NBER WORKING PAPER SERIES

CHANGE IN MARKET ASSESSMENTS OF DEPOSIT-INSTITUTION RISKINESS

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Working Paper No. 2530

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
March 1988

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ABSTRACT

Using the Goldfeld and Quandt switching regression method, this paper investigates variability over 1975-85 in the risk components of bank and saving and loan stock. We develop evidence that the market-beta, interest-sensitivity, and residual risk of deposit-institution stock vary significantly during this period. Reassessing previous event studies in light of these findings suggests that event-study methods tend to overreach their data.

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1. Introduction:

Testifying before the House of Representatives in December 1984, the U.S. Comptroller of the Currency posed the following questions:

> Is banking becoming riskier? Are large banks riskier than small banks? What effect has interest rate deregulation had on the risks faced by the banking system?.. We do not have good answers to these questions.

This paper deploys statistical methods to develop new evidence on these issues. Using capital market data for 1975-1985, we investigate temporal variability in market and interest-rate sensitivity and in unsystematic risk for savings and loan associations (S&Ls) and three size classes of banks. During this decade, deposit institutions faced substantial regulatory and statutory change and interest rates became extraordinarily volatile.

Our methods focus on estimating the parameters of a two-index model, allowing the model to switch parameter regimes over 1975-85 in an unrestricted way. Goldfeld and Quandt's [(1972), (1973), (1976)] switching regression method (GQSRM) is used to determine the temporal variability of model parameters. GQSRM employs a search procedure which finds maximum-likelihood estimates of three attributes of regime variation: the shift dates, the gradualness of each shift, and the parameters of the stochastic process which obtain in each regime. GQSRM is applied by Quandt (1974) and Lin and Oh (1984) to solve specific problems of nonstationarity and Unal (1985) use the approach as a tool of event study.

The statistical findings and the analysis presented in this paper serve to reconcile some of the conflicting results developed in the literature on the two-index model and provide grounds for reassessing traditional event-study methods. We find that the riskiness of bank and S&L stock declined in the late-1970s, but rose again in recent years. We show that event studies focusing on the impact on deposit-institution stock of recent changes in monetary-policy regimes or regulatory adjustments need to control for other significant sources of nonstationarity. Allowing for this nonstationarity suggests that at least over 1979-1982 it is rash to attribute observed shifts in return-generating processes to any particular information event.

2. Model and Data Selection

This paper's focal return-generating process is the two-index model developed by Stone (1974). This asset pricing model expands the familiar market model of asset returns by adding an interest-rate index, in this application as a quasi-industry factor. Including an interest-rate index as a second factor could just as easily be justified by Merton's (1973) intertemporal capital asset pricing model specification. This model expresses return on asset p as:

$$\tilde{R}_p = \beta_o + \beta_m \tilde{R}_m + \beta_i \tilde{R}_i + \tilde{e}_p \quad . \tag{1}$$

β_m and β_i are measures of the asset's systematic market and interest-rate risk; $\tilde{R}_m$ and $\tilde{R}_i$ represent stock-market return and a return on a debt index. Lloyd and Shick (1977), Lynge and Zumwalt (1980), Chance and Lane (1980), Flannery and James (1984a), Scott and Peterson (1986) and Brewer and Lee

(1985) investigate the extent to which this model can explain returns on financial-intermediary stock.[1]

Using the two-index model raises two major problems. First, one must specify whatever simultaneous relation is presumed to exist between the variables. Previous authors deal with what is called "multicollinearity" between R_i and R_m by imposing an arbitrary causal ordering. Because theory does not impose a zero covariance, σ_{mi}, between the market return and an interest rate index, the following orthogonalization, is proposed [Stone (1974), Chance and Lane (1980)].

$$\tilde{R}_i^{*} = \tilde{R}_i - [\, \sigma_{mi} / \sigma_m] \, \tilde{R}_m. \tag{2}$$

An alternative approach is to treat R_i as exogenous and R_m as endogenous [Lynge and Zumwalt (1980), Flannery and James (1982)]. This produces the adjusted market factor, $\tilde{R}_m^{*}$:

$$\tilde{R}_m^{*} = \tilde{R}_m - [\, \sigma_{mi} / \sigma_i] \, \tilde{R}_i. \tag{3}$$

Either an adjusted interest-rate index, R_i^*, or an adjusted market factor, R_m^*, may be incorporated into Equation (1). In both cases, orthogonal series are generated, which eliminate simultaneity by construction. However, as Giliberto (1985) elaborates, using Equation (2) biases t-tests against interest-rate sensitivity, while using Equation (3) imposes the opposite bias. Either auxiliary regression grafts a nonsimultaneous triangular structure onto the two-index model. We would need additional theory to specify either that R_m drives the interest rate, that R_i is logically prior to R_m, or that both variables are governed by omitted exogenous forces. Because the parameter space spanned is the same whether or not we impose either auxiliary regression, this paper employs the unorthogonalized two-index model.

A second problem is to select a proxy for unanticipated changes in R_i. One approach is to pick an interest-rate index, estimate synthetic forecasts *via* an autoregressive model, and use forecast errors from this model as proxies for unanticipated interest-rate changes [Flannery and James (1984a)]. An alternative approach is to use changes in the yield on a given maturity of long-term bonds to capture unanticipated changes in interest rates [Scott and Peterson (1986); Sweeney and Warga (1986)]. Yet another approach is to proxy unanticipated changes in interest rates by the difference between the spot 3-month Treasury bill rate at time t and the forward 3-month Treasury bill rate imbedded in the yield curve at time t-1 [Mishkin (1982) and Brewer and Lee (1985)]. Unal and Kane (1987a) also examine the impact of choosing between unanticipated and anticipated series for both long and short interest-rate indices. Their analysis indicates that, although bank and S&L stock returns are not responsive to short rates, long rates have a significant effect. This pattern holds irrespective of whether interest rates are proxied by errors from autoregressive forecasts or by components orthogonal to R_m. Relying on Occam's Razor, this paper employs unadjusted holding-period returns on long-term government bonds to proxy the unanticipated changes in the interest-rate index.

