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The 1985 Ohio Thrift Crisis, the FSLIC's Solvency, and Rate Contagion for Retail CDs

ELIZABETH S. COOPERMAN, WINSON B. LEE, and GLENN A. WOLFE*

ABSTRACT

This paper uses both an ARIMA transfer-function intervention model and a panel data analysis to examine the effect of the Ohio deposit insurance crisis in 1985 on the pricing of six-month retail certificates of deposit (CDs) for federally-insured Ohio banks and savings and loans. Adjusting for pricing reactions due to changes in market rates, we find a significant, unanticipated rise in CD-rate premiums on the initial event week of the crisis that continued for approximately seven weeks. Consistent with a contingent insurance guarantee hypothesis, rate premiums are found to be risk based.

FROM BOTH A REGULATOR'S and a bank manager's perspective, it is important to understand the nature of depositors' reactions to a depository institution crisis. Recent events, including large real estate loan losses for major banks, the Rhode Island Deposit Insurance crisis and the General Accounting Office's (GAO) announcement that a large bank failure or a severe recession could jeopardize the solvency of the Federal Deposit Insurance Corporation (FDIC), have renewed debate over the potential for a widespread loss of confidence in the banking system. In particular, concern has been expressed over the potential for instability in retail deposit markets in a deregulated environment (see Smith (1984)).[1]

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[1] A recent article on FDIC bailout legislation in the *New York Times* (November 11, 1991, p. D1) exemplifies this type of concern, whereby without new loans for the FDIC, "blameless depositors would not be reimbursed when their banks fail next year, and that could lead to a run on banks like the ones in the Depression that led the Government to begin insuring bank deposits."

Although previous studies have found little evidence of significant market reactions surrounding large bank failures (see Benston et al. (1986)), these studies have generally focused on stock market reactions and were made when the FDIC was adequately capitalized. The Federal Savings and Loan Insurance Corporation (FSLIC) was decapitalized in the mid-1980s; however, an examination of deposit market reactions to a savings and loan (S & L) crisis has never been made.[2] Kane (1989) notes that when insurance funds and depository institutions are decapitalized, depositor confidence may waver in terms of the efficacy of deposit insurance, and silent runs, or the threat of silent runs, may occur for less solvent banks/S & Ls. If a small percentage of depositors threaten to withdraw funds, poorly capitalized banks and S & Ls may be able to retain funds by paying a higher deposit rate, which makes "customers willing to live with their growing doubts" (Kane (1989), p. 12). If higher rates are sustained, depository institutions may suffer losses, however, and regulatory intervention may be needed.

Kane (1989) suggests that silent runs occurred during the 1985 Ohio deposit-insurance crisis, a period when regulators refused to intervene to bail out the Ohio Deposit Insurance Guaranty Fund thrifts.[3] In the mid-1980s, the FSLIC was economically insolvent, which Kane points out regulators attempted to hide.[4] With considerable publicity during the Ohio crisis concerning the inadequacy of FSLIC reserves, however, depositors should have become more cognizant of potential insurance fund problems. Whether or not small, less sophisticated depositors were aware of and cared about the FSLIC's problems as early as 1985, is an unresolved issue.

The purpose of this paper is to examine the effect of the Ohio Deposit Insurance crisis on the pricing of retail (insured) six-month certificates of deposit (CDs) for a sample of federally-insured Ohio banks and S & Ls. We use an Autoregressive Integrated Moving Average (ARIMA) transfer function intervention model which filters out changes in market (T-bill) rates, and allows an examination of the timing, strength, and length of an event's impact on weekly CD rate premiums. We examine a contingent insurance guarantee hypothesis, suggested by Kane (1989), that predicts risk-based pricing during the crisis period. Consistent with this hypothesis, our results

[2] Cook and Spellman (1991) did examine the impact of the GAO's announcement of the FSLIC's insolvency in 1987 on S & L large retail deposit prices. They found a CD rate jump of 120 bp that lasted during 16 months.

[3] Although there is no record of a deposit run on a federally-insured thrift in connection with the state insurance crises in 1985, FSLIC thrift withdrawals exceeded new deposits in seven of the nine months following the Ohio crisis.

[4] Kane (1989, p. 9) states that if accumulated and projected losses for the FSLIC were incorporated in the GAO's $4.6 billion estimate in 1985, reserves would be negative. In 1985 the Federal Home Loan Bank Board first publicly acknowledged that the FSLIC's reserves were significantly below the amount needed to close the nation's regulatorily insolvent thrifts (see Brumbaugh 1988, p. 33). In April, the bank board created a management consignment program to allow large, severely insolvent S & Ls to continue to operate under new, regulator-selected management to protect declining FSLIC reserves. The GAO's announcement of the FSLIC's bankruptcy did not occur until March of 1987.

indicate a significant, unexpected rise in CD prices for less solvent Ohio depository institutions, lasting for approximately seven weeks. The results from an alternative panel data analysis confirm a significant rise in CD rates, and the existence of risk-based pricing during the crisis period.

The paper proceeds as follows. Section I discusses the events of the Ohio crisis and briefly summarizes previous empirical and theoretical studies on deposit market contagion. Section II presents the data and methodology, followed by the empirical results in Section III. Section IV provides a summary and conclusion.

I. The Ohio Crisis and Previous Studies

A. Events

The events of the crisis are shown in Table I.[5] ESM Government Securities, a Fort Lauderdale government securities dealer, was closed on March 4, 1985, resulting in losses of $145 million on repurchase agreements for Home State Savings Bank of Cincinnati, Ohio. These losses exceeded the value of Home State's capital and the $130 million reserves of Home State's deposit insurer, the Ohio Deposit Guaranty Fund (ODGF), a privately-owned, state-chartered fund. ODGF offered full-insurance on deposits, and its funds came from member S & L's deposits, reported as assets on member's balance sheets. Consequently, the ODGF's depletion inflicted losses on all member S & Ls. Runs occurred at Home State Savings during the week of ESM's failure (net withdrawals of $154 million), and on twelve other ODGF S & Ls (net withdrawals of $101 million for the seven most affected Cincinnati firms) the following week (see Cleveland Federal Reserve Annual Report (1985)).

