior of prices. Out-of-sample forecasts

FIGURE 4A

IMPULSE RESPONSE OF CPI TO A SHOCK TO CRB: TWO-VARIABLE VAR



were derived using the CPI equation from the bivariate VAR that included CRB and a univariate equation for CPI inflation with 12 lagged terms. The equations were estimated for the period 1973:01 to 1993:12. Dynamic simulations were used to derive the forecast for CPI inflation for 1994 and 1995, with the changes in commodity prices equal to the actual values. The forecasts for CPI inflation are translated into log levels of the CPI.

Figure 5 shows the forecasted series for the log CPI along with the corresponding actual series. The baseline is the forecasted series obtained from the univariate CPI equation. In the figure, the baseline overestimates the CPI by

FIGURE 4B

IMPULSE RESPONSE OF CPI TO A SHOCK TO CRBRAW: TWO-VARIABLE VAR

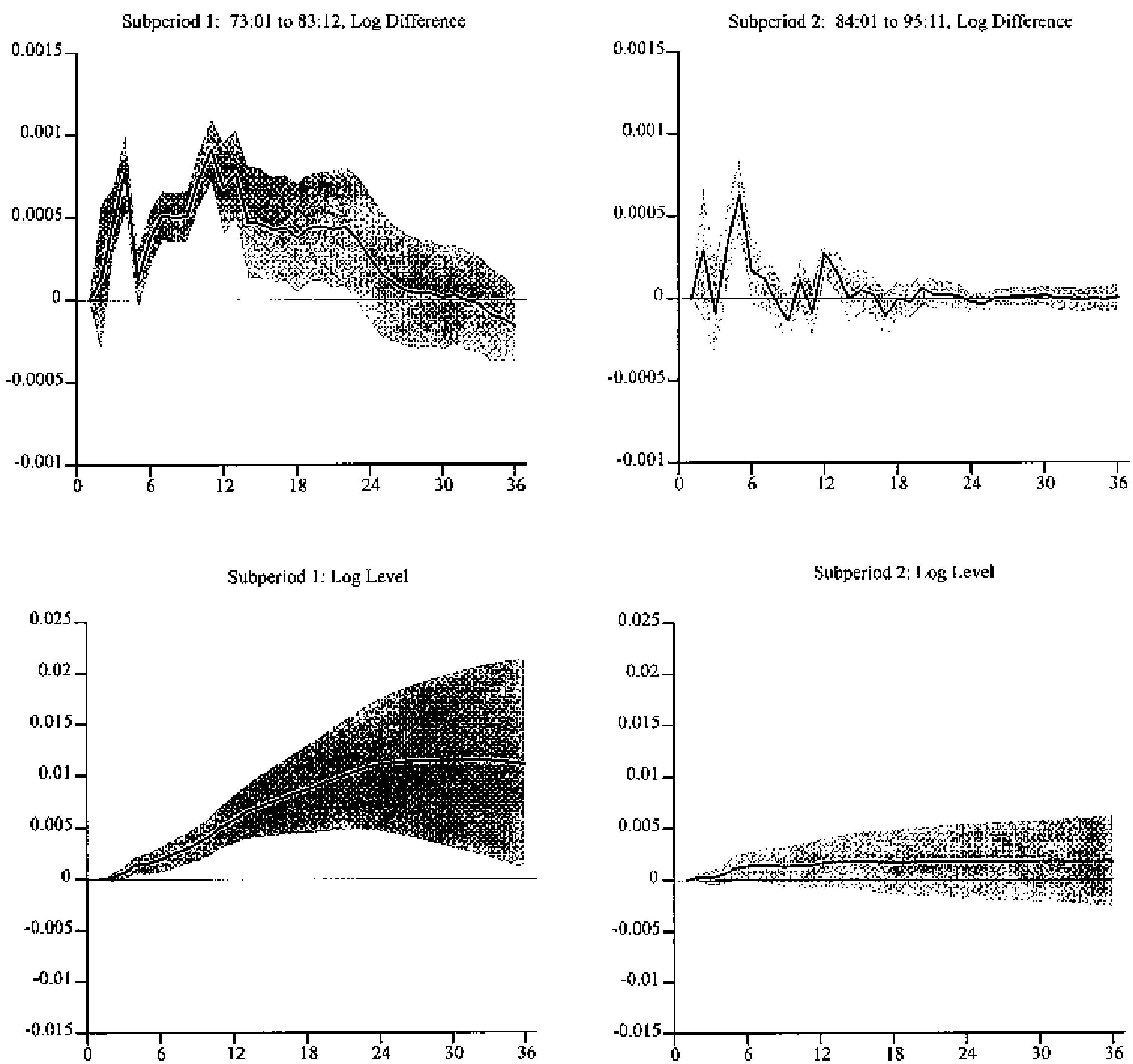
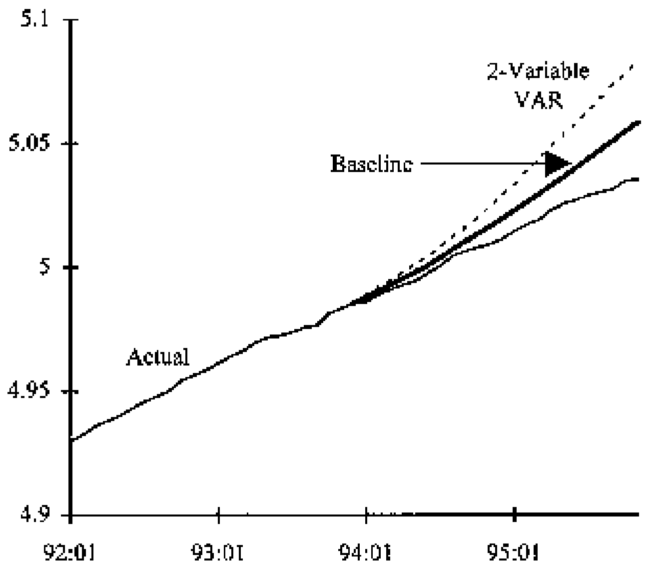


FIGURE 5

FORECASTS OF THE LOG CPI




about $2^{1}/_{2}$ percent over the two years. Using that as the benchmark, the forecasted series from the equation that includes the commodity price index adds another $3^{1}/_{2}$ percent to the estimate for the CPI.[11] This says that based on the historical relationship between commodity prices and inflation, inflation should have picked up noticeably in 1994 and 1995. However, as indicated earlier, CPI inflation did not.

