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RISK-SHIFTING BY FEDERALLY INSURED COMMERCIAL BANKS

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ABSTRACT

Mispriced and misadministered deposit insurance imparts risk-shifting incentives to U.S. banks. Regulators are expected to monitor and discipline increases in bank risk exposure that would transfer wealth from the FDIC to bank stockholders. This paper assesses the success regulators had in controlling risk-shifting by U.S. banks during 1985-1994. In contrast to single-equation estimates developed from the option model by others, our simultaneous-equation evidence indicates that regulators failed to prevent large U.S. banks from shifting risk to the FDIC. Moreover, at the margin, banks that are undercapitalized shifted risk more effectively than other sample banks.

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Risk-Shifting by Federally Insured Commercial Banks

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Until a risk-related premium structure was introduced in the 1990s, the flat explicit deposit insurance premium paid by U.S. banks created an incentive for insured banks to increase their risk exposure[1]. Viewing the bank's deposit guarantee as a put option written by the FDIC, Merton (1977) shows that the value of this option to the bank would increase with asset risk and leverage risk. Higher risk increases the value of the FDIC's insurance services. If the incremental cost a bank pays for these services does not rise to absorb this value, increasing risk exposure extracts an incremental deposit insurance subsidy. Such subsidies may benefit bank stockholders via higher stock price and/or they may be "shifted", in whole or large part, to bank customers.

Four forces act to restrain excessive risk taking by the banks. Benston, Eisenbeis, Horvitz, Kane, and Kaufman (1986) argue that market discipline from uninsured debtholders and other stakeholders limit risk-taking excesses. They also mention a second limiting factor: managerial risk-aversion. To avoid the negative effect that leading a bank to failure may have on their reputational human capital, managers may limit risk-taking to reduce probability of distress. Saunders et al. (1990) develop evidence that supports this hypothesis. Third, though cushioned by deposit insurance, distress costs help to restrain risk-taking. Substantial charter value may be lost in the

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[1] See, for example, Buser et al. (1981), Goodman and Santomero (1986), and Kane (1986, 1987).

event of liquidation[2]. The last but not least important restraining factor is government regulation[3].

Buser et al. (1981) model the price of federal deposit insurance as a sum of two components: an explicit insurance premium and an implicit cost associated with regulatory pressure for safe and sound operation. The implicit premium is conceived to be risk-sensitive so as to stop excessive risk-taking by the banks. In a bilaterally market-driven equilibrium, the marginal benefits from deposit insurance would be exactly offset by its marginal costs (explicit and implicit). However, deposit insurance is a trilateral contract. Agency problems in government decision-making process raise the possibility of an equilibrium in which an uncompensated risk exposure is shifted to taxpayers.

Gorton and Rosen (1995) opine that, despite general agreement among researchers on bank incentives, "empirical research has not reached a consensus on whether deposit insurance is underpriced." Marcus and Shaked (1984) adapt a one-year put option model to estimate the risk-adjusted deposit insurance premiums. Their results suggest that the FDIC insurance is generally overpriced, although the distribution of "fair" premiums is highly skewed. Ronn and Verma (1986) refine the Marcus-Shaked model and incorporate market perceptions of FDIC forbearance policies. Duan, Moreau, and Sealey (1992) use the methods of Ronn and Verma (1986) to estimate fair FDIC premiums. Linking changes in these premiums to the changes in banks' risk exposure, they estimate that successful risk-shifting is a relatively isolated phenomenon.

These studies presume a one-year option settlement framework, which understates the value of deposit insurance. Pennacchi (1987) clarifies that an implicit and counterfactual

[2] See Buser et al. (1981) and Marcus (1984).
[3] See, for example, Buser et al (1981), Flannery (1982), and Pyle (1984, 1986).

assumption in one-year models is that, at each settlement date, insurance premia are adjusted to fair value and/or the FDIC forces banks to raise their capital ratios to a fair level. Studies by Allen and Saunders (1993) and Cooperstein, Pennacchi, and Redburn (1995) propose more realistic multiperiod deposit insurance valuation methods.

Even maintaining the counterfactual assumption of prompt option settlement, we develop strong evidence of active risk-shifting. We use the model of Duan, Moreau, and Sealey (1992) to tie the incremental insurance subsidy received by a bank to changes in the bank's risk exposure. But our paper improves the empirical analysis of risk-shifting in three ways:

(1) Our sample is substantially larger and is more representative of the population of U.S. banks than the samples studied in prior work.

(2) We explore differences in risk shifting between healthy and troubled institutions. There is no reason to expect a uniform risk-shifting pattern across our sample. Corporate finance theory implies that risk-shifting incentives are strongest for troubled institutions. The very large banks studied in the previous literature include some of the safest in U.S.

(3) We use an improved estimator that recognizes and adjusts for simultaneity bias built by definition into single-equation structural models of bank leverage and risk-adjusted insurance premia.

Our estimates indicate that risk-restraining forces are far weaker than earlier studies suggest. The policy implication of our evidence is that ailing institutions seek more actively to shift risks than other banks do and that in 1985-1994 the regulatory system failed to contain risk-shifting by these banks.

The paper is organized as follows. Section I reviews the one-period option pricing model for risk-adjusting deposit insurance. Our sample, methods and results are described in section II. Section III examines the robustness of our findings. Section IV sums up the paper.

I. The Model

Ronn and Verma (1986) model the fair premium for deposit insurance services (IP) as the value of a limited-term (one year) put option:

$$IP = B_1 N(y_2) - \frac{(1-\delta)^n V B_1}{B} N(y_1), \qquad (1)$$

where B_1 is the face value of insured deposits, B is the face value of insured and uninsured debt, V is the value of bank assets, δ is the dividend per dollar of assets, n is the number of times per period the dividend is paid. Variables y_1 and y_2 are defined as follows:

$$y_1 = \frac{\ln[B / V(1-\delta)^n] - \sigma_V^2 T / 2}{\sigma_V \sqrt{T}},$$

$$y_2 = y_1 + \sigma_V \sqrt{T},$$

where σ_V is the instantaneous standard deviation of asset returns, N(•) is the cumulative standard normal distribution, T is the effective maturity (one year).