Our samples are constructed from data tapes prepared by the Center for Research in Security Prices (CRSP) at the University of Chicago. All commercial banks and S&Ls are selected for which CRSP tapes show monthly returns for the 1975-85 period.[2] Appendix 1 lists these large banks and S&Ls and gives their asset size. To test for possible intra-industry differences, banks are further classified into three classes. Following a suggestion from Joseph Sinkey, a class of "money-center banks" is determined from Citicorp's competitor list (Sinkey, 1986, p.249). Other banks are classified as "superregional" and "regional" banks, based on an asset-size threshold of $10 billion. Partitioning banks in this way allows for differences in the extent to which the nondeposit debt of different institutions is effectively guaranteed by the FDIC and develops classes of banks that may be presumed to operate similarly. At the end of 1984, the mean asset sizes of the money-center, superregional, and regional banks are $73.6, $20.6, and $4.5 billion. Sample S&Ls average $13.7 billion. Total assets of the 31 banks and 8 S&Ls in our sample are $973 billion and $109.6 billion, respectively. Expressed as a percentage of respective industries' asset totals, sample banks constitute about 43 percent and sample S&Ls about 12 percent.

Equal-weighted portfolio returns are constructed for each class. The CRSP value-weighted NYSE and AMEX stock index adjusted for dividends is used as the market proxy. Interest rates are proxied by the monthly holding-period return on long-term government bonds, obtained from Ibbotson Associates (1986). As Flannery and James (1984) point out, "since the holding period returns on bonds are negatively correlated with changes in the level of interest rates, a positive value for β_i implies that the firm's market value declines when interest rates rise."

3. Estimating the Model Over Prior Studies' Observation Periods

Unal and Kane (1987a) emphasize that the qualitative force of empirical evidence on the two-index model varies with the market proxy and the interest-rate index used, the time period analyzed, the frequency with which data are observed, and the number of institutions sampled. Because our sample and model specification both differ from previous studies, for comparative purposes Table 1 reports how our sample and proxy specifications behave over the specific time periods analyzed by Flannery and James (1984), Scott and Peterson (1986), and Brewer and Lee (1985).

In these studies and in our parallel runs, bank stock returns show a market-beta of considerably less than unity. This finding portrays the market risk of bank stocks as below average. Other studies develop similar results. For example, Smirlock and Kaufold (1987), relying on 1982 data, estimate market-betas for 23 banks, most of which are below one. They go on to note that, "This ... is consistent with the β's reported for bank portfolios by Flannery and James (1984a, 1984b) and Smirlock (1984)."

Estimates of interest-rate sensitivity prove parallel as well. Bank stock returns are significantly interest-rate sensitive and inversely related to changes in interest rates (which means positively related to holding-period returns). However, on this issue, conflicting evidence may be found in the literature. Chance and Lane (1980) find an insignificant interest-rate coefficient for the two-index model during 1972-76. Sweeney and Warga (1986) report similar results for the same period. In examining variation over the 1960-1979 period, Sweeney and Warga report a lack of interest-rate sensitivity during 1970-1974 and 1975-1979. They find significant interest-rate sensitivity for the 1965-1969 subperiod.

To permit our results to be weighed against 1975-79 estimates by Chance and Lane (1980) and Sweeney and Warga (1986), the fourth panel of Table 1 applies our two-index model to the 1975-1979 period. During this era, our sample banks also show no significant sensitivity to changes in interest rates. Interestingly, the sample shows a market-beta slightly over unity during this subperiod. These comparisons show that when our model specification is applied to our sample banks in periods examined by other studies, similar findings obtain.

We next compare results for the portfolio of our 8 sample S&Ls with an S&L sample studied by Brickley and James (1986). To analyze how access to deposit insurance affects stock returns of S&Ls, Brickley and James estimate shifts in the sensitivity of S&L stock to mortgage bond prices and market returns. They analyze the period 1976-82. To test for a shift in interest-rate elasticity and market beta in January 1980, they fit the following model:

$$R_{pt} = \underset{(0.02)}{0.004} + \underset{(0.287)}{1.41}\ R_{It} - \underset{(0.281)}{0.874}\ R_{It} \cdot D,$$

$$R_{pt} = \underset{(0.001)}{0.002} + \underset{(0.125)}{1.51}\ R_{mt} - \underset{(0.110)}{0.25}\ R_{mt} \cdot D,$$

R_{pt} is the weekly holding-period return on an equally weighted portfolio of 20 S&L stocks, R_{It} is the weekly holding-period return on the GNMA index in week t, and D is a dummy variable that has the value of one except during the calendar year 1979, when it is set at zero. Standard errors are given in parentheses. The slope that is shifting in the first equation relates S&L stock prices to an index of S&L asset returns proxied by the GNMA index. The slope permitted to shift in the second equation is the market-beta of S&L stock.

Estimating these models in the same years with the monthly data of our sample gives:

$$R_{pt} = \underset{(0.01)}{0.009} + \underset{(1.005)}{2.390}\ R_{It} - \underset{(1.056)}{0.794}\ R_{It} \cdot D,$$

$$R_{pt} = \underset{(0.01)}{-0.008} + \underset{(0.613)}{2.184}\ R_{mt} - \underset{(0.668)}{0.132}\ R_{mt} \cdot D.$$

Although our coefficient estimates are qualitatively similar to those of Brickley and James, our negative slope shifts are not statistically significant. The loss of significance may trace to the reduced number of firms in our sample and to the fourfold reduction in observations that occurs in moving from weekly to monthly data.

4. Testing the Parameter Stationarity Hypothesis with Switching Regressions

Table 1 indicates that market and interest-rate sensitivities vary with the subperiod chosen. To run meaningful regressions, one must investigate the possibility of model nonstationarity.

Before drawing substantive inferences from a specific model and data set, researchers should allow for potential inadequacies in measurement, sampling, and specification. Empirical literature in finance supporting the arbitrage-pricing model assures us that generalized market models employed in policy analysis and as benchmarks for event studies are importantly misspecified. Parameters of a misspecified model may be interpreted as coefficients of an incompletely or partially "reduced" form. Estimates of the parameters of a partially reduced form shift with variation in the contributions of various relevant forces whose direct influence is excluded from the model. Researchers using daily or weekly data (for example) must anticipate that, over any time interval of more than a few months, movements in omitted variables (or otherwise misspecified relations) may render the parameters of included variables nonstationary. This potential for nonstationarity imparts unreliability to parameter estimates derived from pooling time-series data that are not put through a battery of prior stationarity tests.

