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The examination of financial ratios of federal credit unions before and after the provision of federal deposit insurance in 1971 shows that their exposure to credit risks grew significantly faster after the provision of insurance. Following a short transition, both the quality of loans and the ability of credit unions to absorb loan losses declined. This evidence supports the recent literature contending that the pricing of federal deposit insurance unintentionally encourages risk taking by insured institutions.

From 1980 to 1983, oil consumption in most industrial countries declined, even though the real dollar price of oil fell and world economic activity increased. Lagged adjustment to the sharp oil price increase in 1979 and appreciation of the dollar more than offset the effects of economic growth and the oil price decline. With adjustment to the 1979 oil price increase nearly complete, continuing economic growth and any decline in the value of the dollar can be expected to result in increased oil consumption.

Analysis of income patterns for Texas agricultural households shows that farmers in West Texas are the least diversified and, consequently, the most susceptible to economic shocks. Blacklands and East Texas farm households generally have smaller incomes, but their income flows are spread more evenly over different sources, making these households better able to withstand unexpected declines in income from agriculture. West Texas farmers and ranchers are also exposed to more risk because of the region's low and highly variable rainfall.

Deposit Insurance, Moral Hazard, and Credit Unions

*By Robert T. Clair**

Much attention has recently been given to the hypothesis that improperly priced federal deposit insurance has induced greater risk taking by depository institutions, especially in a deregulated environment. There have been arguments against bank deregulation on this basis, and the Federal Deposit Insurance Corporation (FDIC) recognized the hypothesis in its 1983 proposal to restructure federal deposit insurance. In a recent article in this *Review*, Eugenie Short and Gerald O'Driscoll proposed regulatory changes to encourage the development of private deposit insurance, because market forces would cause private insurers to price insurance properly and not subsidize risk taking.[1] While a strong theoretical argument can be made supporting this hypothesis of greater risk taking, there has been relatively little empirical underpinning. The purpose of this article is to provide empirical evidence. The analysis is conducted using data from federal credit unions during the postwar period.

Utilizing data from the credit union industry offers a unique advantage over using data from other depository institutions in studying the problem of moral hazard. Federal deposit insurance was made available to credit unions beginning in 1971.[2] It is feasible, therefore, to examine the impact of deposit

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1. See John H. Kareken, "Deregulating Commercial Banks: The Watchword Should Be Caution," *Federal Reserve Bank of Minneapolis Quarterly Review*, Spring-Summer 1981, 2; Federal Deposit Insurance Corporation, *Deposit Insurance in a Changing Environment: A Study of the Current System of Deposit Insurance Pursuant to Section 712 of the Garn-St Germain Depository Institution Act of 1982, Submitted to the United States Congress by the Federal Deposit Insurance Corporation* (Washington, D.C.: FDIC, April 1983), I-7; and Eugenie D. Short and Gerald P. O'Driscoll, Jr., "Deregulation and Deposit Insurance," *Economic Review*, Federal Reserve Bank of Dallas, September 1983, 11–22.

2. Legally, a credit union does not accept deposits but issues shares in the credit union to its members. In fact, credit union shares are so similar to deposits that all references in this paper will be to deposits and deposit insurance.

insurance on credit unions by using postwar time series. Banks and savings and loan associations (S&Ls) received federal deposit insurance in 1934. For these institutions, it would be difficult if not impossible to separate the effects of deposit insurance from the effects of the Great Depression and World War II.

The statistical analysis conducted here suggests that federal deposit insurance has resulted in federal credit unions being more exposed to credit risks. The empirical evidence also indicates that credit union liquidity, which declined over the entire period, was not affected by deposit insurance. On balance, these results support the basic hypothesis that federal deposit insurance induced greater risk taking by insured depository institutions.

The development of moral hazard

The hypothesis that federal deposit insurance has induced greater risk taking by financial institutions is based on the premise that managers of financial institutions exploit the subsidy provided by deposit insurance by paying less heed to risk and investing in higher-yield, higher-risk assets. The cost of deposits is less than it otherwise would be because the deposits are insured. Even if the institution becomes riskier, the cost of deposits will not rise commensurate with the risk undertaken. The reason is that depositors perceive their accounts to be insured and will not demand compensation for their deposits being at greater risk. Furthermore, the premium paid for deposit insurance does not increase as the institution becomes riskier. Hence, management can increase the net interest margin by using lower-cost insured deposits to fund higher-yield, higher-risk assets. The result is higher net interest income, but higher loan losses are also likely to occur. Managers expect an increase in interest income to outweigh the expected increase in loan losses.[3] This argument is one of moral hazard: insuring the deposits of the institution causes the institution to act in a riskier manner.

The moral hazard problem will develop if the deposit insurance subsidizes risk taking. In the case of credit unions, this subsidization is easily established. The insurance premium is the same percentage of insured deposits for all credit unions regardless of risk. The premium is currently one-twelfth of 1 percent of the total amount of deposits in insured accounts, the same as for banks and S&Ls. This disregard of risk in pricing insurance takes place despite the fact that at least some risk characteristics can be measured.[4] Thus, there is a problem of moral hazard from improperly pricing insurance, as opposed to an adverse selection problem from not being able to measure risk.

Given that the moral hazard exists, several other factors will determine how great an increase in risk exposure is likely to occur. One factor is the extent of insurance coverage. Another is the availability of high-yield, high-risk assets that credit unions can acquire. Still another factor is the intensity with which credit union management seeks to maximize profits.