It is convenient to estimate IPP, the insurance premium per dollar of insured deposits:

$$IPP = IP / B_1 = N(y_2) - (1-\delta)^n \frac{V}{B} N(y_1). \qquad (2)$$

Neither the market value of bank assets, V, nor the standard deviation of asset returns, σ_V, is observable. They can, however, be estimated using two additional equations[4]. The first equation states the call option valuation assignable to the equity of a levered firm:

$$E = VN(x_1) - \rho BN(x_2), \tag{3}$$

where E is the value of the bank's equity, ρ is the exercise price, as a percentage of the value of deposits.

Variables x_1 and x_2 are defined as follows:

$$x_1 = \frac{\ln[V / \rho B] + \sigma_V^2 T / 2}{\sigma_V \sqrt{T}},$$

$$x_2 = x_1 - \sigma_V \sqrt{T}.$$

We follow Ronn and Verma (1986) in modeling FDIC forbearance policy as lowering the exercise price of the shareholders' call option below the value of bank deposits by setting ρ = 0.97. This procedure neglects the option value of further FDIC forbearance when the ρ threshold is crossed (Kane (1986)).

The second equation states a relationship between σ_V and σ_E that follows from Ito's lemma:

$$\sigma_V = \sigma_E \frac{E}{V} \frac{1}{N(x_1)}. \tag{4}$$

Here, σ_E is the instantaneous standard deviation of equity returns.

Comparative-static analysis of equation (2) implies that a manager can increase the value of the deposit insurance subsidy by increasing asset risk σ_V and/or leverage B/V.

[4] Equations (3) and (4) are solved simultaneously for V and σ_V using Microsoft Excell 5.0 software.

Duan et al. (1992) translate this implication into two testable hypotheses by approximating the change in the per-dollar insurance premium as follows:

$$\Delta IPP \cong \frac{\partial IPP}{\partial \sigma_V}\Delta\sigma_V + \frac{\partial IPP}{\partial (B/V)}\frac{\partial (B/V)}{\partial \sigma_V}\Delta\sigma_V . \tag{5}$$

Equation (5) is rewritten as:

$$\Delta IPP \cong \beta_1 \Delta\sigma_V , \tag{6}$$

where

$$\beta_1 \equiv \frac{\partial IPP}{\partial \sigma_V} + \frac{\partial IPP}{\partial (B/V)}\alpha_1 , \tag{7}$$

$$\alpha_1 \equiv \frac{d(B/V)}{d\sigma_V} . \tag{8}$$

A positive β_1 would indicate that the bank was successful in increasing its deposit insurance subsidy through risk-shifting during the sample period. Merton (1977) shows that -- absent market and government disciplinary responses -- the partial derivatives $\partial IPP/\partial\sigma_v$ and $\partial IPP/\partial(B/V)$ are positive. Empirically, risk-sensitive capital regulation and other risk-restraining factors introduce negative linkages between σ_V and B/V (negative α_1). The net effect of how disciplinary restraints modify risk-shifting incentives can be assessed by estimating α_1 and β_1. The two hypotheses about the character of net risk-shifting are:

Hypothesis 1: $\alpha_1 \geq 0$,

Hypothesis 2: $\beta_1 \leq 0$.

Treating σ_V as exogenous (a restriction we relax in section III), one can test Hypotheses 1 and 2 for an individual bank j by estimating the following equations:

$$\frac{B_{jt}}{V_{jt}} = \alpha_{0j} + \alpha_{1j}\sigma_{V_{jt}} + \varepsilon_{jt}. \tag{9}$$

$$IPP_{jt} = \beta_{0j} + \beta_{1j}\sigma_{Vjt} + \xi_{jt}. \tag{10}$$

Rejecting Hypothesis 1 would affirm that risk-restraining factors outweigh risk-shifting incentives. Rejecting Hypothesis 2 would indicate that risk-shifting incentives prevail.

Partial derivatives of IPP with respect to σ_V and B/V are complicated functions of these variables. There is no reason to expect that the parameter β defined in (7) is a constant. Even if the majority of banks operate safely and do not actively shift risks, incentives for risk shifting intensify with financial distress. Marcus (1984) shows that healthy banks may prefer a safer policy to protect their charter, while financially weaker banks may reap advantages from strategic risk-taking. Marcus and Shaked (1984) find the distribution of bank risk-shifting to be skewed: almost the entire estimated value of the aggregate deposit insurance subsidy flows to the riskiest 5 percent of insured banks. The hypothesis that it may pay a bank to shift risks aggressively when it is in distress implies that positive α_1 and β_1 should emerge more frequently at distressed institutions.

To test for risk-shifting by distressed institutions, we reestimate equations (9) and (10) on pooled data introducing interactive dummy variables for financial distress:

$$\frac{B_{jt}}{V_{jt}} = \alpha_{0j} + \alpha_1{}^1 d\sigma_{V_{jt}} + \alpha_1{}^0(1-d)\sigma_{V_{jt}} + \varepsilon_{jt}. \tag{11}$$

$$IPP_{jt} = \beta_{0j} + \beta_1{}^1 d\sigma_{Vjt} + \beta_1{}^0(1-d)\sigma_{Vjt} + \xi_{jt}. \tag{12}$$

In these equations, d = 1 if the bank was undercapitalized in the preceding quarter and d=0 otherwise. We augment the analysis with the following two null hypotheses that we expect to reject:

Hypothesis 3: $\alpha_1^1 \leq 0$

Hypothesis 4: $\beta_1^1 \leq 0$

Rejecting these two hypotheses would indicate that undercapitalized banks are able to expand their access to deposit insurance subsidy.

II. Methods and Results

A. Sample Selection and Data

The models derived in the previous section are estimated with quarterly data running from the first quarter of 1985 through the fourth quarter of 1994. Our analysis focuses on chartered commercial banks (SIC codes 6021 and 6022) whose shares trade on NYSE, AMEX, or NASDAQ. To be included into our sample, a bank must have at least sixteen consecutive quarters of balance-sheet and stock-returns data available from, correspondingly, Compustat Industrial Quarterly and CRSP daily returns files. Such a restriction is necessary to allow estimation of regressions for each individual bank. These screening criteria are satisfied for 124 banks, offering us a total of 4,227 bank-quarter observations. Survivorship bias may be introduced if our selection criteria eliminate some banks whose particularly intense risk-shifting incentives resulted in early failure. This bias makes our assessment of risk-shifting by troubled banks conservative.