The event week (Week 0) for the intervention analysis begins on March 4, the date of the first public announcement of a potential crisis. During Weeks 0 through 1, a nonintervention policy was followed by state and federal regulators, with the exception of liquidity provided to ODGF thrifts through the Cleveland Federal Reserve's discount window. Bottlenecks occurred in releasing funds, however, since thrifts were not familiar with the correct procedures and principles on which loans could be made, and runs on several ODGF S & Ls continued (see Meltzer (1986) and McCulloch (1987)). By state law, the Ohio legislature was not allowed to allocate funds to private agencies, and federal regulators did not wish to reduce FSLIC standards to allow ODGF thrifts to convert to federal insurance.[6]

[5] For a more detailed discussion of the events of the crisis see Kane (1987), (1989) and McCulloch (1987).

[6] We are grateful to Walker F. Todd at the Cleveland Federal Reserve for pointing out the problems that regulators faced during the Ohio crisis. Federal regulators were hampered by the potential inflationary effects associated with providing assistance to the ODGF thrifts and by feasibility restrictions in providing sufficient liquidity to ODGF firms. At this time the FHLBB insisted upon newly instated, more stringent requirements for ODGF thrift conversions to FSLIC insurance (Kane, 1987).

Table I

Chronological List of Events in the Ohio S & L Crisis

Dates (1985)	Event
Week 0	
March 4	ESM Government Securities fails, and is found to have a shortfall of $350 million
March 6–8	In response to potential losses, deposit runs occur at Home State of more than $150 million
March 9–10	Home State is closed, and a conservator is appointed. Loss to ODGF fund undermines solvency of other ODGF-insured thrifts.
Week 1	
March 11	Runs occur on Ohio ODGF-insured thrifts. Fed assists S & Ls to obtain documents to borrow from Fed's discount window.
March 13	Ohio legislature authorizes a new ODGF fund based on higher premiums and a special loan. Home State's losses are predicted to wipe out this fund. Heavy deposit runs at four ODGF S & Ls.
March 14	Massive withdrawals occur for six ODGF thrifts. ODGF thrift executives seeking FSLIC insurance are told that ODGF S & Ls must meet new, more stringent requirements for conversion, and the process could take months.
March 15	After all-night meetings with regulators, Celeste invokes emergency powers to close 71 ODGF thrifts. State refuses option of putting full-faith and credit behind the thrifts.
Week 2	
March 18	The dollar declines in foreign exchange markets 2–3%. The Ohio state legislature passes a bill requiring ODGF S & Ls to apply for and receive federal insurance before they can reopen.
March 19	FHLBB agrees to ease state thrift requirements for FSLIC Insurance and to speed applications.
March 20	Dollar begins to recover. State legislature passes bill allowing openings with possibility of limited customer withdrawals and authorizing indemnification of FSLIC for any loss incurred in ODGF S & Ls through July 1, 1987. Fed discount assistance is republicized.
March 22	Bank Board begins approving federal insurance for some of the ODGF thrifts.
Week 3	
March 26	18 S & Ls are fully open with only Home State and a few other thrifts fully closed.

Primary source *Wall Street Journal Index* and cited articles, 1985 and Kane, 1987, pp. 323–324, and 1988, Table 4).

On March 13 in Week 1, state legislators enacted a special bill establishing a new ODGF to be funded by additional levies on member institutions and an emergency loan. This fund, however, proved to be inadequate relative to ODGF losses. At the end of Week 1, in response to continued runs, the ODGF holiday was declared. At this time state and federal regulators refused to provide additional aid in what Kane (1987) notes to be a standoff between the state and federal governments in response to the private insurance fund's crisis. Kane (1987) suggests that regulatory hesitations in bailing out the privately-insured ODGF institutions may have provided a signal that federal regulators might similarly fail to bail out other federally-insured thrifts in the event of a major failure, which may have increased depositors' risk perceptions.[7]

The moratorium on Ohio S & Ls was followed by a strong reaction in national and international markets. On March 18, the beginning of Week 2, the dollar fell 2–3%, and the price of gold and silver respectively increased 3 and 5%.[8] In Week 2, in response to these market reactions, a change in policy occurred, with additional regulatory assistance provided by both governments. The state agreed to indemnify ODGF losses through July 1, 1987 for S & Ls that converted to federal insurance. In turn, the FSLIC relaxed conversion standards. By Week 3, ODGF thrifts were allowed to reopen on a full or partial basis. They received federal insurance or were merged with federally-insured firms by the end of 1985.

B. Previous Studies

Because of Regulation Q and previous data limitations, the pricing effects of a crisis on retail (insured) deposits has never been studied. A few studies, however, have examined the effect of major bank failures in the 1980s on the prices of uninsured CDs. Cramer and Rogowski (1985) found that premiums on negotiable CDs increased 63 basis points after the announcement of Penn Square's failure. Baer and Brewer (1986) and Kutner (1988), however, did not find any short or long-term effects from Penn Square's closure on the general CD market, with the exception of an increase in the CD rate premiums of Continental Illinois. Wall and Peterson (1989) observed a clear increase in banks' CD spreads during the Continental Illinois crisis in May 1984, with spreads increasing to higher levels in late May and in late June, coinciding with Latin American loan problems.

[7] McCulloch (1987, p. 247) notes that the lengthy delays for ODGF depositors to receive insured deposits foreshadowed future delays for FSLIC depositors, since any recapitalization bill entails a considerable time period to be drafted and pushed through legislative channels.

[8] On Monday, March 18, at the Treasury's auction, the average rate on 13-week bills also rose 16 basis points from the previous auction, and 26-week bills, 25 basis points. These reactions were perceived in the financial press as a "flight to quality," by foreign and domestic investors and depositors. Benston et al. (1986) note that when depositors withdraw funds on news that may be related to their bank, they may alternatively: (1) redeposit funds at a bank perceived to be safer, which would have minimum social costs; (2) transfer funds to safer securities, such as Treasury securities, which may raise risk premiums on bank deposits; or (3) transfer funds to a currency or a safer currency, which would have a stronger detrimental economic effect on other banks.