In the case of federal credit unions, insurance coverage is extensive. Deposit insurance for federal credit unions covers nearly 100 percent of total deposits, and deposits are 90 percent of total liabilities plus equity. The vast majority of credit union assets are funded by insured deposits.[5]

3. If managers inappropriately price the loans and do not charge a sufficient risk premium, it is possible for loan losses to exceed any gain in interest income. For a discussion of how regulatory safety-net mechanisms, such as deposit insurance, might result in the inappropriate pricing of loans, see Gerald P. O'Driscoll, Jr., and Eugenie Dudding Short, "Safety-Net Mechanisms: The Case of International Lending," Federal Reserve Bank of Dallas Research Paper no. 8404 (Dallas, May 1984).

4. One characteristic that could be used to price insurance premiums for risk is age of the credit union. The probability of a newly chartered credit union failing is approximately 40 percent. This probability declines over time; after 16 years of operation, the probability of failure approaches zero. Another credit union characteristic that provides information on the probability of failure is the type of membership. For example, a credit union of hotel employees is over three times more likely to fail than a credit union of telephone workers. The probability of failure has been calculated for each major type of membership criterion by the National Credit Union Administration, but this information has not been used to price insurance premiums. See Donald J. Melvin, Raymond N. Davis, and Gerald C. Fischer, *Credit Unions and the Credit Union Industry: A Study of the Powers, Organization, Regulation and Competition* (New York: New York Institute of Finance, 1977), 55.

5. Coverage has changed over the years as the volume of large deposits grew and limits to coverage were raised. Over the 1971–81 period, coverage never fell below 97 percent of total deposits. See Annual Report of the National Credit Union Administration, various issues.

A second factor—ability of credit unions to acquire high-risk, high-yield assets—is restricted in the case of federal credit unions by their charters. Credit unions can only make loans to members, to other credit unions, or through loan participations with other credit unions. The result is a loan portfolio composed primarily of secured consumer loans for the purchase of consumer durables, such as automobiles. The investment portfolio is even more restricted, as funds can only be placed in what appear to be low-risk assets. Credit unions can invest in government obligations, federally sponsored agency obligations, or obligations guaranteed by the government. Credit unions can also place funds in commercial banks or insured thrift institutions.[6]

Because of asset restrictions, the riskiest assets for credit unions are likely to be loans to members. These assets are definitely less liquid than assets in the investment portfolio, and they are more likely to have a greater risk of default. Hence, it can be hypothesized that the effect of moral hazard would be to encourage credit unions to increase their lending to members and reduce their investment portfolios. While this action may not seem to cause a large increase in risk, it demonstrates the effect of moral hazard within the confines of the present regulatory framework.

The willingness of credit union management to accept greater risk in order to obtain higher returns is important in determining the increase in exposure to risk as a result of moral hazard. Credit unions are, by charter, nonprofit cooperative organizations. If credit union management is extremely averse to accepting greater risk or if management has little incentive to increase risk, there would be little motivation to trade risk for return in spite of the incentive of the moral hazard. If this were the case, the provision of federal deposit insurance should have little or no effect on credit union portfolios. If, however, incentives do exist for credit union management to increase profits, revenues, or credit availability, then the moral hazard could result in a sizable increase in exposure to risk. These incentives may take the form of salary increases or increases in perquisites. The results of this investigation support the latter description of credit union management.[7]

Federal deposit insurance

Charters for federal credit unions were first granted in 1934. From that time until 1970, there was no federal deposit insurance for credit unions. In 1970, legislation providing federal deposit insurance was enacted, which created the National Credit Union Share Insurance Fund.

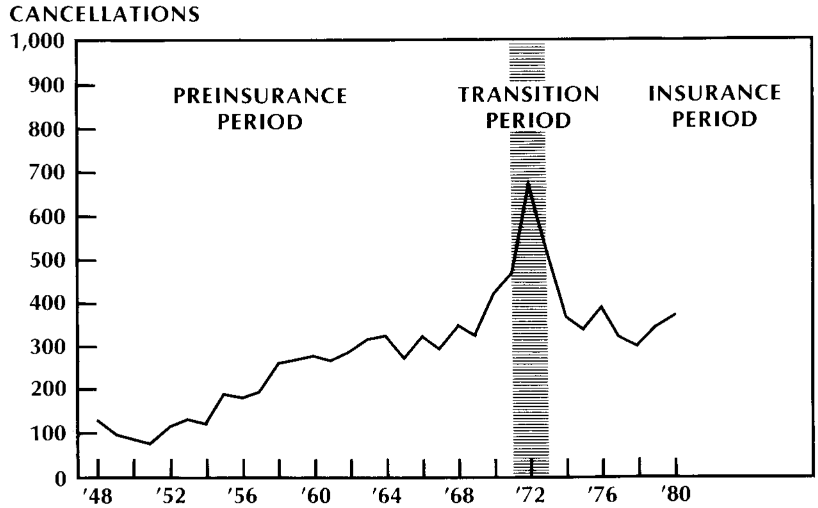
Implementation of the insurance program began in 1971 and spanned several years. Federal credit unions were required to obtain insurance and, to qualify for the insurance, had to meet a standard for financial condition that was higher than had been enforced previously. Any federal credit union that did not qualify for deposit insurance because of poor financial condition was issued a temporary insurance certificate. These certificates provided deposit insurance to the substandard credit unions for two years. At the end of the two-year period, such institutions were required either to have improved their financial condition to the point of being insurable, to have merged with other credit unions, to have switched to state charters, or to face liquidation. By the end of 1973, all federal credit unions were insured, and most transitory effects from the implementation of the insurance program had worked their way through the system.