Table I presents the sample distribution by number of quarters of data available and by chartering authority. A full forty quarters of data are available for eighty-three banks. Data for the remaining forty-one bank cover sixteen to thirty-eight quarters. Forty-seven banks are state-chartered; eighty-seven have a national charter.

Table II reports the distribution of the sample by calendar years. Observations by calendar year vary from a low of 338 in 1985 to a high of 492 in 1991, 1992,and 1993.

Table III presents summary statistics for the sample. The sample banks are diverse in their financial characteristics. While the sample is clearly skewed toward larger institutions, relatively small banks are represented better than in earlier studies. Bank asset size varies from a mere $125 million to $251 billion. Other characteristics, such as standard deviations of stock and asset returns, as well as the estimated values of fair deposit insurance premiums are of similar magnitude to those in the samples analyzed by Marcus and Shaked (1984), and Ronn and Verma (1986). The last line in the table shows that while the majority of the sample shows very low values for risk-adjusted insurance premiums (the median is 0.003%), a distinct minority receives extremely high IPP values. This reinforces our concern that analyzing average risk-shifting for the sample is not sufficient. It is important to look for groups of banks that actively pursue deposit insurance subsidies.

B. Definition of Variables

Statistical proxies for conceptual variables are briefly discussed below:

B, *total debt* : calculated as a difference between book values of assets (Compustat item 044) and common equity (Compustat item 059).

E, *market value of bank's equity* : calculated as the end-of-period stock price (Compustat item 014) times number of shares outstanding (Compustat item 061).

σ_E , *standard deviation of the return on equity* : estimated from daily stock returns over the previous quarter[5].

δ , *dividends per dollar of assets* : calculated as the cash dividends per common share (Compustat item 020) times number of shares outstanding (Compustat item 061) divided over market value of bank's assets (V).

C. Tests of Hypothesis 1

C.1. Linear Time-Series Estimates

To test Hypothesis 1, equation (9) is estimated separately for each of 124 banks. The regressions include corrections for first-order autocorrelation in residuals. Results are reported in the first column of Table IV. At the five percent significance level, 33 negative and no positive estimates of α_1 emerge. At the ten percent level, nine more coefficient estimates prove significantly negative, but only one significantly positive coefficient develops. This evidence rejects Hypothesis 1 that the risk restraints are negligible only for between one-quarter and one-third of the sample banks.

Duan et al. (1992) report that 28 of the 30 banks they examined for 1976-1986 exhibited a significantly negative α_1. This led them to claim that restraints on bank risk-shifting dominate risk-taking incentives. In our expanded sample that covers a later time period, risk-restraining factors prove much weaker. Twenty-one banks are present in both samples. For 20 of these banks Duan et al. (1992) report significantly negative estimates

[5] We require at least 20 stock-return observations for a particular bank-quarter to be included into our sample. The estimated standard deviation of stock returns is annualized assuming 253 trading days per year.

of α_1. At five percent, only 4 of these show negative coefficient estimates in our 1985-1994 time period, and only two more negative coefficients emerge at ten percent.

C.2. Nonlinear Least-Squares Estimates

In regression (9), the left-hand side variable σ_V, is calculated from variables that include the dependent variable B/V. This implies that estimates of α_1 from model (9) are biased by variable definitions. To clarify this, we note from (3) and (4) that:

$$\sigma_V = \sigma_E - \sigma_E \rho \frac{B}{V} \frac{N(x_2)}{N(x_1)}. \tag{13}$$

Therefore, (9) can be rewritten as

$$\frac{B}{V} = \alpha_0 + \alpha_1 \sigma_E \left[1 - \rho \frac{B}{V} \frac{N(x_2)}{N(x_1)} \right] + \varepsilon. \tag{14}$$

With B/V on both sides of equation (14) direct estimates of α_1 suffer from a definitional simultaneity bias. By construction, the endogenous variable B/V enters the regressor. This renders the regressor stochastic and correlated with the error term ε.

To remove this bias from estimates of α_1, equation (14) can be rearranged into a reduced form by collecting B/V terms on the left-hand side[6]:

$$\frac{B}{V} = \frac{\alpha_0}{1 + \alpha_1 \sigma_E \rho N(x_2) / N(x_1)} + \frac{\alpha_1 \sigma_E}{1 + \alpha_1 \sigma_E \rho N(x_2) / N(x_1)} + \zeta. \tag{15}$$

Column 2 of Table IV, presents estimates of α_1 for each bank using the method of non-linear least squares. The regressions include a correction for first-order autocorrelation in residuals. Using five percent significance, across the 124 banks, equation (15) yields

[6] Strictly speaking, B/V still enters the right hand side of equation (15) through the term $N(x_2)/N(x_1)$. As a practical matter, however, $N(x_2)/N(x_1)$ remains virtually constant across the sample observations: 99% of all sample values fall between 0.99 and 1. The mean value of the ratio is 0.9995, its standard deviation is 0.0031, and the correlation with B/V is only -0.058.

twelve positive and five negative estimates of α_1. Thirteen additional estimates become significantly positive at the ten percent significance. Thus, using the more appropriate nonlinear model (15), only about four percent of the sample banks behave as if they are subject to strong risk-restraining factors. The policy implication is that risk-restricting factors were largely neutralized during the 1985 - 1994 period.

For 38 banks, models (9) and (15) yield different results for the sign of α_1. The probability of a discrepancy at least as large as this by pure chance[7] is only 4.3×10^{-18}. It seems clear that, by not recognizing definitional simultaneity built into equation (9), one is bound to overestimate the strength of risk-restraining factors.

C.3. Pooled Estimates

We next estimate equation (11) for a panel of all banks. To resolve the simultaneity problem identified in the previous subsection, we employ lagged values of the independent variable σ_v as instruments. At the cost of introducing specification error, this eliminates contemporaneous correlation between the regressor and the disturbance vector and allows us to obtain a consistent estimator of α_1.

We experimented with several specifications for the pooled regression. Likelihood-ratio tests indicated the existence of both firm- and time-specific effects, while Hausman tests supported the fixed-effects model against the random-effects model. Using the fixed-effects model allows intercepts to vary across sample banks and periods while holding slopes constant. Results are reported in Table V. The estimated α_1 in this pooled

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[7] Assuming that the probability of type I error is 5 percent and that these errors are independent, the probability of having 38 type I errors is equal $(0.05)^{38}\times124!/[38!\times(124-38)!]$.

regression proves significantly negative, indicating that, on balance across the sample, risk-restraining factors are strong.