Although the reaction of retail CD prices to bank failures has not been examined, several studies have found evidence of market discipline by retail depositors in deregulated markets. Cook and Spellman (1991) and Cooperman, Lee, and Sinkey (1990) found S & L CD rates to be a function of insolvency risk proxies during 1987, the year the General Accounting Office (GAO) announced the FSLIC's technical insolvency.[9] Van Drunen and Wikstrom (1989) also found evidence of risk-based premia for the highest yielding bank and thrift CDs in the nation during 1986–1987, and Hirschhorn (1989), for a broad sample of S & Ls in 1988. Other studies by Short and Gunther (1988) and Golding, Hannan, and Liang (1989) found rate premiums to be higher for banks and S & Ls in regions suffering economic distress. While Short and Gunther (1988) observed premia in Texas to be related to a firm's insolvency risk (net worth) in 1988, Golding et al. (1989) examining different states during an earlier time period, did not.

In summary, there is some evidence of risk-based pricing for large retail (insured) deposits during the period after the 1987 announcement by the GAO that the FSLIC was technically insolvent, while there is mixed evidence concerning the existence of risk-based pricing for negotiable (uninsured CDs) for banks during periods of large bank failures. Cook and Spellman (1989) observed that depositors in 1987 appeared to price S & L deposits based on the conjectural value of the FSLIC's guarantee, since there was no evidence of risk-based pricing on a sample date in 1986, or for banks in their sample. In contrast to previous studies, we examine the price reactions of small retail deposits during a S & L crisis period, and examine the issue of whether or not small, less sophisticated depositors were aware of, and cared about the FSLIC's problems in 1985, two years prior to the GAO's announcement, in response to the Ohio S & L crisis.

C. The Contingent Insurance Guarantee Hypothesis

The rationale for many bank and thrift regulations, including the development of deposit insurance, is based on the potential for financial market instability if public confidence is reduced in the banking system as a whole. This instability, in turn, is thought to create social costs in terms of negative externalities including potential deposit runs, a greater risk premium demanded by depositors of other solvent depository institutions, and a potential disruption of economic activity (see Meltzer (1986), Tallman (1988), and Paroush (1988)). Recent theories diverge in terms of the potential causes of bank market instability and predictions regarding depositors' reactions.[10] In this paper, we focus on the potential for bank market instability during the Ohio crisis, as reflected in the risk premia of retail CDs.

[9] Cook and Spellman (1991) examined surveyed bank and S & L CD rates for large ($100,000 denomination) insured deposits on two dates, respectively in 1986 and 1987. Cooperman et al. (1990) examined CD rates for the nation's largest banks and thrifts in seven major metropolitan areas in 1985 and 1987.

[10] See Tallman (1988) for a survey of recent theories of bank contagion and their predictions.

Kane (1989) attests that depositors' expectations are based on the adequacy of insurance fund resources and the government's willingness to intervene in a crisis, i.e., if a private insurance crisis, such as the ODGF crisis, is mishandled, depositors' rational expectations that federal insurance funds will similarly be mishandled will increase. Similarly, uncertainty may be created if the government fails to explicitly guarantee federal deposit insurance funds. Barth and Bradley (1988) point out that in the event that insurance funds are wiped out, there is no explicit guarantee that the government will come to the insurance agencies' rescue. This, in turn, can create a loss of confidence and potential spillover effects. Kane notes that silent runs are more likely to occur during periods when insurance funds and depository institutions are decapitalized. To assuage wary depositors and keep them from removing funds, less solvent depositories may consequently offer higher deposit rates. We refer to this approach as the contingent insurance guarantee hypothesis.

Adverse publicity concerning declining FSLIC reserves during the Ohio crisis may have made depositors more aware of the FSLIC's impending insolvency in 1987. If depositors price deposits on the value of conjectural deposit insurance guarantees, i.e., the financial condition of the insurer (see Cook and Spellman (1991)), deposit premia would be expected to increase. Golding, Hannan, and Liang (1989) suggest that risk premia are also demanded for expected inconvenience costs associated with thrift closure including liquidity, reinvestment risk, and lost accrued interest after a firm's closing date. Moreover, depositors would be expected to look at an individual bank's capital as a cushion against the potential for regulatory closure when federal deposit insurance funds are weak.

In effect, the failure of one guarantor (the ODGF) may serve as a wake-up call to depositors regarding the condition of federal insurance funds and, consequently, may affect the perceived risk and subsequent pricing on the CD issues by banks and thrifts that are implicitly backed by the federal government. The contingent insurance guarantee hypothesis predicts that risk-based CD rate premia will appear during the crisis for federally-insured depository institutions, based on a depository institution's insolvency risk. In contrast, if depositors have full confidence in implicit federal insurance guarantees and the expediency of failure resolutions, no new risk premiums should appear during the crisis period.[11]

II. Data and Methodology

A. Data

Weekly retail, minimum deposit rate data were obtained for a sample of federally-insured Ohio banks and S & Ls through a survey conducted in June

[11] We would like to thank Lewis Spellman, our discussant at the 1991 AFA meeting, for helping to clarify these points.

1988, which was mailed to the 350 thrifts and banks throughout Ohio that were in operation in 1985 and whose names and addresses continued to be listed in the Federal Home Loan Bank of Cincinnati's Membership and Services Directory for 1987, representing two-thirds of the depository institutions in the state. Usable data were received for 35 banks and 27 S & Ls across Ohio (see Appendix A). Our survey was performed two years after the crisis, so a potential survival bias exists, i.e., our sample does not include depository institutions that failed between 1986–1988. Since such firms would be more likely to be judged by the market as risky, our results may understate any observed risk premia.

Data on two additional banks and five additional S & Ls in Cleveland obtained from the Bank Rate Monitor were also included, resulting in a total of 37 Ohio banks and 32 S & Ls. Six-month CDs were employed, since they have little implicit pricing and are the most frequently used maturity by both thrifts and banks, thus, allowing consistent data across sample institutions.[12] The weekly observations coincide with the typical decision time frame for banks. Rates are reported by institutions for each Wednesday. The sample period, March 7, 1984 to December 29, 1987, produced 200 weekly observations for each firm. Corresponding weekly rates for six-month Treasury bills from auctions announced on each Tuesday were collected from the *Federal Reserve Bulletin*. Financial data for individual banks and S & Ls in 1985 were taken from the Lyons, Zomback, and Ostrowski, Inc. *Depository Institutions Performance Directory* (1984–1986).