The method by which the deposit insurance program was implemented suggests that a time series study of the effects of deposit insurance on federal credit unions in the postwar period should be di-

6. Over most of the period under study, 1948 to 1982, such investments would generally be considered risk-free assets. Any nongovernment obligation is, however, risky to some extent. Some credit unions did sustain capital losses as a result of the failure of Penn Square Bank of Oklahoma City. The perception of riskiness of some of these investments may have changed in recent years. For a detailed description of federal credit union asset powers, see Section 107 (12 U.S.C. 1757) of the Federal Credit Union Act as amended to May 1, 1977, in Melvin, Davis, and Fischer, *Credit Unions and the Credit Union Industry*, 237–43.

7. While there exist a number of theoretical models of credit union behavior, almost nothing in the way of empirical tests of these models has been published. Barry Keating did produce empirical results suggesting that credit union managers maximize their own utility subject to a minimum benefit constraint for members. This finding would be consistent with the statement that credit union management may be rewarded indirectly through perquisites for increased returns. See Barry P. Keating, "Prescriptions for Efficiency in Nonprofit Firms," *Applied Economics* 11 (September 1979): 321–32.

Chart 1
Total Cancellations of Federal Credit Union Charters



SOURCE OF PRIMARY DATA: National Credit Union Administration.

vided into three distinct periods. The first period, spanning 1948 to 1970, will be referred to as the preinsurance period. The second, a transition period, includes 1971, 1972, and 1973. Finally, the insurance period includes 1974 to the latest year for which suitable data are available. The problem of moral hazard is hypothesized to have developed during this third period.

It is important to include the transition period as a separate period in analyzing the effects of deposit insurance on federal credit unions. The enforcement of the higher standard of financial condition had a significant effect on the financial ratios used here to assess risk in the credit union industry. Empirical analysis must control for the changes during the transition period so that the results will not be biased.

Two sources of bias might result from not controlling for transition effects. The first is the extremely large number of cancellations of federal credit union charters during the 1971–73 period. In 1972, cancellations reached a record high of 672 (Chart 1). It is reasonable that the credit unions in the poorest financial condition were the first to have their charters canceled. As these credit unions were eliminated, the financial condition of the industry as a whole improved.

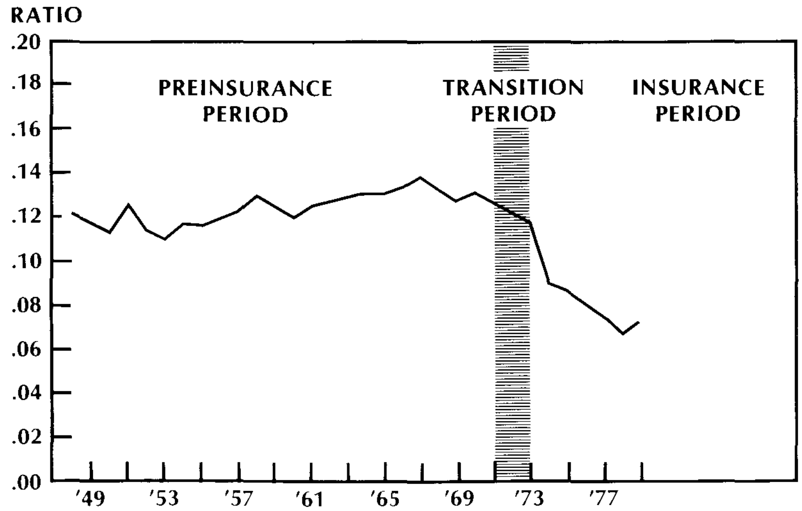
The second source of bias stems from the sudden enforcement of stricter financial standards for credit unions whose charters were not canceled. For example, credit unions wrote off a large amount of delinquent loans during the transition period, so the ratio of delinquent loans to total loans improved substantially. At the same time, writing off the delinquent loans required reductions in capital. These transitory effects should span only a year or two, but it is vitally important to control for them.

Financial ratios as risk measures

The moral hazard hypothesis implies simply that credit unions took on greater risk following the provision of deposit insurance. To test this hypothesis, several financial ratios of federal credit unions are examined to determine if any significant changes occurred in their trends over time that imply greater risk taking during the insurance period. The financial ratios used are designed to measure exposure to two of the basic risks of financial intermediaries: exposure to liquidity risk and exposure to credit risk.[8] Descriptions of the financial ratios are provided next, including rationales for their choice, followed by descriptions of the data and the test procedure.

Chart 2

Capital Ratio for Federal Credit Unions



SOURCE OF PRIMARY DATA: National Credit Union Administration.

Finally, regression results are presented and interpreted.

The first financial ratio examined is a capital ratio, which measures a credit union's ability to absorb loan losses. The ratio used here is the sum of undivided earnings plus regular and other reserves divided by loans. This ratio is the proportion of loans that could default and be absorbed by capital surplus. A decrease in this ratio would indicate a riskier credit union, other things equal.

In Chart 2 the capital ratio is plotted for 1948 to 1979. The capital ratio was about the same in 1970 as in 1949. During the transition period there was a slight decline. It was in the insurance period that the sharp decline occurred, suggesting a greater exposure to credit risks.