We test Hypothesis 3 that risk-restraining factors differ across banks by using interactive dummy variables. We estimate α_1 separately for two groups of banks: those whose leverage in the preceding quarter lay below (d = 0) and above (d = 1) the 90th sample percentile. The results reject Hypothesis 3: the significantly negative coefficient on the interactive variable for d = 0 contrasts with the positive coefficient when d = 1. This finding indicates that restraining forces are weak for banks that become undercapitalized. This implies that regulatory and market disciplines break down in the very cases where they need to be strong.

D. Tests of Hypothesis 2

D.1. Time-Series Estimates

Table VI reports tests of Hypothesis 2 developed from estimates of equation (10) for each bank in the sample. The regressions include a correction for first-order autocorrelation in residuals. At five percent significance, five positive estimates of β_1 emerge and no negative ones. At ten percent, another estimate becomes significantly positive and four negative ones emerge. This indicates that about five percent of sample banks were able to expand deposit insurance subsidies through risk-shifting. Although this percentage is small, the positive coefficients found in these time-series regressions imply that particular banks shifted risk for a prolonged period of time: in fact throughout most of the 4 to 10-year period for which data on these banks were observed.

D.2. Pooled Estimates

Relationship (12) is estimated for a panel of all banks. Again, alternative specifications are compared. Likelihood-ratio and Hausman tests suggest the presence of bank- and time-specific intercepts. The slope is held constant across banks. Results are presented in Panel A of Table VII. The estimated β_1 in this pooled regression differs insignificantly from zero, suggesting that, on average, risk-restraining factors just offset risk-shifting incentives.

To test Hypothesis 4 about how risk-restraining factors differ across banks, we use interactive dummy variables to estimate β_1 separately for two groups of banks: those whose leverage in the preceding quarter lay below ($d = 0$) and above ($d = 1$) the 90th sample percentile. The results reject Hypothesis 4 at the ten percent level: a significantly positive beta emerges from the interactive variable when $d = 1$. This indicates that regulatory discipline failed to prevent undercapitalized banks from increasing their deposit insurance subsidy. For well-capitalized banks, the coefficient is insignificant.

III. Robustness of Results

In this section we examine the sensitivity of our findings to misspecification and other statistical problems. Overall, our sensitivity tests support the hypothesis of risk-shifting and indicate that risk-restraining factors in 1985-1994 were weak.

A. *Simultaneity Bias and Errors-in-Variables*

In principle, a bank's leverage and the volatility of its asset returns are simultaneously determined. Neglecting the simultaneity between these variables could bias inferences about risk restraints. Bias could also exist because the unobservable explanatory variable is inevitably measured with error.

An imperfect way to deal with simultaneity bias and errors-in-variables problem is to use instrumental-variable (IV) estimators. A candidate for an instrument must be correlated with the true regressor, but contemporaneously uncorrelated with the error term. A workable (but potentially unreliable) approach is to substitute the predetermined (lagged) values for the imperfectly observed regressor. Because we correct for first-order autocorrelation in residuals by quasi-differencing, we introduce the second lag of the explanatory variable as our instrument.

We have already used IV estimation for the pooled regressions relating bank leverage ratios to asset risk, σ_V. We now apply the same approach to the pooled regressions of risk-adjusted deposit insurance premiums. The results reported in Panel B of Table 7 strengthen our earlier conclusions: undercapitalized institutions successfully increase the value of their deposit insurance guarantee, but well-capitalized institutions do not and may indeed not want to.

The IV versions of time-series tests of hypotheses 1 and 2 portray risk-restraining factors as even weaker than our more-restrictive tests did. At 5 percent, the relation between leverage and asset risk proves significantly negative for only three banks. The number rises only to 10 banks at ten percent. The relation between deposit insurance premium, IPP, and asset risk is significantly positive for 12 banks (15 at ten percent).

Although a bank can change its risk exposure overnight, using predetermined values requires us to look at relations between leverage and asset risk that are measured half a year apart. These IV-model results support the hypothesis that regulatory controls on risk shifting operate with substantial lags.

B. Ratio Specification of the Model

It is also desirable to investigate the effect of coefficient restraints that are built into every ratio specification. The models we estimated earlier place implicit restrictions on the coefficient estimates. For example, equation (9) for the regressand B/V is mathematically equivalent to:

$$B = \alpha_0 V + \alpha_1 V \sigma_V + V\varepsilon . \tag{16}$$

An unrestricted version of equation (16) introduces two potentially nonzero terms (with coefficients γ_0 and γ_2):

$$B = \gamma_0 + \gamma_1 V + \gamma_2 \sigma_V + \gamma_3 V \sigma_V + \nu . \tag{17}$$

To test for the appropriateness of restricting $\gamma_0 = 0$ and $\gamma_2 = 0$ one may estimate equation (17).

For 34 banks, the coefficient restrictions imposed by (9) are not supported by the data. The number of banks with strong risk-restricting factors ($\gamma_3 < 0$) decreases from 33 in the restricted single-equation regression (15) to only 12. The number of banks with a positive relation between leverage and asset risk increases from none to seven.

Similarly, we estimate an unrestricted version of equation (10). The restrictions are not supported for 21 banks. The number of risk-shifting banks increases from five to nine. Again, these results strengthen our initial findings.

IV. Summary and Conclusions

U.S. experience with state deposit-insurance systems (Calomiris, 1992) and corporate finance theory hold that weaknesses in deposit insurance pricing and capital enforcement encourage banks to extract deposit insurance subsidies by increasing their risk exposure. To test this hypothesis, we decompose it into three parts.

The first part looks to measure the strength of regulatory enforcement as an intensifying risk-restraining factor. We remove definitional simultaneity built in the model used by Duan et al. (1992) to eliminate a bias in previous single-equation estimates of the tested parameters. In contrast to the biased procedure, our tests indicate that risk-restraining factors are weak.

The second part of our tests investigates the frequency of risk shifting among sample banks. Our results indicate that about five percent of the sampled banks succeeded in increasing their deposit insurance subsidy consistently over observation periods of four to ten years.

A third set of tests focuses on endogenous risk-shifting by banks in periods for which benefits from increasing deposit insurance subsidy promise to be extremely high. Consistent with finance theory, we affirm that risk-shifting behavior is concentrated at the financially weakest banks.