Policy analyses and event studies that neglect nonstationarities caused by variation in omitted variables run a serious risk of sorting out empirical evidence incorrectly. Formal tests of statistical hypotheses may develop either falsely positive or falsely negative results. The likelihood of making false inferences about the influence of included variables grows when omitted variables are correlated with included ones. Similarly, the odds of observing a positive event-study result increases whenever, during or in the vicinity of an event window, reinforcing movements occur in variables for which the model does not control. Conversely, the likelihood of observing negative event-study results increases when offsetting movements in uncontrolled forces (i.e., omitted variables) occur during a study's focal time intervals.

We employ the Goldfeld-Quandt search routine as a way of developing policy-analysis or event-study benchmark models that can incorporate the effects of "relevant" movements in unspecified omitted variables. Although it would be useful either for us or for other researchers to identify these omitted variables and to proceed to endogenize the parameter variation we observe, our goal in this paper is more modest than this. We seek merely to extend the range of experimental control to capture movements in variables that standard pricing models omit. In principle, this extension in ex post control renders the sequence of parameter movements estimated by the switching-regression technique a cleaner benchmark from which to infer the influence of actual or potential policy actions.

The Goldfeld and Quandt [(1972, 1973, 1976)] switching-regression method (GQSRM) provides a flexible way to identify changes in the systematic and unsystematic risks of asset portfolios. The strength of the technique is that the number of effective regimes, the parameter values in each regime, the switch dates at which one regime supersedes another, and the gradualness of each regime switch can all be estimated simultaneously. To explain GQSRM, we specify the multi-regime two-index model as follows:

$$\tilde{R}_{pt} = \beta_j + \beta_{mj}\tilde{R}_{mt} + \beta_{ij}\tilde{R}_i + \tilde{e}_{tj} \qquad , j = 1,\ldots,k,\ldots.r \,. \tag{4}$$

In (4), j is the regime index and e_{tj} is the regression residual distributed as $N(0,\sigma_j^2)$. GQSRM introduces a series of transitional dummy variables, D_{tj}. If the observations come from r regimes, r-1 transitions occur with an equal number of cutoff points. To permit the transitions to occur gradually, r-1 sets of variables D_{tj} may be defined:

$$D_{tj} = \int_{-\infty}^{z_t} [(2\pi)^{1/2}\,\sigma_j^*]^{-1} \exp\left\{-\frac{1}{2}\left(\frac{\xi - Z_j^*}{\sigma_j^*}\right)^2\right\} d\xi \quad , \tag{5}$$

where j now runs from 1 to r-1 and the endpoint values are $D_{tr} = 0$ and $D_{t0} = 1$ by definition.

In (4), the equation representing the k-th regime is then multiplied by $\gamma_{tk} = \prod_{j=0}^{k-1} D_{tj} \prod_{j=k}^{r} (1 - D_{tj})$. The resulting equations for r regimes may be added together to obtain the composite equation that we estimate:

$$\sum_{k=1}^{r} \tilde{R}_{pt}\gamma_{tk} = \sum_{k=1}^{r} \left\{ (\beta_k + \beta_{mk}\tilde{R}_{mt} + \beta_{ik}\tilde{R}_i + \tilde{e}_{tk})(\gamma_{tk}) \right\} . \tag{6}$$

The likelihood function for the r-regime two-index model is obtained by assuming $\tilde{R}_{pt}$ to be normally distributed with mean,

$\mu_{pt} = \sum_{k=1}^{r} \{(\beta_k + \beta_{mk}\bar{R}_{mt} + \beta_{ik}\bar{R}_{it})(\gamma_{tk})\}$, and variance, $\sigma_{pt}^2 = \sum_{k=1}^{r} (\sigma_{tk}^2)(\gamma_{tk})$.

The likelihood function becomes:

$$L = -\frac{T}{2}\log 2\pi - \frac{1}{2}\sum_{t=1}^{T}\sum_{k=1}^{r}(\sigma_{pt}^2) - \frac{1}{2}\sum_{t=1}^{T}\frac{[\sum_{k=1}^{r} R_{pt}\gamma_{tk} - \sum_{k=1}^{r}\{(\mu_{pt}\}]^2}{\sum_{k=1}^{r}(\sigma_{pt}^2)}. \tag{7}$$

Maximizing (7) with respect to its unknown parameters gives maximum-likelihood estimates of: the parameters of the regression relation given in (4); the switch points, Z_j^*; and the gradualness parameters σ_j^* which measure the nonabruptness of the associated structural change. Our objective is to describe the evolution of the systematic market risk (β_{mj}), interest-rate risk (β_{ij}), and unsystematic risk (σ_j) of deposit-institution stock over the 1975-85 period.

Formally, the analysis has two key steps. In the first step, the number of regimes operative during the analysis period is found. Initially, maximum-likelihood values for one-regime (L_1) and two-regime (L_2) models are obtained using (7). The null hypothesis that no regime switch occurs is tested against the alternative that two regimes exist (involving one switch point). Asymptotically, the likelihood-ratio test statistic, -2 ln (L_1/L_2), has a chi-squared distribution with degrees of freedom (d.f.) equal to the number of restrictions under the null hypothesis.[3] If the alternative hypothesis is accepted, we proceed to test the possibility of three regimes.

Then, the null and the alternative hypotheses concern the applicability of two-regime and three-regime processes, respectively. More generally, whenever the null hypothesis is rejected, the number of regimes is incremented and the likelihood-ratio test re-run until we fail to reject the null hypothesis.[4]

The second step is to investigate the parametric nature of these shifts. To illustrate, we may assume k regimes to be established in step one. Our procedure would be to remaximize the likelihood function for k regimes, subject to the parameter-equality constraint under examination, where the maximum-likelihood value is written as L_{kR}. The likelihood-ratio test statistic $-2 \ln(L_{kR}/L_k)$ has asymptotically a chi-squared distribution with one d.f.[5]

Summarizing, we first estimate the number of regimes that the two-index model obeys. Once this number is determined, variation in market sensitivity, interest-rate sensitivity, and residual variance is examined across regimes.

Numerical optimization of the likelihood functions developed in this paper use routines contained in Princeton University's GQOPT package: the NMSIMP (Nelder-Mead Simplex Method) and GRADX (an algorithm using the quadratic hill-climbing method). We use NMSIMP to obtain starting points, which are then used as input into the GRADX to produce parameter estimates and t-values.