A second control sample of banks and S & Ls outside of Ohio was obtained from the *Bank Rate Monitor* (BRM) which conducts surveys of the five largest banks and five largest thrifts in major metropolitan areas on a weekly basis. Data for this sample includes the 104 weekly observations during 1984 and 1985.[13] On a caveat, since this sample represents a subset of bank and S & L CDs for the largest firms in six of the nation's largest cities, the results should not be generalized to other smaller non-Ohio banks and S & Ls. This

[12] We requested other maturities from banks and S & Ls in our survey, but the data for other maturities was incomplete. Many S & Ls pointed out that they did not keep this type of historical data. Six-month CDs are the only retail deposit maturity with consistent data for the sample during the time period examined. Ideally, we would have also analyzed deposit withdrawal data; however, regulators consider this data to be confidential, and it is unavailable to the public. Nonetheless, CD prices should reflect volume information, since bank managers will attempt to retain funds by paying a higher deposit rate.

[13] The composition of the BRM sample changed in 1986, so we only use the 1984–1985 period. BRM converts rate data into effective yields incorporating all CD characteristics including the minimum deposit amount and the relevant compounding periods. For the Ohio sample, the majority of the depository institutions surveyed did not have yield data, so rate data based on a particular bank's compounding periods, minimum deposit balances, and prepayment penalties were used. Since the purpose of the study is to observe unexpected changes in rates, as long as consistent compounding periods and deposit balances were used for individual institutions, the use of rate versus yield data should not affect the validity of the study. The correlation between the yield and rate data was found to be above 0.95. Both the BRM and Ohio data behaved similarly, with a one-lag structure in response to T-bill rates.

sample is shown in Appendix B, and includes all the firms with available data surveyed by the Bank Rate Monitor during this time period. The cities surveyed by the BRM with available data included Chicago, Detroit, Los Angeles, Philadelphia, New York and San Francisco.[14]

We made a careful examination of several annual periodical indices, with primary reliance on *The Wall Street Journal Index, Funk and Scott's Predicasts*, and regional newspapers to ensure that simultaneous events that might bias the empirical results did not occur during and surrounding the crisis period that we examine. Of the sample institutions, only American S & L in Los Angeles (Financial Corporation of America) had a simultaneous event, a report of unexpected losses leading to sizable deposit withdrawals during Week 1 of the crisis (March 13), so it was eliminated from the sample. Any impact on rates for California depository institutions could have been affected by this event; hence, all firms in San Francisco and Los Angeles were eliminated.

B. Time-Series Intervention Analysis

Previous studies of bank debt rate contagion have developed multivariate regression models of risk spreads using macroeconomic or firm specific data (see Stover and Miller (1983) and Fraser and McCormack (1978)). Because these data are available only on a monthly or quarterly basis, the exact period when an impact occurred and the nature of an effect could not be precisely detected.

In contrast to previous studies we use an alternative methodology, a multivariate Box Jenkins/ARIMA model, sometimes referred to as a transfer function (see Vandaele, 1983). ARIMA models have advantages in terms of easily removing possible autocorrelation, producing accurate forecasts and allowing structural changes in parameters during pre-event and post-event periods (see Larcker, Gordon, and Pinches (1980)). Edmister and Merriken (1989, pp. 34–35) point out that time-series techniques are particularly appropriate for examining consumer deposit markets, where "information diffusion causes a gradual response," due to the existence of switching costs. With intervention analysis, the strength and structure of an effect can also be quantified in terms of its duration (temporary or permanent) and speed (gradual or immediate).

Due to a significant drift in the mean, first differencing was required to achieve stationarity for the data in the sample, so we use the weekly change in our endogenous variable and the change in the T-bill rate as our exogenous variable.

Changes in retail CD rates (CD) are modeled as a function of national economic factors, proxied by changes in the six-month T-bill rate, and changes

[14] The Federal Reserve Board and the Office of Thrift Supervision (formerly the FHLBB) did not begin collecting rate data on a periodic basis (monthly) until October 1983. Problems with the accuracy of these data in the early years of collection have also been acknowledged (at least informally) by regulators.

in local market factors, represented by the stochastic variation in previous CD rates.[15] Our ARIMA transfer function intervention model can be written as:

$$\mathrm{CD}_t = A(B) I_t^T + \frac{U(B)}{(1-B)^d S(B)} \text{T-bill}_t + \frac{Q(B)}{(1-B)^d Z(B)} a_t, \tag{1}$$

where I_t is the intervention variable, which begins in period T, (equal to 1 in the event period and 0 otherwise), and A is the coefficient of the intervention variable representing the impulse function of the response to an event. B is a backshift operator, specifically (B) T-bill$_t$ = T-bill$_{t-1}$ and (B^2) T-bill$_t$ = T-bill$_{t-2}$. The variable a_t represents the usual random shock or white noise series of the model. The numerator terms U and Q represent the moving average portion of the model; and the denominator terms S and Z, the autoregressive portion, for the respective variables. The term d refers to the degree of differencing necessary to achieve stationarity for the time series data. The identification, parameter estimation, and diagnostic checking of the stochastic variation are discussed in Van Fenstermaker and Filer (1986), Vandaele (1983) and Pankratz (1983) and will not be considered in detail in this paper.

The coefficient A for the intervention term I_t for this analysis is defined as:

$$A(B) = \frac{w(B)}{1 - g(B)} P_t^T, \tag{2}$$

which represents the impulse function of the response to an event, where P_t is the impulse variable (analogous to a dummy variable) and T denotes the time period at which the event occurs.[16] The numerator parameter (NUM)

[15] ARIMA models operate under the assumption that the behavior of a time-series variable can be explained using its past observations. The six-month T-bill rate is assumed to capture macroeconomic variables used by depository institutions in setting their CD rates. This assumption is plausible since T-bill rates are set in weekly auctions that should be informationally efficient, and bankers use these rates along with local market factors, to set their own six-month CD rates. Past observations of CD rates are used to reflect local market and regional factors. These models have been found in practice to frequently outperform the forecasting accuracy of multivariate models (see Pankratz (1983) and Riise and Tjostheim (1984)).

[16] The original event dummy variable is equal to 0 for the pre-event period and 1 for the post-event period. However, to be consistent with the first differencing on both dependent and independent variables, we perform a first differencing on the event dummy as well. As a result, the event dummy is equal to 1 on an event date and zero elsewhere. The first differencing event dummy can be interpreted as the abnormal change of CD rates on that particular date. Unless there is an offsetting effect later, the increase in CD rate is permanent. Nonetheless, one can include as many first differencing event dummies that have a value of one at particular dates as indicated by the cross correlation function.

represents the magnitude of the impulse, and the denominator parameter (DENOM) measures whether the effect is gradual or immediate. A positive DENOM will signify a gradual increase in the effect over time, and a negative value, a gradual decrease (see Vandaele (1983), p. 337). Note, if the data dictate, several intervention terms can be fitted to represent the offsetting reaction, or the complex nature of the intervention.