An array of sensitivity tests establish the robustness of our principal finding: troubled U.S. banks actively engaged in risk-shifting during 1985-1994 and the regulatory system failed to restrain this behavior adequately.

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Table I

Sample Distribution by Number of Quarters of Data Available and by Chartering Authority.

Number of Quarters of data	Number of banks: all banks	federally chartered	state chartered
40	83	57	26
38	1	1	0
36	3	2	1
33	1	0	1
32	3	1	2
31	1	0	1
28	1	1	0
27	1	0	1
25	1	1	0
24	2	1	1
23	1	0	1
22	1	1	0
21	2	0	2
20	4	1	3
17	2	2	0
16	17	9	8
Number of banks	124	77	47
Number of observations	4,227	2,728	1,499

Table II

Sample Distribution by Calendar Year.

Year	Number of observations
1985	338
1986	353
1987	372
1988	381
1989	396
1990	421
1991	492
1992	492
1993	492
1994	490
Total	4,227

Table III

Summary Statistics

Summary statistics for 124 sample banks over the period 1985-94. The statistics are calculated from quarterly data. Leverage ratio is defined as (face value of total debt / market value of assets).

Characteristic	Mean	Median	Min	Max
Market value of assets ($MM)	18,350	6,856	125	251,000
Total deposits ($MM)	17,448	6,344	120	242,700
Leverage ratio (%)	93.5	93.9	65.2	103.2
Annualized std. dev. of rate of return on equity (%)	28.3	25.15	5.3	128.8
Annualized std. dev. of rate of return on assets (%)	2.5	2.2	0.4	26.7
Risk-adjusted deposit insurance premium per dollar of deposits, IPP (%)	0.101	0.003	0.000	4.882

Table IV

The Relation between Bank Leverage and Asset Risk: Time-Series Tests

Time series tests of Hypothesis 1: $\alpha_1 \geq 0$. The α_1 is the slope coefficient in the following structural model:

$$B_{jt} / V_{jt} = \alpha_{0j} + \alpha_{1j}\sigma_{Vjt} + \varepsilon_{jt} . \quad (9)$$

where B_{jt} / V_{jt} is the leverage, σ_{Vjt} is the annualized standard deviation of asset returns for bank j in quarter t.

The table below reports two estimates of α_1: a direct estimate from linear model (9) and an estimate from the following reduced form nonlinear regression:

$$B_{jt} / V_{jt} = (\alpha_{0j} + \alpha_{1j}\sigma_{Ejt}) / [1 + \alpha_{1j}\sigma_{Ejt}\,\rho N(x_{1jt}) / N(x_{2jt})] + \zeta_{jt} . \quad (15)$$

where σ_{Ejt} is the annualized standard deviation of equity returns for bank j in quarter t, $\rho = 0.97$, $N(\bullet)$ is the standard normal c.d.f., and x_1 and x_2 are defined as follows:

$$x_1 = \frac{\ln[V / \rho B] + \sigma_V^2 T / 2}{\sigma_V \sqrt{T}}, \qquad x_2 = x_1 - \sigma_V \sqrt{T}$$

	Linear model		Nonlinear model	
	α_1	t-stat.	α_1	t-stat.
AMSOUTH BANCORPORATION	-0.250	-1.17	0.134	0.62
BANC ONE CORP	-0.412*	-1.97	0.053	0.24
BANCORP HAWAII INC	-0.309	-0.82	0.761*	1.70
BANK OF BOSTON CORP[1]	0.195	1.25	0.509**	2.33
BANK OF NEW YORK CO INC[1]	-0.088	-0.26	0.043	0.03
BANK SOUTH CORP	-0.671***	-4.21	-0.146	-0.89
BANKAMERICA CORP[1]	-0.233	-0.89	0.276	1.33
BANKERS TRUST NEW YORK CORP[1]	-0.137	-0.64	0.265	1.14
BANPONCE CORP	-0.078	-0.85	0.005	0.06
BARNETT BANKS INC[1]	-0.403	-1.49	0.532	1.52
BAYBANKS INC	-0.402	-1.29	0.528	1.31
BOATMENS BANCSHARES INC	0.498	1.59	0.987***	2.71
CENTRAL FIDELITY BANKS INC	-0.465*	-1.74	0.285	0.91
CHASE MANHATTAN CORP[1]	-0.563*	-1.90	0.468	1.38
CHEMICAL BANKING CORP[1]	-0.657**	-2.34	0.543	1.39
CITICORP[1]	0.131	0.65	0.462*	1.84
CITY NATIONAL CORP	-0.244	-0.91	0.686*	1.89
COMERICA INC	-0.453*	-1.74	0.267	0.76
COMMERCE BANCSHARES INC	0.438	1.32	0.810**	2.41
COMPASS BANCSHARES INC	0.349*	1.88	0.579***	2.64
CORESTATES FINANCIAL CORP	-0.225	-0.80	0.397	1.32
CRESTAR FINANCIAL CORP	-0.170	-0.42	1.248**	2.10
CULLEN/FROST BANKERS INC	-0.217	-0.75	0.249	0.82
DAUPHIN DEPOSIT CORP	-0.414**	-2.27	-0.085	-0.47
DEPOSIT GUARANTY CORP	-0.174	-0.48	0.600	1.62
ELDORADO BANCORP/CA	-0.309***	-3.48	-0.195**	-2.25
FIFTH THIRD BANCORP	-0.025	-0.09	0.557*	1.81
FIRST AMERICAN CORP/TN	0.018	0.06	0.585*	1.78
FIRST OF AMERICA BANK CORP	-0.322	-0.96	0.350	1.07
FIRST BANK SYSTEM INC	-0.339	-1.68	0.033	0.31
FIRST CHICAGO CORP[1]	-0.010	-0.03	0.692*	1.98