5. Have U.S. Deposit Institutions Become Riskier?

For S&Ls and three classes of banks, Table 2 summarizes GQSRM tests of the number of regimes in effect during the 1975-85 period. For each class of institution and each regime, Table 3 presents estimates of the two-index model. Table 3 also notes which parameters prove significant and reports regime-switch dates and measures of the gradualness of each switch. Each class undergoes either three or four shifts in parameter regimes. Most of these shifts prove to be abrupt ones and the switch dates have overlapping confidence intervals. At the same time, the pattern of results indicates that 1975-1985 information flows differentially affected different classes of depository firms.

S&Ls and regional banks show an abrupt first switch in early 1976. (For convenience, we term these first switches the 1976 switch). Money-center and superregional banks' first switch and S&Ls' second switch occur at the beginning of 1977 (the 1977 switch).

All bank groups experience a second switch near or in 1979, (the 1979 switch). In passing, we note that this was a time of marked changes in operative regulatory and monetary policy frameworks. Table 2 shows switch points for money-center, superregional and regional banks of 12/79, 5/79, and 9/78, respectively. The table also provides standard errors for each estimated switch date. Regional banks' second switch is determined to be gradual rather than abrupt. The switch begins 7 months before the mean date of 9/78 and completes itself 7 months after this date.

The third switch occurs for all bank groups in the vicinity of 1982 (the 1982 switch), a year of considerable financial and regulatory turmoil. Except for superregional banks, this switch is determined to be abrupt. Money center, superregional and regional banks experience their third switches in 3/82, 10/81 and 11/82. These point estimates have standard errors of 6 months, 5 months and 8 months, respectively. Using these standard errors produces wide confidence intervals that include late 1981 and all or most of 1982. Even S&Ls' third switch in 9/81 falls within interval estimates of the banks' 1982 switch.

S&L's fourth switch falls in 9/83. Its standard error of 9 months includes late 1982 and early 1984 in its 95 percent confidence interval. The confidence interval for superregional banks' fourth switch is tightly centered on 10/84.

5.1. <u>Changes in Bank Riskiness</u>

Table 4 tests parameter-equality constraints across regimes to identify whether and how each of the three risk parameters varies. We focus first on bank groups noting that interest sensitivity (β_i) varies significantly only for the superregional group. The bulk of the action relates to changes in market-beta (β_m) and unsystematic risk measured as the regression residual variance (σ_r^2).

<u>Variation in market beta</u>. In the 1976 switch, regional banks' market-beta declines dramatically. Market-betas for money-center and the superregional bank groups decrease at the beginning of 1977, but the decline

is not statistically significant for the money-center group. A further statistically significant decrease in market-beta occurs only for money-center and superregional banks in 1979 shift.

In the 1982 shift, every bank class experiences a statistically significant increase in market-beta. In their additional regime shift, superregional banks' slight further increase in market-beta does not prove statistically significant.

The temporal variability of the market sensitivity of bank stock may be described as follows. At the start of our sample period, every bank class had a market-beta in excess of one. In subsequent shifts, market-betas decline broadly until 1982. During 1977-1982, bank market-betas fall below unity. After 1982, market-betas increase back above unity.

These observed changes in the market risk of our very large sample banks help to explain the low levels of market betas reported in Table 1. The time periods examined in Table 1 included eras of below-average market risk for bank stocks. We noted that other studies that bracket the early 1980s develop similar estimates (e.g. Smirlock and Kaufold, 1987). However, Table 3 shows that banks' market beta lies below unity only during the 1977-1982 period. Both before and after this interval, these coefficients lie above unity. Studies which estimate bank systematic risk as if it were stationary improperly pool observations from different regimes. The misleading inferences that develop exemplify the dangers of choosing analysis periods arbitrarily or neglecting the possibility of parameter shifts.

It is natural to hypothesize that the riskiness of bank stocks should vary with changes in bank failure rates and (because of regulatory lags) particularly with the number of banks classified as problem cases by federal examiners. Table 5 shows that failure rates and numbers of problem banks were relatively high in the first two years of our sample and fell back to lower levels between 1977 and 1981. Both series jump sharply in 1982. These data support the GQSRM analysis of temporal variation in market risk.

Variation in other parameters. Unsystematic risk also falls and rises, but for the regionals and superregionals also falls again in the 1980s. In the first switch, all bank groups experience a significant decrease in unsystematic risk. However, as early as the 1979 switch, unsystematic risk begins to rise. During the 1979-82 period of generally high unsystematic risk, market risk and bank failure rates remain relatively low.

As noted earlier, little significant temporal variation occurs in bank interest sensitivity. Positive coefficients occur for all bank groups during all but two of the regimes delineated. This implies a conventional inverse relationship between bank stock returns and market interest rates. No bank group shows significant interest-rate sensitivity before the 1979 switch. Only during 1979-1982 volatile interest rate era do all bank groups show significant interest-rate sensitivity. Our observed great variation in interest-sensitivity can account for the seemingly inconsistent conclusions of empirical investigations using smaller time periods and different size compositions of sample banks. Differences in results may trace to differences in the character of and market environments operative in these authors' samples. Once again, we see the importance of allowing for possible nonstationarities before drawing inferences from regression results.

5.2. A Look at Survivorship Bias in the Money-Center Group

Our sample includes no officially failed banks. However, the Bank of America and Continental Illinois may be described as quasi-failures. As a sensitivity experiment, we exclude Continental Illinois and Bank of America from the money-center group to examine whether these banks distort findings for this group. Table 6 reports regressions for this edited subsample and for Continental Illinois taken by itself. Comparing Table 3 and 6 shows that money-center banks' switch points do not differ significantly between the two subsamples and that risk parameters behave similarly. On the other hand, taking Continental by itself shows some substantial differences. Continental skipped the 1977 shift and developed a negative market beta in 1979. Its 1982 increases in beta move against those for other money-center banks. The April 1984 shift coincides with an announced increase of $400 million in its problem loans, starting the run which led to its de facto nationalization. In its 4/84 shift, all risk components increase, but with only the increase in unsystematic risk proving statistically significant. Continental's federal rescue appears to destroy the two-index model's applicability, inasmuch as neither beta attains significance in the post-rescue regime. Interestingly, in its prefailure period Continental shows no significant interest sensitivity. As our S&L sample, this may indicate that its deposit insurer was absorbing the bulk of interest induced gains and losses dring this troubled time. The plausibility of this interpretation is also supported by the failure of the parameter regime for other money-center banks to shift in the wake of the Continental crisis.