## III. MULTIVARIATE VARS

The results in the previous section indicate that the commodity price indexes have not been reliable stand-alone indicators of inflation. It still is possible that commodity price indexes can provide some unique and reliable information about overall price movements if considered in conjunction with other leading indicators of inflation. The inclusion of additional variables also can help to sort out whether the shift in the bivariate relationship between changes in commodity prices and inflation was due to differences in the extent to which commodity prices conveyed more general economic shocks versus idiosyncratic ones. To investigate these issues, we consider other possible leading indicators of inflation, along with the commodity price indexes.

One of the additional variables is a measure of the tightness in labor markets, namely, the difference between the actual unemployment rate and the Congressional Budget Office estimate of full employment unemployment (NUR). This is meant to capture the notion that a tight labor market tends to be associated with shocks that lead to upward pressure on inflation and a slack labor market with ones that lead to disinflation.

Two other variables are the spot price of oil (OIL) and a multilateral trade-weighted exchange value of the U.S. dollar (FX). The price of oil is considered because oil shocks are widely viewed as having temporary effects on the rate of inflation in the U.S. Also, since CRB and CRBRAW do not include petroleum, the oil price augments them with a potentially important commodity. The foreign exchange rate is included since currency markets are highly liquid and prices can adjust quickly in response to changes in information that has a bearing on future inflation.

Finally, the analysis includes the federal funds (FF) interest rate as one indicator of monetary policy. This allows for the possibility that a shift in the response of monetary policy to movements in commodity prices has affected the simple bivariate relationship between changes in commodity prices and inflation.

Integration tests for the additional variables indicate that they are stationary in log-first differences, with the exception of NUR, which is stationary in levels.[12] For comparison with the bivariate benchmarks in the previous section, the multivariate VARs are estimated in levels for NUR, and log first-differences for the other variables.[13] The ordering

11. The forecasts from the two-variable VAR are more than two standard deviations above the actual CPI from mid-1994 to the end of 1995. We estimated a two-variable system (in difference of logarithms) containing 12 lags each of a commodity term (CRB and CRBRAW, respectively) and inflation. The estimation range was 1973:01 to 1993:12. We then did a dynamic forecast of inflation from 1994:01 to 1995:12 assuming we knew the value of the commodity with certainty. We calculated the standard error bands around the forecast by bootstrapping the residuals from the estimated inflation equation and feeding them back into the forecasting exercise as shocks. We bootstrapped the residuals 1,000 times, generating 1,000 fore-

19. Central Valley banks tended to report relatively high problem loan ratios for most of the period from 1985 until 1989, a period when this

12. In the VAR models used in the analysis, own shocks to NUR dissipate to zero over time.

13. Some studies have found evidence of cointegration for commodity prices, inflation, and other variables. Marquis and Cunningham (1990) find evidence that industrial production, commodity prices, and aggregate prices are cointegrated using data from 1968 to 1986. The results in Kugler (1991) suggest that commodity prices, CPI, and the dollar exchange rate might be cointegrated.

In this study, we also tested for cointegration using the levels of the commodity indexes, CPI, OIL, FX , and FF. The test results suggest cointegration over the period 1973 to 1995 using 12 lags. However, the results from our analysis indicate that, when a common cointegrating vector is used for the subperiods, the results are very similar with and without the cointegrating vector.

of the variables is NUR, FX, OIL, CRB or CRBRAW, CPI, and FF.

### *Effects of Changes in Commodity Prices*

Table 2 and Figure 6 present the results from the estimations of the multivariate VARs. A comparison of the results with those from the bivariate VARs points to two characteristics of the additional variables. One is that the variables contain added information about inflation that is not contained in either the commodity price indexes or lagged inflation. This is indicated by the higher overall explanatory power of the estimated CPI equations and the smaller forecast errors in Table 2 compared to the results in Table 1. The variables contributing the added information, however, differ between the two periods. Separate estimates of CPI equations with lagged values of CPI, a commodity price index, and one of the other variables indicate that NUR, FX, OIL, and FF each add something to the higher overall explanatory power and smaller forecast errors in the multivariate VARs in the first period. In the second period, however, the oil price is the main source of added explanatory power.

The second characteristic is that the added variables contain information about inflation that was attributed to the commodity price indexes in the bivariate VARs. That is, their inclusion reduces the amount of independent information associated with the commodity prices. This is primarily true for the 1973–1983 period. For that subperiod, the marginal levels of significance of the coefficients on the commodity terms are raised appreciably compared to the bivariate cases.

The variance decomposition shares for the commodity indexes in the CPI equation in Tables 2B and 2C also show less of a relative role for commodity price indexes, with the bigger change evident for the first subperiod. The other major differences in the variance decomposition shares are in the roles of the oil price and the CPI. Consistent with the above mentioned differences in the added information content of the variables in the two periods, the table shows that changes in the price of oil account for a small share of the forecast error in CPI inflation in the first period and a larger share in the second period. The variance decompositions also show that a higher share of the forecast error for CPI is attributed to itself in the second period. These results suggest that sources of short-run variation in inflation were different for the two periods.