Table IV continued

	Linear model		Nonlinear model	
	α_1	t-stat.	α_1	t-stat.
FIRST COMMERCE CORP	-0.561	-1.42	1.074	1.52
FIRST EMPIRE STATE CORP	0.055	0.37	0.164	1.12
FIRST FID BANCORPORATION[1]	0.067	0.26	0.565**	2.05
FIRST HAWAIIAN INC	-0.474**	-2.18	0.100	0.42
FIRST INTERSTATE BNCP[1]	-0.385	-1.56	0.265	1.02
FIRST SECURITY CORP/DE	-0.830***	-3.85	0.014	0.05
FIRST TENNESSEE NATL CORP	-0.495*	-1.81	0.191	0.70
FIRST UNION CORP (N C)	-0.019	-0.08	0.566*	1.70
FIRST VIRGINIA BANKS INC[1]	0.091	0.30	0.849**	2.18
FIRSTAR CORP[1]	-0.624**	-2.12	0.271	0.84
FLEET FINANCIAL GROUP INC	-0.208	-1.10	0.223	0.94
FOURTH FINANCIAL CORP	-0.652**	-2.30	0.154	0.44
HUBCO INC	-0.417**	-2.46	0.076	0.33
HIBERNIA CORP	-0.455*	-1.77	0.610	1.45
HUNTINGTON BANCSHARES	-0.337	-1.33	0.362	1.11
INTEGRA FINANCIAL CORP	-0.763**	-2.21	0.470	0.98
KEYCORP	0.104	0.54	0.449*	1.95
MARSHALL & ILSLEY CORP	-0.532**	-2.38	-0.091	-0.38
MELLON BANK CORP	-0.325	-0.64	0.664	1.07
MERCANTILE BANCORPORATION	-0.250	-1.13	0.155	0.85
MERCANTILE BANKSHARES CORP	-1.196***	-3.40	0.373	0.45
MERIDIAN BANCORP INC	-0.278	-1.22	0.219	0.94
MICHIGAN NATIONAL CORP	-0.133	-0.41	0.364	1.21
MIDLANTIC CORP	-0.132	-0.45	0.692	1.52
MORGAN (J P) & CO[1]	0.120	1.08	0.222	1.68
NBD BANCORP INC[1]	-0.394	-1.66	0.115	0.51
NATIONAL CITY CORP	0.026	0.10	0.453	1.52
NATIONSBANK CORP[1]	-0.566*	-1.69	0.977*	1.77
NORTHERN TRUST CORP	-0.470	-1.30	0.547	1.51
NORWEST CORP[1]	-0.084	-0.37	0.282	1.55
OLD KENT FINANCIAL CORP	0.025	0.10	0.402	1.68
PNC BANK CORP	-0.354	-1.17	0.626	1.64
PREMIER BANCORP	-0.533**	-2.19	0.388	1.03
REGIONS FINL CORP	-0.006	-0.02	0.427	1.51
REPUBLIC NEW YORK CORP[1]	-0.263	-1.20	0.095	0.41
RIGGS NATL CORP WASH D C	-0.447**	-2.10	0.423	1.13
SHAWMUT NATIONAL CORP	-0.413	-1.56	0.500	1.35
SIGNET BANKING CORP[1]	-0.060	-0.20	0.647	1.50
SOUTHTRUST CORP	-0.232	-0.93	0.377	1.27
STAR BANC CORP	-0.052	-0.34	0.173	0.96
STATE STREET BOSTON CORP	0.138	0.57	0.718**	2.24
TRUSTMARK CORP	-0.006	-0.05	0.193	1.42
UJB FINL CORP[1]	-0.617***	-2.88	0.101	0.35
UMB FINANCIAL CORP	-0.291	-1.29	0.203	0.78
UNION PLANTERS CORP	-0.951**	-2.24	0.879	1.32
U S BANCORP	-0.356	-1.42	0.332	1.18
U S TRUST CORP	-0.414	-1.50	0.269	0.70
WELLS FARGO & CO[1]	-0.563***	-2.64	0.034	0.20
WEST ONE BANCORP	-0.412	-1.67	0.225	0.86

Table IV continued

	Linear model		Nonlinear model	
	α_1	t-stat.	α_1	t-stat.
WESTAMERICA BANCORP	-0.337**	-2.20	-0.010	-0.07
WILMINGTON TRUST CORP	-0.085	-0.27	0.860**	2.07
ZIONS BANCORPORATION	-0.506	-1.61	0.698	1.61
SUNTRUST BANKS INC	-0.483	-1.60	0.307	0.93
FIRSTIER FINANCIAL INC	-0.591***	-4.71	-5.379***	-42.32
SUMMIT BANCORPORATION	0.219	0.95	0.813*	1.95
WACHOVIA CORP	0.081	0.32	0.453*	1.73
FIRST CITIZENS BANCSH	0.254	0.84	1.065**	2.06
MID AMERICA BANCORP/KY	0.066	0.27	0.725*	1.85
NORTH FORK BANCORPORATION	0.231	0.88	1.125*	1.92
SYNOVUS FINANCIAL CP	-0.156	-1.19	0.222	1.02
SANTA MONICA BK CALIF	-0.077	-0.32	0.339	0.64
SOUTHERN NATIONAL CORP	-0.320	-1.21	0.311	0.97
GUARDIAN BANCORP	-0.095	-0.59	0.874	1.00
LIBERTY BANCORP INC/OK	-0.064	-0.25	0.229	0.97
IMPERIAL BANCORP	-0.385	-0.97	1.637	1.48
MAGNA GROUP INC	0.069	0.14	1.505**	2.22
UNION BANK SAN FRANCISCO	-0.734	-1.53	1.803	1.60
INTERCHANGE FINL SVCS CP/NJ	-0.798**	-2.26	0.008	0.02
CVB FINANCIAL CORP	-1.238***	-5.03	-0.914***	-3.80
METROBANK	-0.780***	-3.77	-0.566**	-2.56
CITIZENS BANCORP/MD	-1.014***	-3.09	2.652	1.46
FIRST MIDWEST BANCORP INC	-0.781	-1.41	1.354	1.12
PROFESSIONAL BANCORP INC/PA	-0.759***	-3.65	-0.376	-1.50
TRUST CO NJ JERSEY CITY	-0.506*	-1.78	0.116	0.30
CENTURY BANKS INC	-1.549***	-3.84	-0.661	-1.48
FIRST MARYLAND BANCORP	0.026	0.18	0.106	0.72
ASSOCIATED BANC CORP	-1.038***	-2.87	-0.436	-1.08
CCB FINANCIAL CORP	-0.793**	-2.05	-0.233	-0.46
CHARTER BANCSHARES INC	-0.510**	-2.58	-0.154	-0.75
CITIZENS BANKING CORP	-1.021***	-5.25	-0.710***	-2.80
FIRST COMMERCIAL CORP	-0.522**	-2.65	-0.257	-1.08
FIRST MICHIGAN BANK CORP	-0.677	-0.99	1.538	1.03
FULTON FINANCIAL CORP	-0.299	-0.66	0.495	0.93
MBNA CORP	0.084	0.30	0.383	1.38
MARK TWAIN BANCSHARES	-1.115***	-4.12	-0.438	-1.70
NATIONAL BANCORP ALASKA INC	-0.098	-0.27	0.515	1.00
NATIONAL COMM BANCORP/TN	-1.247	-1.45	3.884	1.27
PROVIDENT BANCORP	-0.875**	-2.07	0.145	0.21
SC BANCORP/CA	-1.154***	-4.12	-0.173	-0.27
SUSQUEHANNA BANCSHARES INC	-0.063	-0.24	0.226	0.84
UST CORP	-0.444	-1.65	-0.098	-0.54
UNITED CAROLINA BANCSHARES	-0.644	-1.77	-0.155	-0.41
VALLEY NATIONAL BANCORP	-1.316	-6.06	1.332**	2.11