We estimate the transfer function intervention model for subsamples categorized by location and capital ratios. Our Ohio sample data fell into three distinct geographical regions by consolidated metropolitan statistical areas (CMSAs): Southern Ohio (Cincinnati, Dayton, Columbus), Northeastern Ohio (Cleveland, Mansfield, Youngstown), and Northwestern Ohio (Toledo, Lima), see Appendix A. The Ohio data was also subdivided into quartiles according to capital ratios to test for the existence of risk-based pricing. For the Bank Rate Monitor sample, intervention models were estimated for S & Ls and banks in each respective city.

The steps in estimating our transfer function consisted of first identifying and estimating univariate ARIMA models for the T-bill and CD series in order to prewhiten or factor out the systematic or seasonal component of each series to obtain the residuals.[17] Then the residuals were used to estimate the autocorrelation and partial autocorrelation functions of each series and to calculate the cross correlation function between the T-bill and CD series. These functions were utilized to identify the structure of the transfer function model which was estimated using a maximum likelihood, nonlinear estimation procedure. The cross correlation between the transfer-function residual and the event dummy was employed to determine the pattern of the intervention term.[18] In the final stage, the transfer function model was reestimated with the inclusion of the event dummy to yield the final estimate of the exact intervention.

C. Panel Data Analysis

To verify that the detection of any significant intervention effects on CD rates is not methodology specific, we perform a panel data analysis that considers simultaneously both the time-series and cross-sectional characteris-

[17] The pre-whitening procedure is necessary because when two series are highly self-autocorrelated, the relationship as indicated by the cross-correlation function between the two time series can be difficult to interpret and can even be misleading (see Vandaele, 1983, p. 268).

[18] Assuming the event dummy is not related to T-bill rate changes, a cross correlation is computed to identify the magnitude and the timing of the event's impact on the change in CD rates at time t. Residual plots, chi-square statistics, as well as residual cross-correlations with the independent variables, are used to check the adequacy of the model. To make the model more complete, the rebound of the change in CD rates at time t was also modeled.

tics of the data. Our model is specified as follows:

$$\begin{aligned}\text{CDRATE}_{it} = B_0 &+ B_1\,\text{T-BILL}_t + B_2\text{REGION}_i \\ &+ B_3\text{SIZE}_i + B_4\text{CAPITAL}_i \\ &+ B_5\text{EVENT}_i + B_6\text{EVENT} \times \text{CAPITAL} + e_{it} \qquad (3)\end{aligned}$$

The variables are described below:

The dependent variable is CDRATE_{it} = six-month CD rates for depository institution i at time t. Alternatively, we used rate spreads over the six-month T-bill rate, with similar results.

Control variables include: T-BILL = six-month T-bill rates to control for national market movements, REGION = mean regional six-month CD rates to control for local market effects,[19] SIZE = the asset size of individual depository institutions (in thousands), to control for implicit pricing factors, and CAPITAL = depository institution's tangible (primary) capital to total assets ratio as a proxy for insolvency risk to control for any risk-based pricing in the noncrisis period.

Test variables include: EVENT = a dummy variable equal to 1 during the crisis period and 0 otherwise, to test for an average increase/decrease in rates during the crisis period, EVENT × CAPITAL = an interactive variable that is the product of the variables EVENT and CAPITAL to test for risk-based pricing during the crisis period, e = error term with mean of zero and a constant variance.

We estimate the model using a cross-sectional, time-series analysis for the 37 Ohio banks and 32 Ohio S & Ls in our sample for 200 weeks, with respectively 7,400 observations for the bank sample and 6,400 observations for the S & Ls. A Chow test indicated that a structural difference exists between S & Ls and banks;[20] hence, separate regressions were performed for each type of institution.

The control variables are factors found to be significant in a recent study of CD-rate pricing by Cooperman, Lee, and Sinkey (1990). The test variables allow tests for the predictions of the contingent guarantee hypothesis of depositor reactions. The contingent insurance guarantee hypothesis predicts that deposit premiums during the Ohio crisis should be risk-based, i.e., based on the financial condition of individual firms, as proxied by a depository

[19] The regional rate variable is included to control for local market (supply and demand) effects. Studies examining deposit price/concentration relationships for a broad cross-section of banks have found that deposit markets continue to be affected by local market conditions (See Berger and Hannan (1989)).

[20] A Chow test is conducted based on the sets of variables included in Table 3, with the exception of the event dummy and the interaction term between the dummy and capital ratio. We strongly reject the null hypothesis that the bank and S & L regressions can be pooled for a joint estimation. Further tests also indicate that differences occur in both the intercept and slopes. This result is in agreement with previous studies.

institution's tangible capital ratio. Consequently, both the EVENT and the EVENT × CAPITAL variables would be expected to be negative and significant under this hypothesis.[21] The EVENT variable along with the EVENT × CAPITAL variable would reflect a change in depositors' rational expectations towards the condition of the FSLIC, with the EVENT × CAPITAL variable reflecting greater potential inconvenience costs associated with reimbursement delays for less solvent depository institutions.

IV. Empirical Results

A. Intervention Analysis

For each of the ARIMA models, banks and thrifts were found to respond to T-bill rate changes with a one-week lag. This relationship was observed for both the Bank Rate Monitor and Ohio data. This result is consistent with other recent studies examining retail rate adjustments in deregulated markets indicating that six-month CDs are interest-rate sensitive and adjust quickly to market rates.

Table II summarizes the empirical results of the intervention analysis for the Ohio sample. The DENOM variables were not significant for any of the analyses, indicating that the effect of the crisis on CD rates was a sudden versus a gradual effect. Since first-differencing was performed for each of the time-series rate data to achieve stationarity, the NUM variable can be interpreted as the unexpected change in rates that occurred during the initial week of the crisis (March 6, 1985), after adjusting for local and national market effects.