The effect of including the additional macroeconomic variables is most striking in the responses of CPI inflation to shocks to the commodity price indexes. A comparison of the lower-left panels in Figures 6A and 6B with those in Figures 3A and 3B indicates that the average responses of inflation to the same size shock to the commodity price indexes are much less in the multivariate cases than in the

TABLE 2A

CPI EQUATION RESULTS
SIX-VARIABLE VAR: NUR, FX, OIL, CRB/CRBRAW, CPI, AND FF

| SPECIFICATION | CRB | | CRBRAW | |
|---|---|---|---|---|
| PERIOD | 1973–1983 | 1984–1995 | 1973–1983 | 1984–1995 |
| *F*-STAT: NUR (MSL) | 1.68<br>(.096) | 1.11<br>(.367) | 1.80<br>(.068) | 1.30<br>(.237) |
| *F*-STAT: FX (MSL) | 1.96<br>(.045) | 1.36<br>(.206) | 2.05<br>(.035) | 1.82<br>(.062) |
| *F*-STAT: OIL(MSL) | 1.16<br>(.333) | 4.08<br>(.000) | 0.80<br>(.652) | 4.04<br>(.000) |
| *F*-STAT: CRB/CRBRAW (MSL) | 0.38<br>(.967) | 0.79<br>(.656) | 0.75<br>(.698) | 1.09<br>(.384) |
| *F*-STAT: CPI (MSL) | 1.70<br>(0.89) | 3.46<br>(.000) | 2.12<br>(0.28) | 3.48<br>(.000) |
| *F*-STAT: FF (MSL) | 1.92<br>(.051) | 1.13<br>(.350) | 1.76<br>(.076) | 1.29<br>(.246) |
| $R^2$ | .61 | .42 | .63 | .44 |

## TABLE 2B

### Variance Decompositions: NUR, FX, OIL, CRB, CPI, FF; 1973–1983

| Forecast Horizon | Standard Error | Share of Forecast Errors for CPI due to: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRB | CPI | FF |
| 12 Months | .00223 | 14.2 | 11.2 | 8.3 | 8.8 | 44.7 | 12.8 |
| 24 Months | .00270 | 13.8 | 16.3 | 8.9 | 11.6 | 32.9 | 16.5 |
| 36 Months | .00306 | 12.4 | 13.9 | 7.9 | 20.3 | 28.2 | 17.3 |

## TABLE 2C

### Variance Decompositions: NUR, FX, OIL, CRB, CPI, FF; 1984–1995

| Forecast Horizon | Standard Error | Share of Forecast Errors for CPI due to: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRB | CPI | FF |
| 12 Months | .00162 | 10.2 | 5.9 | 22.9 | 7.6 | 40.7 | 12.7 |
| 24 Months | .00175 | 11.2 | 6.6 | 23.0 | 8.1 | 38.4 | 12.7 |
| 36 Months | .00179 | 11.3 | 6.7 | 23.1 | 7.9 | 38.4 | 12.6 |

## TABLE 2D

### Variance Decompositions: NUR, FX, OIL, CRBRAW, CPI, FF; 1973–1983

| Forecast Horizon | Standard Error | Share of Forecast Errors for CPI due to: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRBRAW | CPI | FF |
| 12 Months | .00225 | 13.5 | 15.4 | 4.9 | 12.7 | 42.3 | 11.2 |
| 24 Months | .00274 | 13.9 | 23.6 | 6.3 | 12.6 | 31.2 | 12.4 |
| 36 Months | .00304 | 12.0 | 21.4 | 6.2 | 19.4 | 28.1 | 12.9 |

## TABLE 2E

### Variance Decompositions: NUR, FX, OIL, CRBRAW, CPI, FF; 1984–1995

| Forecast Horizon | Standard Error | Share of Forecast Errors for CPI due to: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRBRAW | CPI | FF |
| 12 Months | .00162 | 10.6 | 8.7 | 21.1 | 6.2 | 40.7 | 12.7 |
| 24 Months | .00176 | 11.5 | 8.2 | 21.5 | 6.9 | 37.9 | 14.0 |
| 36 Months | .00181 | 12.2 | 8.4 | 21.4 | 6.8 | 37.4 | 13.8 |