The adequacy of a credit union's capital is directly related to the quality of its loan portfolio. The second financial ratio, the delinquency rate, is a measure of the quality of the loan portfolio. The delinquency rate is calculated as the amount of loans that are delinquent in payments two months or more divided by total loans. Other things equal, an increase in the delinquency rate indicates a decline in the quality of the loan portfolio and greater exposure to credit risks.

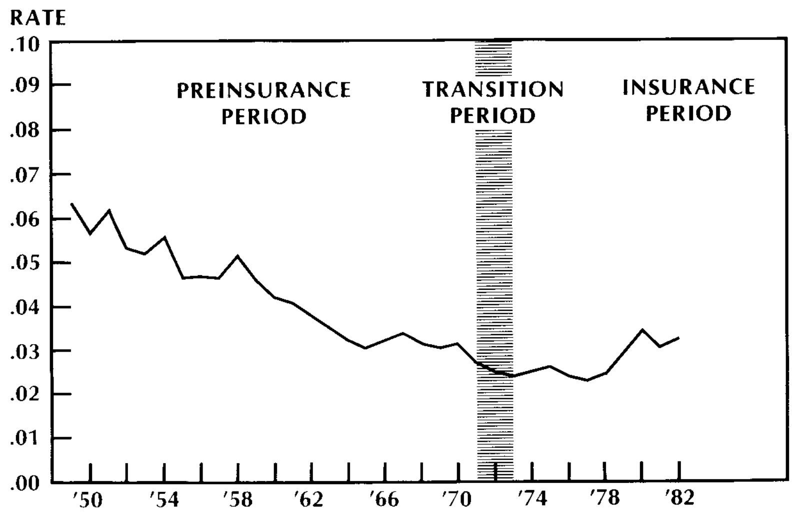
The delinquency rate, plotted for 1949 to 1982 in Chart 3, trended downward from 1949 to the middle 1960s, indicating an improving loan quality. From the middle 1960s to the transition period, the rate was fairly constant. During the transition period the delinquency rate declined sharply, likely a result of the enforcement of higher financial standards. As credit unions wrote off questionable loans, the delinquency rate improved at the expense of the capital ratio. Following the transition period, the delinquency rate was flat initially and then began an upward trend. The rising delinquency rate beginning in 1977 suggests an increase in exposure to credit risks.

The third financial ratio measures exposure to liquidity risk. A credit union is illiquid if it cannot meet its short-term obligations resulting from either depositors withdrawing funds or borrowers exercising existing lines of credit. The ratio chosen to

8. Most depository institutions are also subject to interest rate risk. It is likely that credit unions, like other depository institutions, borrow short and lend long. Unexpected increases in interest rates can cause a decline in earnings. Unfortunately, there are no data for measuring interest rate risk exposure.

Chart 3

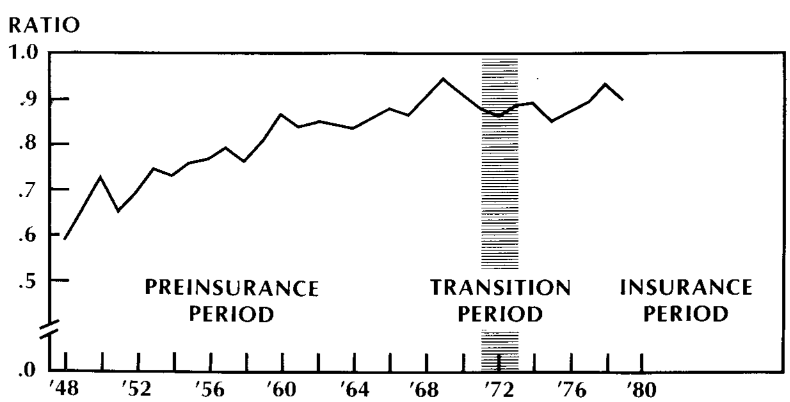
Delinquency Rate for Federal Credit Unions



SOURCE OF PRIMARY DATA: National Credit Union Administration.

Chart 4

Loan-to-Deposit Ratio for Federal Credit Unions



SOURCE OF PRIMARY DATA: National Credit Union Administration.

measure illiquidity is the proportion of loans to deposits. Loans are relatively illiquid assets compared with a credit union's investments. Thus, as loans increase relative to deposits, the credit union has fewer liquid assets available to meet sudden needs for liquidity, such as deposit withdrawals. As the loan-to-deposit ratio rises, the credit union is more exposed to liquidity risks.

The loan–deposit ratio, plotted in Chart 4, maintained a steady upward trend during the preinsurance period. During the transition period, this ratio declined, consistent with the enforcement of higher financial standards during the period. Subsequently, the loan–deposit ratio appears to follow the same upward trend evident for 1960 through 1970. It seems that the higher financial standards enforced during the transition period had only a minor effect on this ratio.

A test of the moral hazard hypothesis

The hypothesis that the provision of federal deposit insurance resulted in greater risk taking was tested using the financial ratios measuring risk as dependent variables in regressions. These regressions determined the time trend before insurance, the change in trend during the transition period, and the time trend during the insurance period. The hypothesis would predict a decline in riskiness during the transition period and an increase in risk taking during the insurance period.