The top panel of Table II indicates the intervention results by location. Rate premiums for S & Ls in all regions were affected by the crisis, with the

[21] As suggested by Johnston (1984, Chapter 6), an example interpreting the coefficients of interactive variables can be provided using a simplified version of our regression equation:

$$\mathrm{CD} = B_0 + B_1\mathrm{Capital} + B_2\mathrm{Event} + B_3\mathrm{Capital} \times \mathrm{Event} + e. \tag{1}$$

In a nonevent period, the expected CD rate is:

$$E(\mathrm{CD}/\mathrm{Event} = 0) = B_0 + B_1\mathrm{Capital}. \tag{2}$$

During an event period, the expected CD rate becomes:

$$E(\mathrm{CD}/\mathrm{Event} = 1) = B_0 + B_1\mathrm{Capital} + B_2 + B_3\mathrm{Capital}. \tag{3}$$

Subtracting (2) from (3) we obtain:

$$E(\mathrm{CD}/\mathrm{Event} = 1) - E(\mathrm{CD}/\mathrm{Event} = 0) = B_2 + B_3\mathrm{Capital}. \tag{4}$$

Differentiating equation (4) with respect to capital, we obtain B_3, which is the sensitivity of the expected CD rate to changes in a depository institution's tangible capital ratio during the event period. Therefore, the incremental change in the expected CD rate during the crisis period can be obtained using equation (4), i.e., the coefficient on the event dummy variable plus the product of the coefficient on the CAPITAL × EVENT variable and a firm's capital ratio.

Table II

Intervention Analysis Results Ohio Banks and Thrifts

	NUM (Start)	NUM (End)	Start	End
Intervention by Location				
Savings and Loans				
Cincinnati	0.06†	−0.0828*	0	7
Cleveland	0.07§	−0.115*	0	7
Toledo	0.162*	−0.116§	0	5
Banks				
Cincinnati	0	−0.11*	0	3
Cleveland	0.097†	−0.145*	0	1
Toledo	0	−0.09†	0	7
Intervention by Capital Ratio				
Savings and Loans				
Low	0.106*	−0.117*	0	7
Medium	0	0	0	0
High 1	0	0	0	0
High 2	0	0	0	0
Banks				
Low	0.131†	−0.118*	0	5
Medium	0.106§	−0.182*	0	2
High 1	0.063§	−0.107*	0	1
High 2	0.082§	−0.152*	0	1

Start indicates when the impact began, as indicated here in week 0 (March 6, 1985) for each category, and End indicates the date in which rates returned to normal levels. Since first differencing was performed for each of the time-series rate data to achieve stationarity, NUM (Start) indicates the unexpected change in rates that occurred on the start date, and NUM (End), the size of the sudden decline in rates when rates returned to normal. Location categories are for the Cincinnati (Southern Ohio), Toledo (Northwest Ohio), and Cleveland (Northeast Ohio) regions. Our sample fell into these three distinct regions. Capital categories represent different percentiles for tangible capital for S & Ls and primary capital for banks. The DENOM variables (not shown) were insignificant, indicating that the effect of the crisis on CD rates was sudden and ended suddenly, versus a gradual effect. $N = 200$ time-series observations.

* t-statistic significant at 1% level.

† t-statistic significant at 5% level.

§ t-statistic significant at 10% level.

largest reaction in the Toledo region.[22] An unexpected rise in rates occurred for S & Ls in Cincinnati and Cleveland of 6 to 7 basis points (bp) in Week 0, with premiums recovering in approximately seven weeks. In Toledo, the mean rise was larger, 16 bp, and rates recovered in five weeks.

Rate premiums for banks in the Cleveland region were also affected, with a 9.7 bp rise in Week 0. In Week 1 of the crisis, however, an unexpected drop in

[22] This reaction does not include State Home Savings of Bowling Green, Ohio, which suffered deposit runs and a 48.9 basis point rise in CD rates during the crisis, due to confusion between its name and Home State Savings in Cincinnati. When State Home was included, the Toledo S & L reaction was even larger.

the mean rate premium occurred of 14.5 bp, more than offsetting this rise. Cincinnati and Cleveland did not have a significant intervention term during the initial crisis week. However, the mean rate premium in Cincinnati fell unexpectedly in Week 3 by 11 bp, suggesting that Cincinnati banks received a net benefit from the crisis. In Toledo, a sudden premium drop occurred of 9 pb, but not until Week 7. The results indicate some geographic effects of the crisis, with S & Ls in Cincinnati and outside Cincinnati experiencing a rise in rate premiums during the initial event week which lasted about seven weeks. Ohio banks in two regions appeared to experience a net drop in rate premiums during Weeks 1 and 3 of the crisis.

When the intervention analysis was performed by capital ratio categories, as shown in the lower panel of Table II, the intervention term was only significant for the lowest capital (technically insolvent) thrifts, with a significant rise in the average CD rate of 10.6 basis points, again occurring during Week 0 and returning to a normal level in Week 7. This result suggests that the unexpected rise in CD rates during the initial week of the crisis was a function of a S & L's insolvency risk, proxied by its tangible capital ratio.

For the banks in the sample, which all met the minimum FDIC capital requirement of 6% at this time (with primary capital ratios ranging from 6.91–12.35%, versus −8.11–14.5% for the tangible capital ratios for S & Ls), there was a significant interaction variable across all capital categories. The highest intervention term occurred for the lowest capital category, 13.1 bp, with the risk premium returning to normal in Week 5.

At first glance, it appears that the other more well-capitalized bank categories were also negatively affected by the crisis, as a function of capital. However, the second, NUM(End) column indicates that each of the three higher capital categories also experienced a sudden drop in rate premiums during Weeks 1 or 2. In each case, the drop in the premium was larger than the previous rise in Week 0, resulting in a net decline in rates during the crisis. For instance, the highest capital category (High 2) had a sudden rise in its rate premium of 8.2 bp in Week 0, but a subsequent fall of 15.2 bp in Week 1, or a net decline of 7 bp. Similarly, the next two capital categories, High 1 and Medium, had a net decline during Weeks 1 and 2, of 4.4 and 7.6 bp, respectively. Bank managers may have viewed the crisis in Week 0 as an opportunity to attract new deposits from less solvent firms by initially offering higher rates. These rates were subsequently lowered in Weeks 1 or 2. The rise in the mean rate premium that lasted for five weeks for the lowest bank capital category, in contrast, may be related to the tendency for bank customers to be more risk averse than S & L depositors (i.e., self-selection bias). Managers of low capital banks may have feared losing deposits to better capitalized firms, despite the fact that the FDIC was well-capitalized at this time.[23]

[23] In a separate analysis we also examined the response in rates by size percentiles. There was no clear pattern in the response by size category, suggesting that size is not a good proxy for risk.