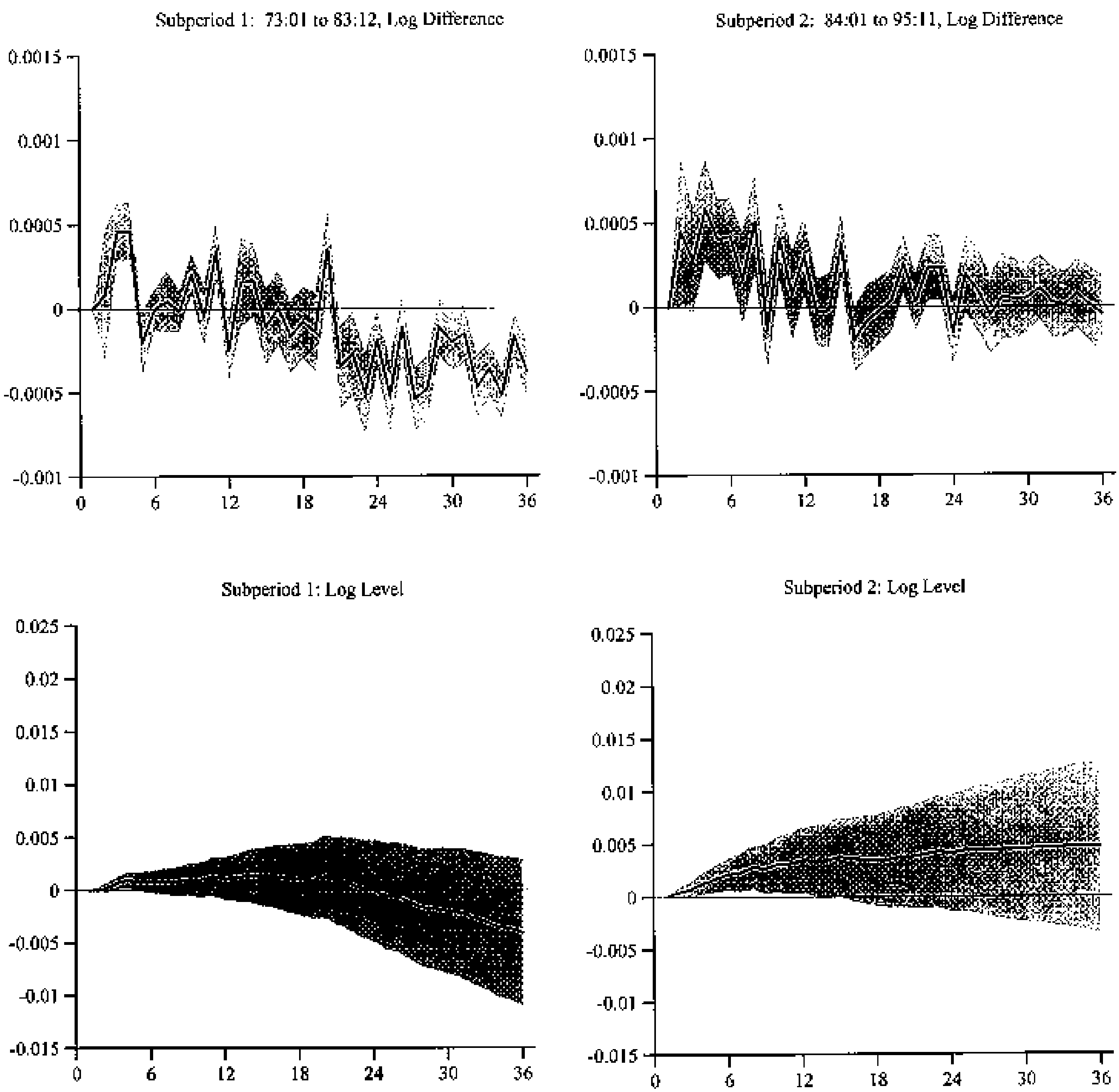
bivariate cases for the first period.[14] For that period, the average responses of the CPI to a shock to CRB or CRBRAW in the multivariate VARs generally are below the lower two standard deviation bands for the responses in the bivariate models. Moreover, the responses from the multivariate VAR are not statistically significant beyond the very near-term horizons in the first period.

The results relating to the effects of the commodity price indexes are not sensitive to their ordering in the VAR models.