* significant at the 10 percent level using one-tailed test.
** significant at the 5 percent level using one-tailed test.
*** significant at the 1 percent level using one-tailed test.
[1] each of these banks were part of the sample in Duan et al. (1992).

Table V

The Relation between Bank Leverage and Asset Risk:

Pooled Time-Series-Cross-Sectional Tests

Pooled regression tests of Hypotheses 1 and 3 using the fixed-effects model with bank- and time-specific intercepts. σ_{Vjt-2} is used as the instrumental variable for σ_{Vjt}.
Hypothesis 1: $\alpha_1 \geq 0$. The α_1 is the slope coefficient in the following regression:

$$B_{jt} / V_{jt} = \alpha_{0j} + \alpha_{1j}\sigma_{Vjt} + \varepsilon_{jt} . \quad (9)$$

where B_{jt} / V_{jt} is the leverage, σ_{Vjt} is the annualized standard deviation of asset returns for bank j in quarter t.
Hypothesis 3: $\alpha_1^{1} \leq 0$. The α_1^{1} is the first slope coefficient in the following regression:

$$B_{jt} / V_{jt} = \alpha_{0jt} + \alpha_1^{1} d\sigma_{Vjt} + \alpha_1^{0}(1\text{-}d)\sigma_{Vjt} + \varepsilon_{jt} , \quad (11)$$

where d = 1 for undercapitalized institutions and 0 otherwise.

	α_1	t-stat.	α_1^{1}	t-stat.	α_1^{0}	t-stat.
Without dummies	-0.693***	-23.43				
With dummies			0.285***	3.82	-0.749 ***	-23.49

*** significant at the 1percent level using one-tailed test.

Table VI

The Relation between Fair Deposit Insurance Premium and Asset Risk: Time-Series Tests

Time series tests of Hypothesis 2: $\beta_1 \leq 0$. The β_1 is the slope coefficient in the following regression:

$$IPP_{jt} = \beta_{0j} + \beta_{1j}\sigma_{Vjt} + \xi_{jt}. \quad (10)$$

where IPP_{jt} is the risk-adjusted insurance premium per dollar of insured deposits, σ_{Vjt} is the annualized standard deviation of asset returns for bank j in quarter t.

	β_1	t-stat.
AMSOUTH BANCORPORATION	0.056***	4.22
BANC ONE CORP	-0.003	-0.38
BANCORP HAWAII INC	0.005	0.56
BANK OF BOSTON CORP	-0.041	-0.57
BANK OF NEW YORK CO INC	0.123*	1.76
BANK SOUTH CORP	0.001	0.08
BANKAMERICA CORP	0.117	1.31
BANKERS TRUST NEW YORK CORP	0.002	0.05
BANPONCE CORP	0.020	0.49
BARNETT BANKS INC	-0.003	-0.14
BAYBANKS INC	-0.055	-1.37
BOATMENS BANCSHARES INC	0.012	0.41
CENTRAL FIDELITY BANKS INC	0.142	1.36
CHASE MANHATTAN CORP	0.030	1.62
CHEMICAL BANKING CORP	-0.086	-1.27
CITICORP	0.049	0.43
CITY NATIONAL CORP	-0.018	-0.41
COMERICA INC	0.091	1.01
COMMERCE BANCSHARES INC	-0.025*	-1.91
COMPASS BANCSHARES INC	0.024	0.93
CORESTATES FINANCIAL CORP	0.004	0.73
CRESTAR FINANCIAL CORP	-0.003	-1.13
CULLEN/FROST BANKERS INC	-0.036	-0.99
DAUPHIN DEPOSIT CORP	-0.004	-0.70
DEPOSIT GUARANTY CORP	-0.028	-0.71
ELDORADO BANCORP/CA	0.079***	3.44
FIFTH THIRD BANCORP	-0.001	-0.92
FIRST AMERICAN CORP/TN	-0.009	-0.49
FIRST OF AMERICA BANK CORP	-0.113	-1.08
FIRST BANK SYSTEM INC	0.001	1.07
FIRST CHICAGO CORP	-0.069	-0.69
FIRST COMMERCE CORP	-0.006	-0.70
FIRST EMPIRE STATE CORP	-0.005	-0.38
FIRST FID BANCORPORATION	-0.010	-0.12
FIRST HAWAIIAN INC	-0.007	-0.33
FIRST INTERSTATE BNCP	-0.027	-0.92
FIRST SECURITY CORP/DE	-0.029	-1.03
FIRST TENNESSEE NATL CORP	0.073	1.65
FIRST UNION CORP (N C)	0.004	0.21
FIRST VIRGINIA BANKS INC	0.024	0.52