5.3 Changes in S&L Riskiness

For S&Ls, the last column of Table 4 tests for parameter inequality across regimes. As do regional banks, S&Ls experience significant shifts in market and unsystematic risk in early 1976, developing a significantly negative market-beta and greatly enhanced interest sensitivity. However, S&Ls show an abrupt second shift in 4/77, one combining increases in unsystematic risk and market-beta with reduced interest sensitivity.

In contrast to the 1979 shift for bank groups, the return-generating process for our small sample of S&Ls shows no switch during the 1979-1980 era. S&Ls undergo their third shift in late 1981, with market and unsystematic risk doubling and interest sensitivity declining to insignificance.

In the late-1983 shift, market risk declines and interest sensitivity rises (both significantly), while the fall in unsystematic risk fails to achieve statistical significance.

As we did for banks, we may compare temporal variation in market and unsystematic risk with fluctuations in problem and failed institutions. Unfortunately, time series of failures and problem S&Ls are not routinely published by the Federal Home Loan Bank Board. For 1975-1985, Table 5 reports, as a proxy for problem S&Ls, the number of S&Ls whose net worth is less than or equal to zero under Generally Accepted Accounting Principles (GAAP). We call these institutions "GAAP-insolvent S&Ls."

Table 5 shows that the number of failed S&Ls surged sharply in 1982, with a parallel rise in GAAP-insolvent institutions. Although the number of failed S&Ls declines significantly in 1984, a second sharp surge in the number of GAAP-insolvent S&Ls is observed. This suggests that the reduced failure rate reflects FSLIC's own growing economic insolvency and staffing weakness and not an across-the board improvement in industry credibility. In S&Ls' fourth regime, market and unsystematic risk reach peaks for the 1975-1985 period. This peak is consistent with S&L failure rates but not with the trend in GAAP insolvency. With only 8 extremely large S&Ls in our sample (and most of these headquartered in California), it is doubtful that the regression shifts we observe are representative of the S&L industry as a whole.

Interest-rate sensitivity varies over a far wider range of values for S&Ls than for banks. Moreover, three out of the four shifts in S&L interest-rate sensitivity prove significant. During their first regime and third regime when market-betas were at high levels, S&L interest sensitivity is not significantly different from zero. In the remaining three regimes, S&L stock proves significantly market and interest-rate sensitive. During the 9/81-9/83 regime, although interest rates were highly volatile, S&L stock shows near-zero interest sensitivity. This may trace not only to a higher incidence of adjustable-rate mortgages and to mortgage prepayments in the last part of this era, but also to the extent of hidden economic insolvency at sample S&Ls. Deep insolvencies could have forced the FSLIC to absorb the bulk of interest induced profits and losses on short-funded positions in long assets during this particularly troubled era.

5.4 Some Observed Implications

The observed pattern of coefficient change indicates that the riskiness of sample bank and S&L stocks declined in the late 1970s, but rose again in recent years. Of course our sample of banks and S&Ls is not necessarily representative of banks and S&Ls as a whole. The sample includes only very large institutions. Moreover, using monthly observations reduces the power of our tests. The wide confidence intervals that attach to the 1979 and 1982 switch dates are particularly unfortunate. Macroeconomic and regulatory events taking place in these eras call out for tighter estimates. Daily or weekly data might let us develop closer estimates of the timing of the structural changes in return-generating processes.

This paper does not attempt to link its statistical findings with particular information events. Although the switches could be shown to coincide with various information flows, without analyzing a full chronology of potential information events and considering potentially relevant omitted variables, asserting such links would overreach the regression experiments we conduct.

A growing literature has sought to explain cross-sectional differences in bank stocks' interest-sensitivity (e.g. Flannery and James 1984a, Brewer and Lee 1985, and Tarhan 1987). These studies all assume interest-sensitivity is stationary over the periods they analyze. However, because our findings indicate that interest sensitivity shifted importantly over these authors' analysis periods, such nonstationarity needs to be taken into account. Modelling this nonstationarity endogenously over time would amount to developing a model of the repricing process for deposit-institution stock.[6]

6. Reassessing Event Studies

Several authors use 1975-85 data to investigate the impact of selected information events on bank stocks. We focus on papers concerned with events occuring in 1979 (1979 event studies) or in 1982 (1982 event studies). Our estimated switch dates establish a perspective from which to reinterpret these studies.

6.1 1979 Event Studies

Aharony, Saunders, and Swary (1986) examine the impact on commercial banks of the well-publicized October 6, 1979 change in Federal Reserve operating procedures. Using weekly data and a two-index model, they postulate the possibility of a parameter shift on this date. To test for risk changes, they estimate market risk, interest-rate risk, and unsystematic risk for the two years bracketing the October 1979 announcement date. For every bank group they examine, tests of risk-parameter equality uncover no significant change in market or interest-rate risk across the two years.

Brickley and James (1986) investigate how modification of insolvency rules by the FSLIC in 1980 might have affected common-stock returns for financial institutions. Using weekly returns, and taking January 1980 as the date that insolvency rules were relaxed, these authors test the hypothesis that market risk and interest sensitivity prove lower in the 1980-1982 period than in the 1976-1979 period. For S&Ls but not for banks, they find significant decreases in these sensitivity measures.

In the finance literature, it is commonplace to investigate the risk impact of regulatory changes in this fashion. Among others, Benston (1973), Aharony, Jones and Swary, Aharony and Swary (1981), Smirlock (1984), and Lamy and Thompson (1986) employ similar methods. But two problems inhere in this approach (Unal 1987). First, precise dates for regime switches are imposed a priori and never tested. Second, possibilities for parameter change are restricted to a narrow subset of regulatory and legislative "event dates." Event-study researchers neither identify nor control for the effects of other information events that might develop at nearby dates. For example, while Aharony _et al_. (1986) associate their regression findings with the Federal Reserve's October, 1979 announcement, Brickley and James link the same class of regression results to a January, 1980 action by FSLIC. Failure to impose statistical controls for other information events increases the probability of falsely accepting or rejecting the null hypothesis.

Our method does find switches for every bank group in 1979. However, because the switch-point estimates have large standard errors, we cannot tightly bracket the switch dates on the Fed's October, 1979 announcement. 95 percent confidence intervals of the switch dates for all bank groups span much of 1979. We observe that after these shifts bank groups experience significant interest-rate sensitivity, lower market risk and higher unsystematic risk. It is striking that the return-generating process for our small sample of S&Ls shows no switch in the 1978-1980 era.