The model used here assumes that financial ratios trend toward optimal values that are functions of several variables. The optimal value should be sensitive to changes in the expected rates of return on various assets and should also change in response to technological changes. These technological changes might include improvements in cash management techniques, the availability of new financial instruments, and changes in the regulatory environment. The test of the moral hazard hypothesis is to show evidence that the optimal values of the financial ratios changed, indicating greater exposure to risk in response to the provision of deposit insurance.

The optimal value is affected by both cyclical fluctuations and secular trends. In the model, balance sheet variables are used to capture the effects of cyclical fluctuations and unanticipated shocks. The secular effect is modeled with a time trend variable, and the changes in the secular trend resulting from the provision of deposit insurance are modeled using a binary variable that interacts with the time variable. The binary variable separates the data into two subperiods. If the coefficient of the binary variable interacted with the time variable is significantly different from zero, the implication is that the trend of the optimal ratio of the dependent variable has changed between the two periods.

Suppose credit union management behavior can be modeled as seeking an optimal value for the financial ratio, r^*, which is a linear function of time, t:

$$(1) \qquad r_t^* = a_1 + b_1 t, \qquad t = 1, 2, ..., n.$$

The management behaves as if it seeks to increase the financial ratio by b_1 in every period and to adjust for any difference between the actual and optimal ratio in the previous period. The financial ratio that actually occurs in period t will also be affected by any exogenous shocks or cyclical fluctuations that might occur in period t. The financial ratio in period t, r_t, is a function of the ratio in the previous period, r_{t-1}; the trend of the optimal ratio, b_1; the difference between r_{t-1}^* and r_{t-1}; and any exogenous shocks, e_t:

$$(2) \qquad r_t = r_{t-1} + b_1 + \lambda(r_{t-1}^* - r_{t-1}) + e_t, \quad 0 < \lambda < 1.$$

If the financial ratio in the previous period was equal to its optimal value and there are no exogenous shocks in the present period, then

$$(3) \qquad r_t = r_{t-1} + b_1 = r_{t-1}^* + b_1 = r_t^*.$$

The first two terms of equation 2 are expected to pick up the secular trend of the optimal value for the financial ratio. In this analysis, a simplification is made whereby the definition of equation r_t^* is substituted into equation 2, replacing the first two terms:

$$(4) \qquad r_t = a_1 + b_1 t + \lambda(r_{t-1}^* - r_{t-1}) + e_t.$$

The term $(r_{t-1}^* - r_{t-1})$ measures the impact of past exogenous shocks on the present ratio. The term e_t measures present exogenous shocks. Neither term can be measured directly. Assume that variable y_t is related to e_t as a direct proportion with a random error term. Then,

$$(5) \qquad e_t = \delta y_t + \mu_t, \qquad \mu \sim N(0,\sigma^2).$$

Furthermore, assume that credit union management reacts quickly to the exogenous shocks. The quick reaction implies that λ is very close to 1. In this structure, exogenous shocks before period $t-1$ have a negligible effect on r_t, as most of the adjustment for a shock will be accomplished in the next period. Then equation 4 can be rewritten as

$$(6) \qquad r_t = a_1 + b_1 t + \lambda \delta y_{t-1} + \delta y_t + \varepsilon_t, \quad \varepsilon_t = \lambda \mu_{t-1} + \mu_t.$$

The provision of deposit insurance changes the secular trend of r^* and, consequently, changes a_1 and b_1. If period x is the period during which deposit insurance was provided, then

$$(7) \qquad \begin{aligned} r^* &= a_1 + b_1 t, & t &= 1, ..., x-1. \\ &= a_2 + b_2 t, & t &= x, ..., n. \end{aligned}$$

To include this formulation of r^* in the above equation, first define a binary variable, D_1, that equals 1 in period x or later and equals zero in all earlier periods. Equation 6 can be restated as

$$(8) \; r_t = a_1 + D_1(a_2 - a_1) + b_1 t + D_1(b_2 - b_1)t + \lambda \delta y_{t-1} + \delta y_t + \varepsilon_t.$$

This equation can be placed in a regression format, as follows:

$$(9) \; r_t = \alpha + \beta_1 D_1 + \beta_2 t + \beta_3(D_1 \cdot t) + \beta_4 y_{t-1} + \beta_5 y_t + \varepsilon_t,$$

where $\alpha = a_1$
$\beta_1 = (a_2 - a_1)$
$\beta_2 = b_1$
$\beta_3 = (b_2 - b_1)$
$\beta_4 = \lambda \delta$
$\beta_5 = \delta$
$\varepsilon_t =$ normally distributed random error.

Equation 9 requires one more change in specification. The first four terms on the right-hand side of equation 8 model the secular trend of the optimal value of the financial ratio, including its shift in period x. The remaining terms capture the effects of exogenous shocks and cyclical fluctuations in this period and past periods. Under this structure the introduction of deposit insurance will cause a change in the secular trend. The change in the secular trend is estimated using the first four terms in equation 9. Adjustments are also required to reach this new secular trend. In order to separate these adjustments from other adjustments for exogenous shocks or cyclical fluctuations, an additional binary variable, D_2, is introduced into the equation. Transition effects are captured by $(D_2 \cdot t)$. The interactive variable $(D_2 \cdot t)$ was chosen to determine the change in slope.[9]

$$(10) \qquad r_t = \alpha + \beta_1 D_1 + \beta_2 t + \beta_3(D_1 \cdot t) + \beta_4 y_{t-1} + \beta_5 y_t + \beta_6(D_2 \cdot t) + \varepsilon_t.$$

Equation 10 is the format of the estimated regressions. The binary variable D_1 equals 1 for 1974 and later years, separating the insurance period from earlier periods. The binary variable D_2 is defined as equal to 1 for the years 1971-73 and zero for all other years. D_2 is used to isolate the changes that occurred during the transition period. The variable y_t, which captures exogenous shocks and cyclical movements in r_t, is different depending on the dependent variable. For the capital ratio and the loan-deposit ratio, y_t is defined as the percentage change in total assets over the past year divided by 100. For the delinquency rate, y_t is defined as the percentage change in total loans over the past year divided by 100.