Table VI continued

	β_1	t-stat.
FIRSTAR CORP	-0.053	-1.42
FLEET FINANCIAL GROUP INC	-0.011	-0.63
FOURTH FINANCIAL CORP	0.001	0.18
HUBCO INC	0.008	1.20
HIBERNIA CORP	0.000	0.04
HUNTINGTON BANCSHARES	0.021***	3.99
INTEGRA FINANCIAL CORP	-0.010	-1.20
KEYCORP	-0.011	-0.53
MARSHALL & ILSLEY CORP	-0.001	-0.08
MELLON BANK CORP	0.001	0.09
MERCANTILE BANCORPORATION	-0.099	-0.39
MERCANTILE BANKSHARES CORP	0.021	1.30
MERIDIAN BANCORP INC	-0.099	-1.11
MICHIGAN NATIONAL CORP	0.045	0.80
MIDLANTIC CORP	-0.008	-0.53
MORGAN (J P) & CO	-0.006	-0.78
NBD BANCORP INC	-0.020	-0.23
NATIONAL CITY CORP	-0.002	-0.14
NATIONSBANK CORP	0.025	1.32
NORTHERN TRUST CORP	-0.002	-0.23
NORWEST CORP	-0.087*	-1.99
OLD KENT FINANCIAL CORP	-0.021	-1.00
PNC BANK CORP	-0.011	-0.95
PREMIER BANCORP	0.006	0.18
REGIONS FINL CORP	0.039	1.08
REPUBLIC NEW YORK CORP	-0.040	-1.04
RIGGS NATL CORP WASH D C	0.027	0.48
SHAWMUT NATIONAL CORP	-0.165	-1.43
SIGNET BANKING CORP	-0.010	-1.38
SOUTHTRUST CORP	0.007	0.75
STAR BANC CORP	0.001	0.19
STATE STREET BOSTON CORP	-0.022	-0.94
TRUSTMARK CORP	0.024	0.71
UJB FINL CORP	-0.013	-0.70
UMB FINANCIAL CORP	0.025**	2.07
UNION PLANTERS CORP	-0.001	-0.12
U S BANCORP	-0.047	-1.51
U S TRUST CORP	0.025	0.91
WELLS FARGO & CO	-0.014	-0.62
WEST ONE BANCORP	0.007	0.93
WESTAMERICA BANCORPORATION	0.006	0.45
WILMINGTON TRUST CORP	-0.077*	-1.79
ZIONS BANCORPORATION	-0.043	-1.13
SUNTRUST BANKS INC	-0.121*	-1.96
FIRSTIER FINANCIAL INC	0.018	0.44
SUMMIT BANCORPORATION	0.018	1.41
WACHOVIA CORP	-0.004	-0.67
FIRST CITIZENS BANCSH	-0.041	-1.24
MID AMERICA BANCORP/KY	-0.019	-1.13
NORTH FORK BANCORPORATION	-0.033	-0.73

Table VI continued

	β_1	t-stat.
SYNOVUS FINANCIAL CP	-0.000	-0.17
SANTA MONICA BK CALIF	0.001	0.37
SOUTHERN NATIONAL CORP	-0.269	-1.11
GUARDIAN BANCORP	-0.009	-0.89
LIBERTY BANCORP INC/OK	-0.058	-0.65
IMPERIAL BANCORP	-0.073	-1.67
MAGNA GROUP INC	-0.023	-0.20
UNION BANK SAN FRANCISCO	0.004	0.19
INTERCHANGE FINL SVCS CP/NJ	-0.000	-0.60
CVB FINANCIAL CORP	0.160***	4.56
METROBANK	-0.030	-0.55
CITIZENS BANCORP/MD	-0.022	-0.85
FIRST MIDWEST BANCORP INC	0.011	0.34
PROFESSIONAL BANCORP INC/PA	-0.006	-1.30
TRUST CO NJ JERSEY CITY	0.002	0.06
CENTURY BANKS INC	-0.006	-1.24
FIRST MARYLAND BANCORP	-0.004	-0.31
ASSOCIATED BANC CORP	-0.001	-1.06
CCB FINANCIAL CORP	-0.035	-0.39
CHARTER BANCSHARES INC	-0.000	-0.07
CITIZENS BANKING CORP	-0.001	-1.17
FIRST COMMERCIAL CORP	-0.008	-1.77
FIRST MICHIGAN BANK CORP	-0.065	-1.16
FULTON FINANCIAL CORP	0.013	0.61
MBNA CORP	0.005	0.95
MARK TWAIN BANCSHARES	0.009	0.70
NATIONAL BANCORP ALASKA INC	-0.039	-1.17
NATIONAL COMM BANCORP/TN	0.075	1.46
PROVIDENT BANCORP	-0.000	-0.10
SC BANCORP/CA	0.005	0.24
SUSQUEHANNA BANCSHARES INC	-0.000	-0.54
UST CORP	-0.000	-0.17
UNITED CAROLINA BANCSHARES	0.081	1.03
VALLEY NATIONAL BANCORP	0.001	0.46

* significant at the 10 percent level using one-tailed test.
** significant at the 5 percent level using one-tailed test.
*** significant at the 1 percent level using one-tailed test.

Table VII

The Relation between Fair Deposit Insurance Premium and Asset Risk: Pooled Time-Series-Cross-Sectional Tests

Hypothesis 2: $\beta_1 \geq 0$. The β_1 is the slope coefficient in the following regression:

$$IPP_{jt} = \beta_{0jt} + \beta_1 \sigma_{Vjt} + \xi_{jt}, \qquad (10)$$

where IPP_{jt} is the risk-adjusted insurance premium per dollar of insured deposits, σ_{Vjt} is the annualized standard deviation of asset returns for bank j in quarter t.

Hypothesis 4: $\beta_1^1 \leq 0$. The β_1^1 is the first slope coefficient in the following regression:

$$IPP_{jt} = \beta_{0jt} + \beta_1^1 d\sigma_{Vjt} + \beta_1^0(1\text{-}d)\sigma_{Vjt} + \xi_{jt}, \qquad (12)$$

where d = 1 for undercapitalized institutions and 0 otherwise.

Panel A: Pooled regression tests of Hypotheses 2 and 4 using the fixed-effects model with bank- and time-specific intercepts.

	β_1	t-stat.	β_1^1	t-stat.	β_1^0	t-stat.
Without dummies	0.004	0.86				
With dummies			0.019*	1.78	0.006	1.19

Panel B: Instrumental-variable estimation: pooled regression tests of Hypotheses 2 and 4 with bank- and time-specific random effects. σ_{Vjt-2} is used as the instrumental variable for σ_{Vjt}.

	β_1	t-stat.	β_1^1	t-stat.	β_1^0	t-stat.
Without dummies	0.004	1.05				
With dummies			0.027***	2.68	0.003	0.66

* significant at the 10 percent level using one-tailed test.
** significant at the 5 percent level using one-tailed test.
*** significant at the 1 percent level using one-tailed test.