A hallmark of the late 1970s is the many ways in which banking and S&L regulators sought to help their regulatees to reduce burdens that deposit-rate ceilings would otherwise have imposed. For example, Money Market Certificates were authorized in May, 1978. In this same month the Fed authorized commercial banks to offer automatic transfer service (ATS) accounts. A lawsuit challenging the authority of the Fed and other regulators to authorize this and various other forms of implicit interest was denied by a District Court in October 1978, but in April 1979 the Court of Appeals for that District reversed this finding. Congress was saddled with a yearend deadline either to authorize ATS accounts and other popular regulatory innovations or to see them lapse. After voting itself a three-month extension, on March 31,1980 Congress legalized the challenged regulatory innovations and set up a six-year phase-out of deposit-rate ceilings on time and savings accounts.

Presuming a sharp causal connection between specific monetary or regulatory events and shifts in regression parameters goes beyond the inferential reach of the data actually examined. Clearly, given the ebb and flow of regulatory, judicial, and legislative events, it is rash to associate parameter shifts (nonshifts) for commercial-banks or S&Ls in 1978-1979 with specific dates. Who is to say which regulatory actions, monetary-policy changes, modification of insolvency rules, court decisions, or steps in the passage of the 1980 legislation, competitive developments, were more important than other events? A cautious observer can only say that the 1978-79 period includes numerous developments that might jointly or individually have supported expectations revisions large enough to induce a shift (or nonshift) in the return-generating processes for deposit-institution stocks.

6.2 1982 Event Studies

To put the specific events analyzed by 1982 event studies into broad perspective, it is instructive to consider a partial chronology of information events that might have proved relevant in 1982. Early in 1982, nonperforming loans increased sharply at large banks and the SEC authorized shelf-registration. Shelf-registration promised to make it easier for large firms to issue open-market securities as an alternative to bank loans. In March a Joint Congressional Resolution held it to be the "sense of Congress" that the full faith and credit of the U.S. Treasury stood behind federal deposit insurance. Other worrisome events in early 1982 include: Drysdale Government Securities' slide into bankruptcy by mid-May, the developing Penn Square crisis, and the LDC debt crisis. The office of the Comptroller of the Currency closed the Penn Square Bank on July 5. In mid-August, Mexico declared a moratorium on its foreign debt. However, negative information about the value of Latin America debt arrived throughout 1982. In late summer or early fall the Fed is said to have readopted a policy of interest-rate smoothing. In November, Congress passed the Garn-St Germain Act, aimed at forestalling the collapse of the S&L industry. The Act also authorized Money Market Deposit Accounts and Super NOW accounts. These instruments hit the market in December, when the Fed also reduced its discount rate to the lowest level in four years. Completing the picture, we recall from Table 5 that 1982 is a year when the number of failed and problem institutions surges sharply.

Thus, 1982 shapes up as a crucial year for the deposit-institution industry. Despite these many and variegated events, Lamy and Thompson (1986) treat the Penn-Square crisis as the dominant event in this time interval, using the market model to estimate shifts in market risk and unsystematic risk for the 100 trading days preceding and 25 trading days following July 6, 1982. They find no significant shift in market risk between these periods, but a significant shift in unsystematic risk. They use this as "evidence of a structural change in the pricing mechanism for bank stocks after the Penn Square failure."

Mexico's moratorium announcement in August, 1982 is the only development we know to spawn four independent event studies [Smirlock and Kaufold 1985, Schoder and Vankudre 1986, Cornell and Shapiro 1986, and Bruner and Simms, 1987]. Broadly, these studies analyze the response of bank stock prices on or around the announcement-day for the moratorium. In focusing on announcement-day "abnormal returns" or parameter shifts, these authors do not control for the possibility of parameter nonstationarity elsewhere in 1982. The flow of information events throughout 1982 provides reason to believe that deposit-institution riskiness may be fluctuating. Estimating abnormal returns without allowing for alternative shift dates or nonstationarities may lead a researcher falsely to accept or to reject event-study hypotheses.

Also focusing on 1982, Fraser, Richards and Fosberg (1985) seek to determine the direction of the impact of Super NOWs on bank shareholder wealth. They use a two-index model to estimate abnormal returns, but without controlling for changes in slopes or residual variance. They report that "..while money center banks were essentially unaffected by the announcement of Super NOWs, (excess returns for) regional banks were strongly (and negatively) affected." However, the significant shift our methods find in the market-beta for regional banks explains the negative abnormal returns these authors report as the consequence of an abrupt market repricing necessary to let these stocks offer appropriate ex ante compensation for their now-heightened market risk.

Our statistical findings reflect the financial and regulatory turmoil of 1982. S&Ls and every bank class we examine experience switches in or near 1982 that show a significant increase in market risk. Money-center banks and S&Ls also show increased unsystematic risk. The estimated shift is a gradual one for superregional banks. For the other bank classes, because observed switch dates have large standard errors, almost any date in 1982 could pass muster as a potential switching point. The overriding problem is that so many potential events can be identified in 1982 that it is presumptuous to label any one of them as the precise cause of the nonstationarity our methods uncover.

7. Summary

Two types of evidence are developed in this paper: substantive and methodological. Substantively, we show that our sample of deposit-institution stocks became riskier investments in the wake of the many regulatory relaxations made in the 1980s. 1979 and 1982 are affirmed as years when information events substantially affected return-generating processes for deposit-institution stock. Bank market risk lies below unity only during the 1977-1982 period. Both before and after this interval, bank market risk lies above unity. Our methods find bank equity returns to be interest-sensitive primarily during the 1979-82 era, but S&L equity returns to be interest-sensitive during the bulk of the observation period.

Methodologically, we show that using GQSRM to identify nonstationarities in deposit-institution equity returns supports the hypothesis that information flows have differentially affected different types and sizes of institutions. These same nonstationarities underscore the unreliability of traditional event-study methods. It is inappropriate to designate a potential shift date without also allowing for the effects of other information events or controlling for observable nonstationarities in returns that occur in the neighborhood of a researcher's event window.

NOTES

1. Flannery and James (1984a) and Brewer and Lee (1985) focus on the relation between banks' interest-rate sensitivity and their balance-sheet composition. This leads them to test the interest-rate sensitivity of the institutions included in their sample. In a related study, Martin and Keown (1977) investigate the importance of extra-market sources of covariation. They find significant covariance among the unsystematic-risk parameters for financial institutions and suggest that this may reflect an interest-rate factor.