14. The size of the shocks to the commodity indexes are the same as those used in the bivariate analysis.

FIGURE 6A

Impulse Response of CPI to a Shock to CRB: Six-Variable VAR

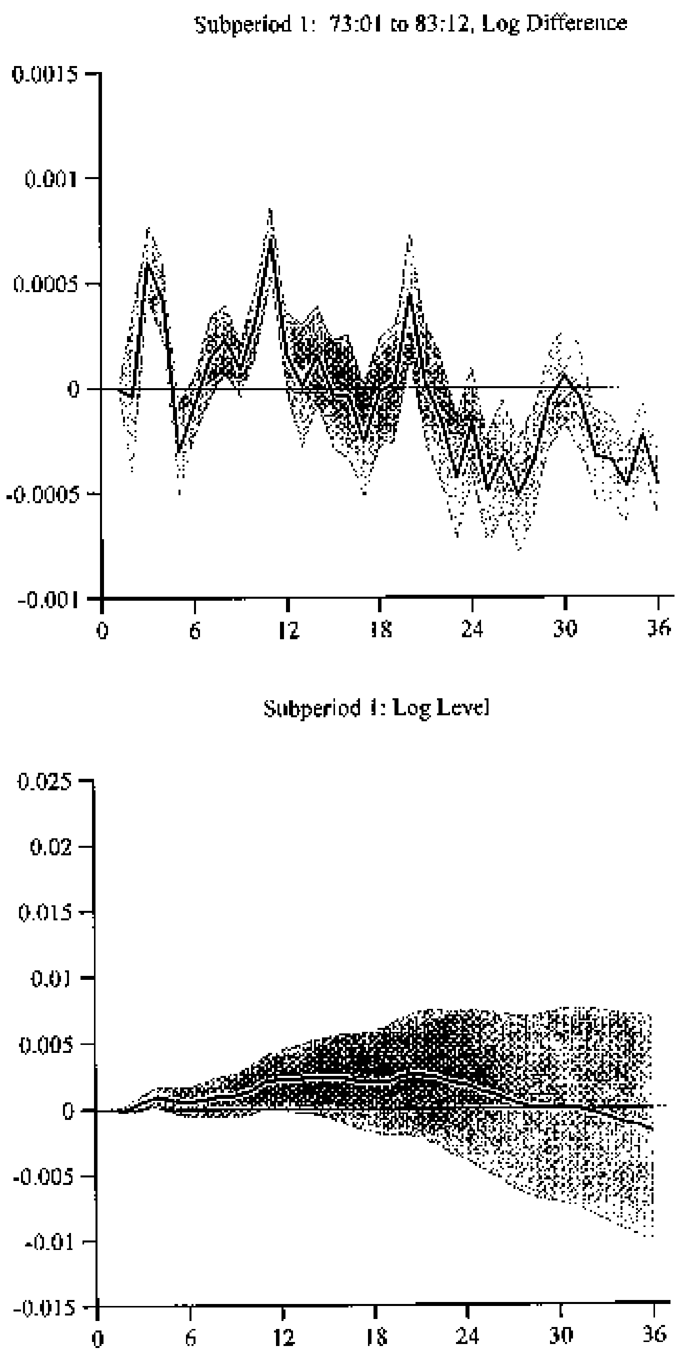


The variance decomposition shares and the responses of CPI when the indexes are ordered first are almost identical to those shown in the tables and figures. This suggests that commodity prices in the first period were responding to the shocks affecting employment and the foreign exchange value of the dollar. In other words, commodity prices likely were signaling more general economic shocks affecting inflation in the first period and not just idiosyncratic shocks.

A comparison of the results for the second period presents a different picture. The lower-right panels in Figures 6A and 6B and those in Figures 3A and 3B show that the average response of the CPI to a shock to CRB or CRBRAW

FIGURE 6B

IMPULSE RESPONSE OF CPI TO A SHOCK TO CRBRAW: SIX-VARIABLE VAR



in the second period is very similar in the multivariate and bivariate cases. Comparing the multivariate results across the periods (the lower two panels in Figures 6A and 6B) shows that the average response is actually a bit larger in the second period, though the difference is not significant. However, taken by itself, the average response of CPI to a CRB shock in the lower-right panel of Figure 6A is more than two standard deviations above zero for a horizon of more than a year. The comparable horizon for CRBRAW shocks in Figure 6B is about nine months. Again, the results relating to the commodity price indexes are not sensitive to their ordering in the VAR models.

The results, then, indicate that the information content in shocks to commodity prices about future inflation in the second period did not overlap significantly with other macroeconomic variables. The shocks to commodity prices that conveyed information about future inflation in the second period were more idiosyncratic than in the first period. These differences in the information content of shocks to commodity prices in the two periods could account for the results in the previous section showing the commodity indexes were relatively robust stand-alone indicators in the 1970s and early 1980s, but not in more recent years. This could be because idiosyncratic shocks to commodity prices (those associated with a positive response in overall prices) tend to affect the relevant commodity prices more than overall prices, while more general economic shocks could have a more balanced long-run impact on commodity prices and overall prices. This appears to have been the case in the sample period covering the 1970s and early 1980s. For that period, shocks to the commodity price indexes led to larger responses of those indexes relative to the responses of CPI than did shocks to either NUR or FX.

Pinning down the reasons for the difference in the information content of the commodity price indexes, however, is problematic. One explanation is that the mix of shocks changed.[15] It is possible, for example, that general economic shocks were more important relative to idiosyncratic commodity price shocks in the first period compared with the second. That is, while supply shocks may have had some role, commodity prices performed relatively well as stand-alone inflation indicators in the first period because the relatively high inflation rates ultimately reflected persistent aggregate demand shocks. Such a shift in the relative importance of shocks would be consistent with the relatively stable and low CPI inflation, the general decline in the relative price of commodities, and the relatively greater role of oil price shocks in explaining CPI inflation in the multivariate systems for the subperiod since the early 1980s.

A second explanation is that the changes in commodity price indexes have become less effective in conveying shocks generally. One possibility is that the role of commodities has changed. Bloomberg and Harris (1995), for example, point out that the role of commodities in total output has declined over time. In the case of supply shocks to a commodity or basket of commodities, this should mean that over time a given change in commodity prices would have a smaller impact on overall prices. This may have played some role in the change in the empirical relationship between commodity prices and inflation. However, it seems unlikely that it could account for an eightfold difference in the response of inflation to a shock to commodity prices in the first period compared with the second period as is found in the bivariate models. Moreover, the analysis above suggests that the response of CPI inflation to idiosyncratic shocks to the commodity is more likely to have been larger than smaller in recent years.

Another possibility relates to commodity prices signaling aggregate demand shocks. The argument is that over time commodities have been used less for hedging against inflation because of the availability of alternative financial instruments (Bloomberg and Harris). If so, this could reduce the demand for some durable commodities and contribute to the drift in the level of their prices relative to other prices. The implications for the short-run link are less straightforward, however. Prices of durable commodities still should respond to aggregate demand shocks, though possibly with less overshooting—that is, smaller initial commodity price movements for a given shock. However, if that were the case (and the ultimate response of overall inflation to a given aggregate demand shock were the same) we should find evidence of larger, not smaller, responses of inflation to shocks to commodity prices in the bivariate VARs in recent years compared to the more distant past.

A third general explanation is that the response of monetary policy to shocks to commodity prices has changed. The idea is that, if monetary policy were to respond to shocks to commodity prices to head off inflation, the observed relationship between commodity prices and inflation would be changed. Since the observation is that the link between commodity prices and inflation is weaker, the argument would have to be that monetary policy has responded more in recent years to offset the pending inflation. This raises two issues: Has monetary policy responded more to commodity prices, and, if so, how much has it affected the empirical relationship between the commodity price indexes and inflation?

15. Another explanation is that the relationship between CPI inflation and the other macroeconomic variables has been unstable.

Evidence regarding the monetary policy response can be gleaned from the federal funds rate equations in the multivariate VARs discussed above. These turn out to be inconclusive on whether the federal funds rate has been more responsive to changes in commodity prices in recent years. In Table 3 the results for the FF equations in the multivariate VARs indicate that the lagged values of the CRB terms are jointly significant in the second period but not the first period, and those for the CRBRAW terms are not jointly significant for either period. The variance decompositions for the forecast errors of FF differ noticeably for some variables in the two periods. The shares for NUR are smaller in the second period, while those for CPI are larger, suggesting monetary policy may have responded relatively more to movements in inflation. The share for the price of oil also is larger in the second period. However, the shares for non-oil commodity price indexes in the two periods are virtually unchanged. Finally in Figures 7A and 7B while the average responses of FF to shocks to commodity prices are larger in the second period than in the first, the difference generally is not statistically significant at the 5 percent level.