The data used in the liquidity ratio and capital ratio regressions are for 1948 to 1979. Post-1948 data were used to avoid any effects from World War II. The data after 1979 are not comparable with those for earlier periods because of a binding usury ceiling in late 1979 and 1980 and the imposition of credit controls in early 1980.[10]

9. Because an adjustment is modeled as a process and not a sudden shift, no binary variable for a change in the intercept term has been included.

10. The data are collected and published by the National Credit Union Administration. All the ratios are calculated from the aggregated outstanding balances reported for all federal credit unions in the United States, Guam, Puerto Rico, and the Virgin Islands. Assets and liabilities are reported as outstanding balances on December 31.

The two economic policies, usury ceilings and credit controls, resulted in a severe decline in loan growth. The decline in loan growth caused a sharp drop in the liquidity ratio and an increase in the capital ratio. These ratios did not recover

The delinquency rate regression uses data from 1949 to 1982—all the available postwar data. The years 1980-82 are included for the delinquency rate because this ratio should not be overly affected by the usury ceiling or the credit controls.

The postwar period was a relatively stable one for credit unions. Asset powers were not altered in any significant way during this period. Credit unions weathered the business cycles in relatively good form. Credit unions have a propensity to be organized among workers with stable employment patterns—for example, government employees and public utility employees—as opposed to workers in less secure occupations. In addition, credit unions were probably not affected as much as savings and loan associations or banks by the problems resulting from rising interest rates.

Interest rate ceilings on deposits have typically been higher at credit unions than at banks or other depository institutions, making credit unions less exposed to disintermediation. When their interest rate ceilings did become binding, credit unions were innovative in creating unregulated instruments, and the National Credit Union Administration was quick to respond by raising the ceiling. Also, credit unions maintain only a small share of assets in mortgages, resulting in a relatively shorter average asset maturity than in the case of S&Ls. Consequently, credit unions did not face the severe earnings pressure S&Ls did during periods of rising interest rates.

The moral hazard hypothesis implies that the coefficient β_3 is significant, indicating greater risk-taking behavior. Increased exposure to credit risk would be evidenced by β_3 being positive in the delinquency rate regression and negative in the capital ratio regression. Greater exposure to liquidity risk would be denoted by a positive β_3 in the loan-deposit ratio regression.

The enforcement of higher financial standards during the transition period should have two effects. First, the β_6 coefficient should show improving financial condition. This would mean a positive β_6 in the capital ratio regression and a negative β_6 in the regressions for the delinquency rate and the loan-deposit ratio. Second, credit unions are expected to begin the insurance period in a better financial condition than existed at the end of the preinsurance period. Consequently, the β_1 coefficient is expected to be positive in the capital ratio regression and negative in the delinquency rate and loan-deposit ratio regressions.

The sign of the coefficient β_5 can be predicted in some instances. In the capital ratio regression, β_5 is expected to be negative. In this regression, y_t is asset growth. If an exogenous shock occurs that results in an increase in assets, it is likely that the capital ratio will decline. As earnings from the additional assets accrue, undivided earnings and reserves can be increased. The time lag between the growth in assets and the growth in capital causes the capital ratio to decline. As a result, β_5 is expected to be negative.

In the delinquency rate regression the coefficient β_5 is also expected to be negative. In this regression, y_t is loan growth. An exogenous shock resulting in higher loan growth would cause the denominator of the delinquency rate to increase. Furthermore, economic conditions that encourage loan growth, high employment, and strong income growth are also conditions that indicate consumers are less likely to let their payments lapse. Hence, it would be likely that the numerator in the delinquency rate would decrease. Positive loan growth would be associated with a decline in the delinquency rate.

It is not possible to hypothesize *a priori* the sign of β_5 in the loan-deposit ratio regression. The variable y_t in this regression is asset growth. It is impossible to determine whether growth in total assets would be correlated with growth in loans or deposits or both.

Regression results

The regressions were estimated using data for all

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to their previous levels in 1981 and 1982. It is most likely that the recession from July 1981 to November 1982 stifled loan growth. The lack of growth in lending prevented the two ratios from returning to their previous levels.

If the data from 1980 through 1982 are included in the regressions, the results indicate that there were no significant changes in the trend of the capital ratio and the loan-deposit ratio. In the case of the capital ratio, nothing could be further from the truth. The capital ratio was fairly constant in the years before deposit insurance. Following the provision of insurance, this ratio dropped sharply from 1970 to 1979, reaching a record postwar low. From 1980 through 1982, it increased dramatically because of the usury ceiling and the credit controls. To describe this 12-year period as essentially flat would be misleading. Rather than attempt to explicitly model the effects of the usury ceiling and the credit controls, the last few observations were eliminated.