2. Although very few institutions of the size sampled on CRSP tapes disappeared over the sample period (e.g., Penn Square Bank and Franklin National Bank), a survivorship bias is built into this approach.

3. Goldfeld and Quandt (1973, p 479) indicate that this statistic "appears in finite samples to be well approximated" by the Chi-squared distribution.

4. The conditionality of higher-round tests on the outcome of previous rounds of testing means that we should tighten the test criterion to maintain a fixed level of significance. Our procedure may be said to fix a maximal number of regimes that might be operative during the analysis period.

5. Again, although this statistic tests the parameter-equality hypothesis, some small-sample bias may exist.

6. A related study (Unal and Kane, 1987) constructs and estimates a model of this repricing process.

Table 1: Estimates of Market and Interest Rate Sensitivity for Bank Portfolios Reported in the Literature and Comparison of These Results with Kane and Unal (KU) Sample

		β_m	β_I	Analysis Period	Observations	No. of Banks
1.	FJ	0.56 (18.5)	0.13 (3.5)	1/76-11/81	Weekly	68
	KU	0.68 (7.38)	0.33 (3.11)	1/76-11/81	Monthly	31
2.	SP	0.67 (8.15)	-0.40 (-4.41)	1/77-12/84	Monthly	78
	KU	0.75 (8.10)	0.35 (3.48)	1/77-12/84	Monthly	31
3.	BL	0.53 (80.83)	-0.02 (6.71)	1/78-6/84	Daily	44
	KU	0.74 (7.03)	0.36 (3.23)	1/78-6/84	Monthly	31
4.	KU	1.05 (9.66)	0.27** (1.30)	1/75-12/79	Monthly	31

1. Flannery and James, 1984 (FJ), use NYSE Composite Index as the market return. They use three interest-rate indices. The interest-rate coefficient reported here is estimated using the residuals of an AR(3) model for the weekly holding-period return for GNMA 8 percent certificates.
2. Scott and Peterson, 1986 (SP), use the S&P500 return index and monthly percentage changes in 30-year Treasury bond yields as proxies for the market index and the interest-rate index, respectively.
3. Brewer and Lee, 1986 (BL), use the value-weighted NYSE and AMEX composite index obtained from CRSP to proxy market returns. Their proxy for the interest-rate index is the difference between the 3-month Treasury bill rate at time t and the forward 3-month Treasury bill rate imbedded in the yield curve at time t-1. The reported estimates are obtained from cross-section time-series data.
4. The analysis period is chosen to coincide with that of Chance and Lane (1980) and Sweeney and Warga (1986).

Note: t-values are given in parentheses and are significant at the 1 percent level unless marked with a double asterisk (**).

Table 2: Likelihood-Ratio Test to Determine the Number of Regimes in Effect During the 1975-85 period by institution class.

Institution/Regime	(L*/L)	-2 ℓN(L*/L)
1. Money Center Banks		
R_1 vs. R_2	201.56/211.52	19.92
R_2 vs. R_3	211.52/222.43	21.82
R_3 vs. R_4	222.43/235.67	26.48
R_4 vs. R_5	235.67/238.10	4.86*
2. Superregional Banks		
R_1 vs. R_2	251.12/259.83	17.42
R_2 vs. R_3	259.83/269.18	18.70
R_3 vs. R_4	269.18/277.64	16.92
R_4 vs. R_5	277.64/285.82	16.36
R_5 vs. R_6	285.82/288.42	5.20*
3. Regional Banks		
R_1 vs. R_2	240.54/247.32	13.56
R_2 vs. R_3	247.32/255.16	15.68
R_3 vs. R_4	255.16/261.93	13.54
R_4 vs. R_5	261.93/267.03	10.20*
4. Savings and Loan Associations		
R_1 vs. R_2	161.09/168.66	15.14
R_2 vs. R_3	168.66/176.41	15.50
R_3 vs. R_4	176.41/184.09	15.36
R_4 vs. R_5	184.09/192.29	16.40
R_5 vs. R_6	192.29/195.47	6.36*

Notes: Critical value for 6 d.f. at 5 percent significance is 12.592. (*) indicates that the hypothesis of an additional regime is rejected at 5 percent significance.

Table 3: Maximum-Likelihood Estimates of Regime Parameters for the 1975-1985 period

Parameters	Money Center Banks	Superregional Banks	Regional Banks	S&Ls
Starting Date	1/75	1/75	1/75	1/75
β_{m1}	1.40*	1.36*	1.33*	1.84*
β_{11}	.20	.33	.43	1.35
σ_1	.06	.04	.05	.09
Switch Point	25.20 (.98)	24.75 (.57)	13.76 (.42)	14.00 (.03)
Implied Date	1/77	1/77	2/76	2/76
Switch Std. Dev.	.65	.12	.25	.01
β_{m2}	1.03*	.92*	.43*	-1.12*
β_{12}	.35	-.26	.35	3.02*
σ_2	.02	.02	.02	.03
Switch Point	60.43 (3.11)	53.19 (1.47)	45.19 (5.93)	28.15 (.9)
Implied Date	12/79	5/79	9/78	4/77
Switch Std. Dev.	1.21	.15	7.00*	.02
β_{m3}	.07	.45*	.57*	1.17*
β_{13}	.45*	.54*	.27*	1.27*
σ_3	.04	.04	.04	.05
Switch Point	86.53 (6.28)	81.62 (4.98)	95.34 (8.22)	81.30 (.9)
Implied Date	3/82	10/81	11/82	9/81
Switch Std. Dev.	2.1	2.8	.06	.1
β_{m4}	1.35*	1.19*	1.03*	2.67*
β_{14}	.44	.16	.38	.43
σ_4	.06	.04	.03	.10
Switch Point	-	118.12 (.1)	-	105.3 (9.1
Implied Date	-	10/84	-	9/83
Switch Std. Dev.	-	.59	-	1.59
$\beta m5$	-	1.37*	-	1.01*
β_{15}	-	-.24	-	1.36*
σ_5	-	.01	-	.06
Ending Date	12/85	12/85	12/85	12/85

Notes: (*) indicates significantly different from zero at the 5 percent level The standard error of each switch point is given in parentheses. All intercep estimates are small and fail to be statistically significant at 5 percent.