The intervention terms for the banks and S & Ls outside Ohio in each of the respective cities were insignificant, so for the sake of brevity are not shown in Table II. Since the Bank Rate Monitor sample consists of a segment of the nation's largest banks in the nation's largest cities, the lack of a CD pricing reaction may reflect a consensus by depositors that these banks are "too big to fail." On a note of caution, since this sample represents a subset of the largest firms operating in five of the nation's largest cities, the results should not be generalized to other smaller institutions or bank markets.

On balance, the initial crisis event resulted in significantly higher CD rates for less solvent banks and S & Ls in Ohio with the increase in CD rates lasting approximately seven weeks, consistent with the contingent insurance guarantee hypothesis.

B. Cross-Sectional / Time-Series Analysis

The results of our panel data analysis are shown in Table III. The regression results appear to be quite robust to the model specification.[24] The sensitivity of CD rates to T-bill rate changes is confirmed, with CD rates rising only slightly less than 1% with a 1% increase in T-bill rates. For both banks and S & Ls, CD rates fell as the asset size of firms increased, suggesting that depositors may have more confidence in larger firms. Alternatively, larger firms may be able to offer lower CD rates, because of greater implicit pricing in terms of a larger number of services offered. Bank and S & L CD rates also appeared to be strongly related to local market competition and local economic conditions, with a large and significant coefficient on the regional rate variable.

The capital ratio variable indicates a small amount of risk-based pricing for S & Ls during the nonevent period. A 10% increase in a firm's capital ratio would reduce CD rates by 10 basis points. During the event period, however, as indicated by the EVENT × CAPITAL interaction variable, risk-based pricing occurred for both banks and S & Ls with a larger coefficient for banks. Also, the event dummy variable was significant for both S & Ls and banks, and the coefficient is larger for banks. However, one should not infer that banks were more seriously affected by the crisis; the EVENT × CAPITAL interaction variable coefficients must be interpreted in conjunction with the distribution of the capital ratios, which are presented in Table IV. The statistics indicate that the range of S & L capital ratios is wider, and the

[24] Several alterations were made to check the robustness of the regression analysis. Since CD rate changes are found to lag behind the T-bill rate changes by at least one week, we replaced the current of T-bill rate by a lagged T-bill rate. The coefficient of determination becomes larger, and the conclusion is largely unchanged. The results of risk-based pricing are also not sensitive to the width of the event window from 1 week to 6 weeks. Using spread as the dependent variable and removing the T-bill rate as a variable, we had similar results.

Table III

Cross-Sectional / Time-Series Analysis

Dependent variable: Weekly six-month CD rate in percent for the *i*th firm for Ohio Banks and Thrifts

Independent Variables	Equations: Banks	Equations: S & Ls
Intercept	0.372	0.842
	(5.28)*	(11.67)*
Event dummy variable	0.702	0.062
	(3.39)*	(1.58)
T-bill rate (%)	0.955	0.956
	(271.30)*	(224.25)*
Capital ratio (%)	0.013	−0.011
	(2.42)*	(−7.28)*
Event × capital interaction variable	−0.072	−0.016
	(−3.01)*	(−2.38)*
Ln(asset size) (thousands)	−0.021	−0.089
	(−5.49)*	(−8.77)*
Regional rate	0.807	0.928
	(6.16)*	(15.79)*
Adjusted R^2	0.933	0.906
F-value	12620.67*	8653.54*
N	7,400	6,400

The coefficients on the independent variables were similar when spreads were used as the independent variable, and the T-bill rate was removed as an independent variable. The event dummy represents the period $t = 0$ to $t = 10$, where $t = 0$, represents the first event week, March 6, 1985 of the crisis.

* Significance at the 1% level per both t-values and F-values.

distribution is more platykurtic than that of banks. The full effect of the crisis for banks and S & Ls at a particular capital level have been calculated in column two of this table, where the full effect is equal to the event dummy plus the product of the capital interaction variable and a firm's capital ratio. The calculations in column two indicate that only S & Ls with negative capital ratios and banks with primary capital ratios below approximately 9% are adversely affected. As predicted by a contingent insurance guarantee hypothesis, banks and S & Ls with minimum capital ratios would have to offer the highest rate premiums, in this case approximately 20 bp. The results strongly suggest the emergence of risk-based pricing during the event period, whereby as Kane argues a weakening of depositor confidence in insurance funds may result in the threat of silent runs, necessitating a rise in deposit rates by less solvent firms as compensation for depositors.

Table IV
Full Impact of the Crisis Event by Capital Ratio Distribution

Distribution	Capital Ratio (%)	Full Impact of Event (Basis Points)
Banks		
Maximum	12.35	−18.72
75th percentile	9.22	3.81
Median	8.44	9.43
Average	9.54	8.71
25th percentile	7.90	13.32
Minimum	6.91	20.45
Savings and Loans		
Maximum	14.50	−17
75th percentile	5.96	−3.34
Median	3.85	0.04
Average	2.38	2.39
25th percentile	−0.61	7.18
Minimum	−8.11	19.18

The full impact of the event for each individual thrift can be calculated by multiplying the adjusted coefficient on the capital × event interaction variable by the firm's capital ratio and then adding the coefficient on the event dummy variable. The first column lists the capital ratio distribution for banks and thrifts, respectively. The second column indicates the full impact of the crisis for each capital ratio.

IV. Summary and Conclusion

This paper uses an ARIMA transfer function intervention model and alternative panel data analysis to examine the effect of the Ohio thrift crisis in 1985 on the pricing of six-month retail CDs for federally-insured Ohio banks and S & Ls. Adjusting for local and national market rate changes, we find an unanticipated rise in CD rate premium for less solvent banks and S & Ls during the initial event week with rates returning to normal in approximately seven weeks. Our panel data analysis confirms the existence of risk-based pricing during the crisis period, consistent with a contingent insurance guarantee hypothesis, suggested by Kane (1989).

Kane (1990) argues that the FSLIC was insolvent much earlier than the GAO announcement in March 1987. Due to potential post-government jobs, regulators found ways to window dress the FSLIC's financial position and denied its bankruptcy. The finding of risk-based pricing during the crisis period indicates the public's awareness and concern over federal insurance fund problems as early as 1985.