Table 1
REGRESSION RESULTS FOR FINANCIAL RATIOS

	Coefficients						
	α	β_1	β_2	β_3	β_4	β_5	β_6
		Independent variables					
Regression	Constant	D_1	t	$(D_1 \cdot t)$	y_{t-1}	y_t	$(D_2 \cdot t)$
Capital ratio	.1349	.1847	.00016	-.00524	-.02926	-.05838	-.00012
	(13.22)***	(3.86)***	(.69)	(-4.81)***	(-1.24)	(-2.80)***	(-1.25)
	$\bar{R}^2$ = .9642; DW = 1.82; n = 32; rho = .2345. (1.36)						
Delinquency rate	.0963	-.1061	-.00176	.00259	.00551	-.02919	.00001
	(16.72)***	(-4.88)***	(-10.18)***	(5.14)***	(1.41)*	(-7.10)***	(.18)
	$\bar{R}^2$ = .9757; DW = 1.56; n = 34; rho = .7328. (6.28)						
Loan-deposit ratio	.4158	.0326	.01341	-.00393	.04836	.04048	-.00236
	(5.68)***	(.09)	(8.17)***	(-.50)	(.26)	(.24)	(-3.48)***
	$\bar{R}^2$ = .8931; DW = 1.79; n = 32.						

Independent variables

D_1 = 1 for 1974 or later years; zero otherwise.

t = a trend variable that takes the values 1, 2, . . ., 50; 1948 corresponds to (t = 16).

y_t = (total assets$_t$ − total assets$_{t-1}$) ÷ total assets$_{t-1}$ for the capital ratio regression and the loan-deposit ratio regression.

= (total loans$_t$ − total loans$_{t-1}$) ÷ total loans$_{t-1}$ for the delinquency rate regression.

D_2 = 1 for 1971, 1972, and 1973; zero otherwise.

Dependent variables

Capital ratio = (undivided earnings + reserves) ÷ total loans.

Delinquency rate = loans with payments delinquent for two months or more ÷ total loans.

Loan-deposit ratio = total loans ÷ total savings.

NOTE: Figures in parentheses are t statistics; * indicates significance of the coefficient at the .10 level, ** at the .05 level, and *** at the .01 level, using a one-tailed test when the sign of the coefficient can be hypothesized.
$\bar{R}^2$ is the coefficient of determination adjusted for degrees of freedom.
DW is the Durbin-Watson autocorrelation test statistic.
n is the number in sample.
Rho is the first-order autocorrelation coefficient.
Capital ratio and delinquency rate regressions were corrected for autocorrelation.

Table 2
PREDICTED AND ACTUAL SIGNS OF REGRESSION COEFFICIENTS

Regression	Moral hazard hypothesis β_3 Predicted	Actual	Transition period effects β_1 Predicted	Actual	β_6 Predicted	Actual	Adjustments to exogenous shocks β_5 Predicted	Actual
Capital ratio	−	−*	+	+*	+	n.s.	−	−*
Delinquency rate	+	+*	−	−*	−	n.s.	−	−*
Loan-deposit ratio	+	n.s.	−	n.s.	−	−*	n.a.	n.s.

* Significant with α equal to .01.
n.s.—Not significantly different from zero.
n.a.—Not applicable.

federal credit unions, and the results are reported in Table 1. The model fits the data quite well, as most of the variation in the financial ratios is explained. The regressions for the capital ratio and the delinquency rate indicate increased exposure to credit risk following the provision of deposit insurance. The loan-deposit ratio was unaffected by the provision of deposit insurance; however, it did maintain a steady upward trend, indicating increasing illiquidity, over the entire 1948–79 period.

The fact that credit union liquidity was unaffected does suggest that examination of the basic hypothesis within the context of a richer theoretical model of a depository institution might prove productive. It may be that deposit insurance lowered the cost of bankruptcy to credit union members but did not lower the cost that results from a credit union being illiquid. Credit unions would then be more willing to accept greater risk of bankruptcy for a greater return but not the greater risk of being illiquid.

The implications of the regression results are summarized in Table 2. The first two columns of the table provide the predicted and actual signs for the β_3 coefficient based on the moral hazard hypothesis. The coefficient was of the appropriate sign and significant for the capital ratio and the delinquency rate. These results imply that federal credit unions are more exposed to credit risks as a result of the moral hazard. The β_3 coefficient was insignificant in the loan-deposit ratio regression. This result implies that deposit insurance had no measurable effect on federal credit union exposure to liquidity risks.

The coefficients β_1 and β_6 both model the effect of the transition period. Three of the six coefficients were significant and appropriately signed. This supports the hypothesis that the transition period was marked by improving financial conditions as federal regulators enforced higher standards. The coefficient β_5, modeling the effect of cyclical fluctuations and exogenous shocks on the financial ratios, had the predicted sign and was significant in both cases where the sign could be hypothesized.

Conclusion

Regression analysis offers support of the hypothesis that the provision of federal deposit insurance resulted in credit unions taking on greater risk. In particular, there is clear evidence that credit unions face increasing credit risks in the insurance period. Both the quality of the credit union industry's loan portfolio and its ability to absorb loan losses have diminished during the insurance period. Deposit insurance seems to have had no effect on credit union liquidity; however, credit union liquidity declined significantly over the entire 1948–79 period.

The conditions that resulted in federal credit unions becoming more exposed to credit risks also exist for other depository institutions. The National Credit Union Share Insurance Fund was modeled on the Federal Deposit Insurance Corporation and the Federal Savings and Loan Insurance Corporation. Subsidized deposit insurance is provided to commercial banks, savings and loan associations, and mutual savings banks. The problems resulting from moral hazard that developed in the credit union industry are likely to exist for all these insured institutions.