Even if monetary policy has responded more to commodity prices in recent years, the results in Figures 4 and 6 show that the inclusion of the federal funds interest rate did not fundamentally change the response of inflation to a given commodity price shock in the second period. That finding suggests that any difference in response of monetary policy to commodity prices is not the fundamental cause of their decline in usefulness as stand-alone indicators of inflation.

## *Recent inflation behavior*

While the commodity price indexes may be poor stand-alone indicators of inflation, the results from the multivariate analysis indicate that they still may provide some information about future in inflation when considered in conjunction with other inflation indicators. That is, a shock to commodity prices still might be expected to have a positive, though modest, impact on inflation. In this section we look at whether the net rise in commodity prices from mid-1993 to mid-1995 provided any additional information about inflation.

To investigate this, the multivariate VAR that included the CRB was estimated for the period 1984 to 1993. Two forecasts were derived for the CPI for the period 1994–1995. The baseline forecast sets NUR, FX, OIL, and FF at their historical values and allows both CPI inflation and changes in the CRB to be forecasted dynamically. The second forecasts CPI inflation dynamically while setting the values for all the other variables to their historic values.

The implied forecasts for log level of the CPI along with the actual CPI are shown in Figure 8. The two forecasted series overpredict actual CPI by a large margin. By the end of 1995, the forecasts are about 5 percent above actual, and that spread is greater than two standard errors of the forecasts. This suggests that, at least relative to the model used

TABLE 3A

FF-Equation Results
Six-Variable VAR: NUR, FX, OIL, CRB/CRBRAW, CPI, and FF

| Specification | CRB | | CRBRAW | |
|---|---|---|---|---|
| Period | 1973–1983 | 1984–1995 | 1973–1983 | 1984–1995 |
| *F*-Stat: NUR (MSL) | 1.43<br>(.177) | 0.85<br>(.598) | 1.47<br>(.160) | 0.51<br>(.904) |
| *F*-Stat: FX (MSL) | 0.55<br>(.871) | 1.03<br>(.433) | 0.52<br>(.895) | 0.88<br>(.567) |
| *F*-Stat: OIL(MSL) | 0.63<br>(.803) | 1.32<br>(.226) | 0.72<br>(.727) | 1.36<br>(.204) |
| *F*-Stat: CRB/CRBRAW (MSL) | 0.77<br>(.678) | 1.91<br>(.048) | 1.91<br>(.669) | 0.95<br>(.507) |
| *F*-Stat: CPI (MSL) | 1.45<br>(.170) | 1.18<br>(.317) | 1.15<br>(.337) | 1.06<br>(.408) |
| *F*-Stat: FF (MSL) | 2.83<br>(0.04) | 1.08<br>(.393) | 2.29<br>(.018) | 0.81<br>(.638) |
| $R^2$ | .29 | .33 | .29 | .23 |

## TABLE 3B

### VARIANCE DECOMPOSITIONS: NUR, FX, OIL, CRB, CPI, FF; 1973–1983

| FORECAST HORIZON | STANDARD ERROR | SHARE OF FORECAST ERRORS FOR FF DUE TO: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRB | CPI | FF |
| 12 Months | .87153 | 20.1 | 5.8 | 5.2 | 7.7 | 2.3 | 58.9 |
| 24 Months | 1.02324 | 17.5 | 9.3 | 5.7 | 10.8 | 4.1 | 52.6 |
| 36 Months | 1.05454 | 17.2 | 9.6 | 6.6 | 10.9 | 4.3 | 51.4 |

## TABLE 3C

### VARIANCE DECOMPOSITIONS: NUR, FX, OIL, CRB, CPI, FF; 1984–1995

| FORECAST HORIZON | STANDARD ERROR | SHARE OF FORECAST ERRORS FOR FF DUE TO: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRB | CPI | FF |
| 12 Months | .25939 | 11.0 | 6.3 | 13.4 | 11.8 | 15.1 | 42.4 |
| 24 Months | .28163 | 11.2 | 7.7 | 12.6 | 10.6 | 19.3 | 38.6 |
| 36 Months | .28967 | 11.4 | 7.8 | 12.3 | 10.5 | 20.8 | 37.2 |

## TABLE 3D

### VARIANCE DECOMPOSITIONS: NUR, FX, OIL, CRBRAW, CPI, FF; 1973–1983

| FORECAST HORIZON | STANDARD ERROR | SHARE OF FORECAST ERRORS FOR FF DUE TO: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRBRAW | CPI | FF |
| 12 Months | .88695 | 20.8 | 7.5 | 5.3 | 7.6 | 2.7 | 56.1 |
| 24 Months | 1.02991 | 18.0 | 10.1 | 5.4 | 11.2 | 4.4 | 50.9 |
| 36 Months | 1.05712 | 17.9 | 10.1 | 5.8 | 11.8 | 4.6 | 49.8 |

## TABLE 3E

### VARIANCE DECOMPOSITIONS: NUR, FX, OIL, CRBRAW, CPI, FF; 1984–1995

| FORECAST HORIZON | STANDARD ERROR | SHARE OF FORECAST ERRORS FOR FF DUE TO: | | | | | |
|---|---|---|---|---|---|---|---|
| | | NUR | FX | OIL | CRBRAW | CPI | FF |
| 12 Months | .26673 | 9.2 | 5.1 | 11.5 | 12.0 | 16.2 | 46.0 |
| 24 Months | .28299 | 9.5 | 5.7 | 11.7 | 11.4 | 18.2 | 43.5 |
| 36 Months | .28849 | 10.0 | 5.9 | 11.4 | 11.3 | 18.8 | 42.6 |