Table 4: Likelihood-Ratio Test of Shifts in Risk Parameters During the 1975-85 period.

Parameter Restriction	Money Center Banks	Superregional Banks	Regional Banks	S&Ls
	1977 Switch		1976 Switch	
$\beta_{m2}-\beta_{m1}$	1.68	3.94*	7.70*	15.58*
$\beta_{i2}-\beta_{i1}$	.38	1.34	.12	1.42
$\sigma_2-\sigma_1$	27.50*	23.58*	9.56*	16.82*
	1979 Switch			1977 Switch
$\beta_{m3}-\beta_{m2}$	20.78*	6.16*	.58	15.58*
$\beta_{i3}-\beta_{i2}$	.22	7.98*	.44	6.22*
$\sigma_3-\sigma_2$	8.60*	19.68*	13.76*	5.00*
	1982 Switch			
$\beta_{m4}-\beta_{m3}$	18.44*	10.34*	5.36*	6.40*
$\beta_{i4}-\beta_{i3}$	.04	2.36	.32	2.56*
$\sigma_4-\sigma_3$	6.40*	.04	1.04	17.26*
		1984 Switch		1983 Switch
$\beta_{m5}-\beta_{m4}$	-	1.08		6.46*
$\beta_{i5}-\beta_{i4}$	-	3.94*		4.02*
$\sigma_5-\sigma_4$	-	5.10*		.92

Notes: The test statistic is $-2\,\ln(L^*/L)$, where L^* and L are restricted and unrestricted maximum likelihood values. Critical values for 1 d.f. at 5 percent and 10 percent significance are 3.84 and 2.71, respectively. (*) indicates that the coefficient shift is significant at 5 percent. No shift in intercept proves significant at 5 percent.

Table 5: Number of Problem Banks, Failed Banks and S&Ls, and GAAP insolvent S&Ls.

Year	Failed Banks	Problem Banks	Failed S&Ls	GAAP-Insolvent S&Ls
1970	8	251	...	...
1971	6	239	...	...
1972	3	190	...	...
1973	6	155	...	...
1974	4	181	...	...
1975	14	347	11	17
1976	17	385	9	11
1977	6	368	9	8
1978	7	342	3	4
1979	10	287	3	13
1980	10	217	35	17
1981	10	223	81	65
1982	42	369	252	201
1983	48	642	101	287
1984	79	847	42	434
1985	120	1140	70	450

Sources: Information on banks is compiled from Kane (1985) and the Federal Deposit Insurance Corporation's 1985 Annual Report. S&L information comes from FHLBB files and Barth, Brumbaugh and Sauerhaft (1986).
Notes: "Failure" is defined as a regulator-induced cessation of autonomous operations. It includes supervisory mergers or acquisitions and loose forms of conservatorship such as the Federal Home Loan Bank Board's Management Consignment Program. "Problem banks" are those that are classified as such by FDIC examiners. "GAAP-Insolvent S&Ls" is defined as those S&Ls whose net worth is less than or equal to zero under Generally Accepted Accounting Principles (GAAP).

Table 6: Maximum-Likelihood Estimates of Regime Parameters for Money-Center Banks (excluding Continental Illinois and Bank of America) and Continental Illinois for the 1975-1985 period

Parameters	Money Center Banks (excluding Continental)	Continental Illinois
Starting Date	1/75	1/75
β_{m1}	1.48*	1.26*
β_{i1}	-.18	.36
σ_1	.06	.07
Switch Point	25.41 (.76)	59.01 (.08)
Implied Date	1/77	11/79
Switch Std. Dev.	.62	.07
β_{m2}	.96*	-.48
β_{i2}	.52*	.80*
σ_2	.02	.05
Switch Point	62.42 (2.88)	82.14 (4.23)
Implied Date	2/80	10/81
Switch Std. Dev.	3.28	1.97
β_{m3}	.12	1.73*
β_{i3}	.39*	-.41
σ_3	.04	.08
Switch Point	86.49 (6.21)	111.95 (4.44)
Implied Date	3/82	4/84
Switch S.D.	.53	.09
β_{m4}	1.25*	2.54
β_{i4}	.48	1.44
σ_4	.05	.27
Ending Date	12/85	12/85

Notes: (*) indicates significantly different from zero at the 5 percent level. The standard error of each switch point is given in parentheses.

Appendix 1: List of Sample Banks and Savings and Loan Associations

Institution	Asset Size Million $. End of 1984
Money Center Banks	
1. BankAmerica Corp	117680
2. Bankers Trust of New York Corp.	45208
3. Chase Manhattan Corp.	86883
4. Chemical Bank of New York Corp.	52236
5. Citicorp	150586
6. Continental Bank of Illinois Corp.	30413
7. First Chicago Corp.	39846
8. Manufacturers Hanover Corp.	75713
9. J.P. Morgan and Co. Inc.	64126
Superregional banks	
1. Bank of Boston	22079
2. Bank of New York, Inc.	15156
3. First City Bancorporation Texas	17318
4. First Interstate Bancorp.	45544
5. Interfirst Corp.	21617
6. Irving Bank Corp.	18982
7. Marine Midland Banks, Inc.	22056
8. NBD Bancorp, Inc.	14232
9. Norwest Corp.	21346
10. Republic New York Corp.	12382
11. Southeast Banking Corp.	9869
12. Texas Commerce Bancshares, Inc.	20732
13. Wells Fargo and Co.	28184
Regional Banks	
1. Bank of Virginia	4134
2. Equimark Corp.	2593
3. First Pennsylvania Corp.	5355
4. First Virginia Banks, Inc.	2686
5. First Wisconsin Corp.	5516
6. Fleet Financial Group, Inc.	5747
7. General Bancshares Corp.	2024
8. United Jersey Banks Hackensack	4050
9. Wachovia Corp.	8717
Savings and Loan Associations	
1. Ahmanson H.F. & Co.	24307
2. Far West Financial Corp.	2050
3. Financial Corporation of America	28518
4. Gibraltar Financial Corp.	9273
5. Golden West Financial Corp. DFL	10620
6. Great Western Financial Corp.	23555
7. Imperial Corporation of America	8465
8. TransOhio Financial Corp.	2797

Note: Asset sizes are from consolidated holding-company balance sheets a shown in Moody's Bank and Finance Manual, 1984 and 1985.

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