Studies by Cook and Spellman (1991) and others have demonstrated that depositors may price deposits based on risk if their rational expectations increase concerning future costs in terms of a guarantor's handling of an insolvency, or the government's handling of an insolvent guarantor. Accordingly shocks, such as the announcement of an insurer's potential insolvency, may change depositors' rational expectations. Bank and savings and loan managers may anticipate depositors' reactions when a shock occurs and, accordingly, price deposits based on a firm's insolvency risk. The Ohio S & L crisis may have represented this type of shock.

Our results also suggest that in theoretical models of deposit prices and bank contagion, a consideration of the explicit financial strength of a federal deposit insurer needs to be made.

Appendix A: Ohio Banks and Thrifts in Sample (1985 asset size in thousands)

Banks	
Southern Ohio (Cincinnati, Dayton, and Columbus CMSAs)	
Star Bank, N.A., Cincinnati	$2,197,555
Bank One, Athens	165,755
Dime Bank, Marietta	23,917
Farmers & Traders National Bank, Hillsboro	66,761
Ohio Valley Bank Co., Gallapolis	133,939
The Second National Bank, Hamilton	187,154
Glouster Community Bank, Glouster	23,670
The First Bremen Bank, Bremen	21,240
Citizens Bank, DeGraff	12,923
First Knox National Bank, Mt. Vernon	185,275
Park National Bank, Newark	499,136
North East Ohio (Cleveland, Mansfield, and Youngstown CMSAs)	
Ameritrust, Cleveland	7,571,547
Central National Bank, Cleveland	2,319,000
National City Bank, Cleveland	5,524,549
Society National Bank, Cleveland	5,043,000
United Bank, Bucyrus	79,600
Bank One of Mansfield	318,547
Old Phoenix National Bank, Medina	267,120
The Republic Banking Co., Republic	10,844
Dollar Savings & Trust Co., Youngstown	873,230
Northwest Ohio (Toledo, Lima, CMSAs)	
First National, Toledo	739,908
Huntington National Bank, Toledo	6,163,560
Mid-American National Bank & Trust, Toledo	355,644
Ohio Citizens, Toledo	907,144
Toledo Trust, Toledo	2,183,716
Century Bank, Toledo	121,000
Hamler State Bank, Toledo	22,912
State Bank & Trust Co., Defiance	124,117

Home Banking Co., Gibsonburg	75,371
Exchange Bank, Luckey	48,736
Old Fort Banking Co., Old Fort	36,952
National Bank of Montipelier, Montpelier	46,690
National Bank of Oak Harbor, Oak Harbor	50,467
Croghan Colonial Bank, Fremont	170,815
The Metropolitan Bank, Lima	180,956
Citizens Commercial Bank & Trust, Celina	140,860
Commercial Bank, Delphos	91,317
Savings and Loans	
Southern Ohio (Cincinnati, Dayton, and Columbus CMSAs)	
Belmont Federal S & L, Bellaire	125,094
Brentwood Savings Association, Cincinnati	67,062
Cottage Savings Association, Cincinnati	165,771
Fidelity Federal Savings and Loan, Cincinnati	132,852
Harvest Home Savings Association, Cincinnati	57,426
Merit Savings Association, Cincinnati	86,266
North Cincinnati Loan & Bldg, Cincinnati	53,362
Suburban Federal Savings & Loan, Cincinnati	100,123
Mid Fed Savings Bank, Middleton	249,814
Mutual Federal S & L, Zanesville	337,849
Gem Savings, Dayton	1,516,134
Mt. Vernon S & L, Mt. Vernon	91,838
People's Savings & Loan, Xenia	159,378
North East Ohio (Cleveland, Mansfield, and Youngstown CMSAs)	
Broadview Savings & Loan, Cleveland	2,096,000
Cardinal Federal Savings, Cleveland	1,594,000
Ohio Savings Association, Cleveland	1,438,000
Horizon Savings and Loan, Cleveland	180,864
Superior Savings and Loan, Cleveland	121,188
Third Federal Savings, Cleveland	1,491,000
Transohio Savings, Cleveland	3,214,000
Women's Federal Savings Bank, Cleveland	660,655
First S & L Co., Masillion	201,309
First Federal S & L, Lakewood	284,286
First Federal Savings Association, Youngstown	362,803
First Federal S & L, Warren	266,088
Trumbull S & L, Warren	253,069
People's Federal Savings Bank, Wooster	403,745
Northwest Ohio (Toledo, Lima, CMSAs)	
First Federal S & L, Lima	191,163
People's S & L, Toledo	402,714
United Home Federal, Toledo	583,488
First Federal Savings, Bowling Green	83,621
State Home Savings, Bowling Green	293,847

Appendix B:
Depository Institutions in Sample in Major Metropolitan Areas

Commercial Banks	Thrifts
Chicago	
First National Bank	Talman Home Federal
Continental Illinois	Citicorp Savings
Harris Bank	St. Paul Federal
Northern Trust Co.	Bell Federal Savings
American National Bank	Cragin Federal
Detroit	
National Bank of Detroit	First Federal-Michigan
Comerica	Standard Federal Savings
Manufacturer's National	First Federal Savings Bank
First of America	Heritage Federal Savings
Los Angeles	
Security Pacific	American Savings and Loan
Union Bank	Home Savings
Lloyd's Bank	Great Western
City National	California Federal
	Glendale Federal
New York	
Citibank	Goldome (MSB)
Chase Manhattan	Dime Savings (MSB)
Chemical Bank	Bowery Savings (MSB)
Bankers Trust	Emigrant Savings (MSB)
Philadelphia	
Philadelphia National Bank	PSFS (MSB)
First Pennsylvania Bank	Beneficial Mutual Savings (MSB)
Provident National Bank	First Federal Savings and Loan
Mellon Bank	Germantown Savings Bank (MSB)
Continential Bank	Horizon Financial
San Francisco	
Wells Fargo	World Savings and Loan
Crocker	First Nationwide Savings
California First	Imperial Savings and Loan
First Interstate	Citicorp Savings

(MSB) indicates a mutual savings bank, insured by the FDIC. All institutions had assets of $1 billion or larger, except Heritage Federal Savings in Detroit which had $.5 bil.

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