FIGURE 7A

IMPULSE RESPONSE OF THE FEDERAL FUNDS RATE TO A SHOCK TO CRB: SIX-VARIABLE VAR

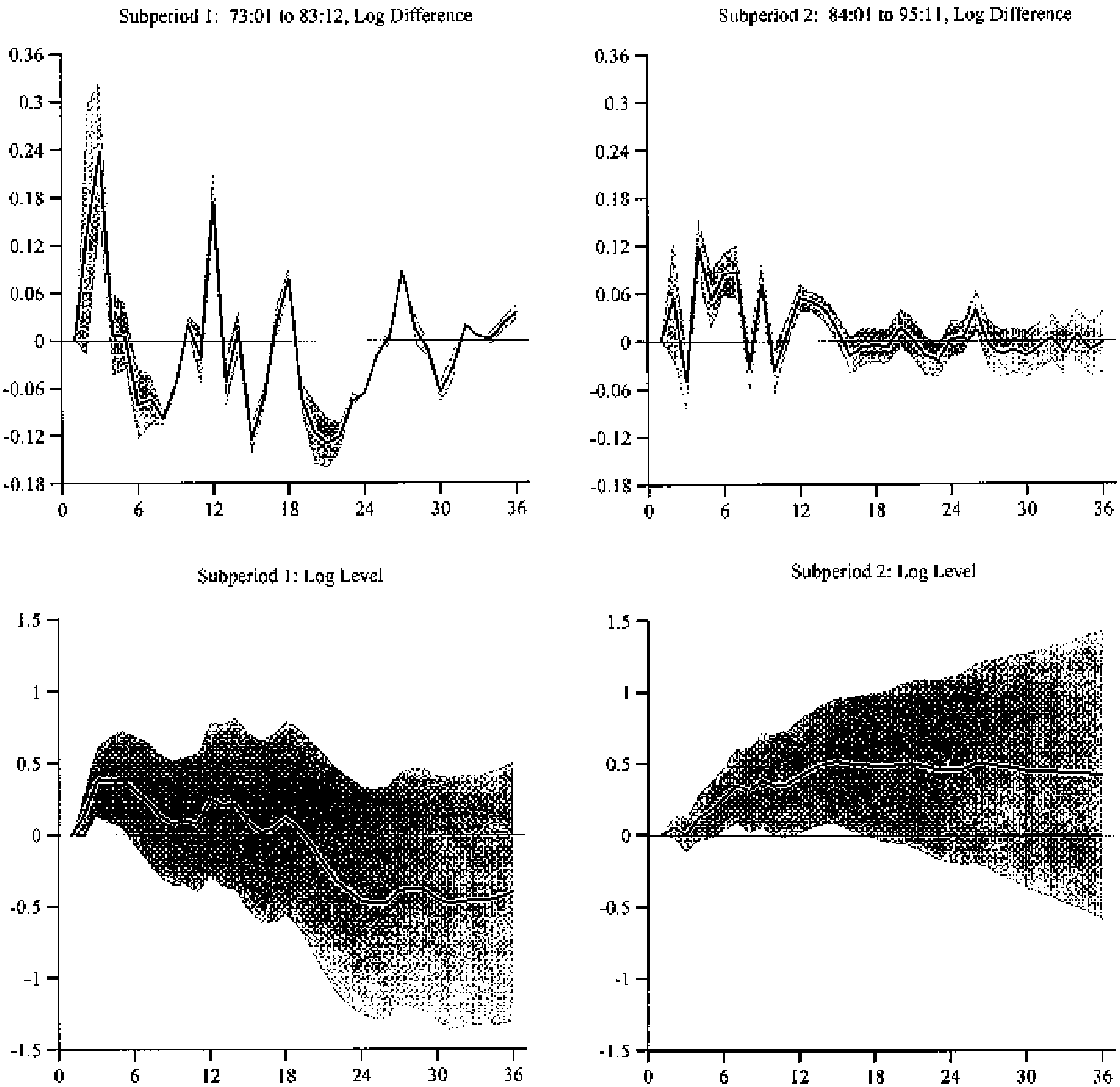


FIGURE 7B

IMPULSE RESPONSE OF THE FEDERAL FUNDS RATE TO A SHOCK TO CRBRAW: SIX-VARIABLE VAR

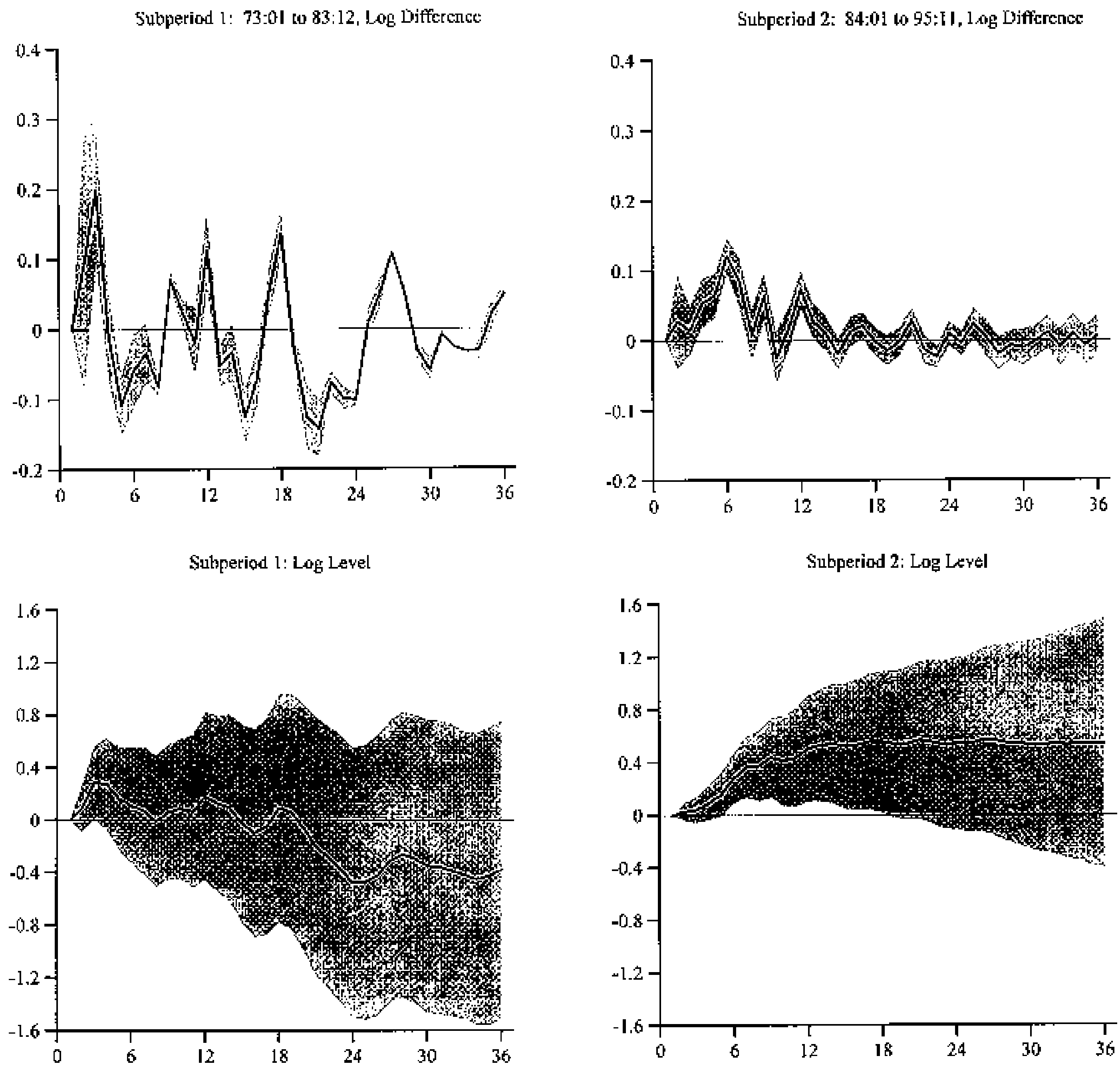
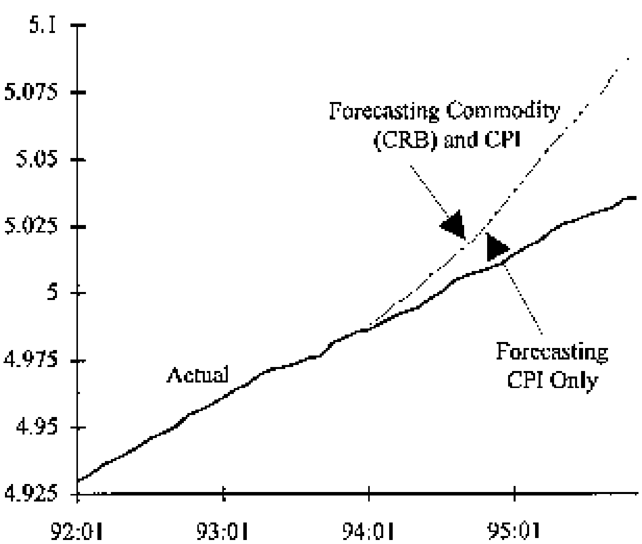


FIGURE 8

COMPARISON OF FORECASTS OF LOG CPI:
SIX-VARIABLE VAR




in this study, inflation was lower than expected in the 1994–1995 period. The similarity of the two forecast series, however, indicates that shocks to commodity price indexes were relatively small and do not help explain the overprediction of inflation.

## IV. CONCLUSION

The simple two-way relationship between CPI inflation and the commodity price indexes has changed significantly over time. The non-oil commodity prices were relatively strong and statistically robust leading indicators of overall inflation for a period covering the 1970s and early 1980s, but they have performed poorly in more recent years. As a result, using the past relationship between commodity prices and inflation to forecast inflation leads to a sizeable overprediction of inflation in recent years.

The deterioration in the role of non-oil commodity prices as stand-alone indicators of inflation appears to reflect a change in the extent to which the movement in the prices of these commodities reflected general economic shocks ultimately affecting overall inflation versus more idiosyncratic shocks to commodities. We find the non-oil commodity indexes performed relatively well as stand-alone indicators of inflation when the commodity prices conveyed the effects of factors affecting inflation that were reflected first in the tightness in labor markets and the foreign exchange rate of the dollar, while they performed poorly when they did not.

Pinpointing the reasons for the difference in the information content of commodity prices is problematic. Explanations such as the decline in the commodities' share in overall output, less use of commodities for inflation hedging, or offsetting response of monetary policy appear inadequate to account for the deterioration in empirical relationships between changes in commodity prices and overall inflation. Another possibility suggested in our analysis is a change in the mix of shocks affecting prices. Such a change occurring would be consistent with the relatively stable and low CPI inflation, the general decline in the relative price of commodities, and the more important role of oil price shocks in explaining inflation since the early 1980s.

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