The problems would be dealt with most directly by pricing the premiums for deposit insurance to reflect the risk of individual institutions. This might be accomplished within the present system of federal insurance. It could also be accomplished by encouraging the development of a private deposit insurance industry. A strong argument can be made that the private industry would have a greater incentive to price risk properly than would the federal insurance programs. In either case, given that deregulation of the financial industry will continue, it is vitally important to remove the subsidy inherent in the present insurance system in order to eliminate moral hazard and to provide a "level playing field" for insured depository institutions to compete with uninsured financial institutions.

The Effects of Oil Prices and Exchange Rates on World Oil Consumption

*By Stephen P. A. Brown and Keith R. Phillips**

From 1980 to 1983, oil consumption in most industrial countries declined, even though the real dollar price of oil fell and world economic activity increased. A common explanation for this decline is that consumers continued to adjust to the sharp oil price increase occurring in 1979. A more complete analysis reveals that exchange rate movements have also reduced oil consumption. Because world oil prices are denominated in U.S. dollars, movements in exchange rates can alter the price of oil faced by countries other than the United States. In fact, increases in the value of the dollar raised the effective price of oil for some major industrial countries to levels that were higher in 1983 than in 1980.

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For the seven largest industrial economies (Canada, France, Italy, Japan, the United Kingdom, the United States, and West Germany), simple econometric models and simulation analysis were used to estimate the effects that three factors—lagged adjustment to oil prices, changes in the dollar price of oil, and appreciation of the dollar—have had on world oil consumption. In 1983, these seven economies consumed 27 million barrels of oil per day, accounting for about three-fourths of free-world oil consumption and about half of total world oil consumption.

Results of the simulations indicate that a lagged adjustment to the oil price shock occurring in 1979 was the largest single factor reducing world oil consumption from 1980 to 1983 but that the adjustment was largely completed by the end of 1983. The results further indicate that the stimulative effects of a decline in the real dollar price of oil from 1980 to 1983 were more than offset by appreciation of the dollar against other major currencies.

Modeling oil demand

A simple model of oil demand was developed for each of the seven economies. Each model was constructed with two equations. The first equation indicates the quantity of oil that would be demanded in the long run for a given gross national product and price of oil. The second indicates the rate at which adjustment to the long run occurs. A reduced form was estimated to obtain both short-run and long-run price elasticities of oil demand and to obtain the rate of adjustment.

The consumption of oil products is commonly thought to be a function of the general level of economic activity, the real prices of oil products, and other variables. If other variables are ignored as being of secondary importance and if the elasticity of aggregate oil demand with respect to gross national product (GNP) is assumed to be 1,[1] oil demand can be expressed as a long-run consumption-to-GNP ratio that is a function of price:

$$(1) \qquad \left(\frac{C^*_{j,t}}{GNP_{j,t}}\right) = A_j \, PRICE_{j,t}^{\eta_j},$$

in which

$C^*_{j,t}$ = long-run value of oil consumption in country j during period t (the consumption that would prevail with $GNP_{j,t}$ if there was complete adjustment to the price of oil, $PRICE_{j,t}$)

$GNP_{j,t}$ = real gross national product in country j during period t, as measured in country j's own currency

A_j = a scaling constant for country j's crude oil demand

$PRICE_{j,t}$ = real price of crude oil faced by country j during period t, denominated in the country's own currency (a proxy for internal oil product prices)

η_j = long-run price elasticity of crude oil demand for country j.

Reflecting the difference between short-run and long-run adjustment, oil consumption is modeled as achieving only a partial adjustment to changes in price during a given quarter. This approach creates a gap between the consumption–GNP ratio that prevails in a given quarter and the long-run consumption–GNP ratio. The gap was modeled as closing at a constant rate of δ_j percent in each time period:

$$(2) \qquad \frac{C_{j,t}/GNP_{j,t}}{C_{j,t-1}/GNP_{j,t-1}} = \left(\frac{C^*_{j,t}/GNP_{j,t}}{C_{j,t-1}/GNP_{j,t-1}}\right)^{\delta_j}.$$

Substituting (2) into (1), reducing and rearranging terms, and then taking the natural logarithm yield the reduced-form equation that was used for estimation:

$$(3) \qquad \ln\left(\frac{C_{j,t}}{GNP_{j,t}}\right) = \alpha_j + \beta_j \ln(PRICE_{j,t}) + \gamma_j \ln\left(\frac{C_{j,t-1}}{GNP_{j,t-1}}\right),$$

in which

$\alpha_j \equiv \delta_j \ln(A_j)$

$\beta_j \equiv \delta_j \eta_j$, the short-run price elasticity of oil demand

$\gamma_j \equiv (1 - \delta_j)$.

The estimated parameters from equations 1 and 2 are obtained as follows:

$\hat{\delta}_j = (1 - \hat{\gamma}_j)$, the estimated rate of adjustment

$\ln(\hat{A}_j) = \hat{\alpha}_j/(1 - \hat{\gamma}_j)$

$\hat{\eta}_j = \hat{\beta}_j/(1 - \hat{\gamma}_j)$, the estimated long-run price elasticity of oil demand.

The model used for estimation has two advantages. It allows oil demand to show a partial (or lagged) adjustment to changes in the oil price, and it requires only a limited number of observations.[2]

Real oil prices

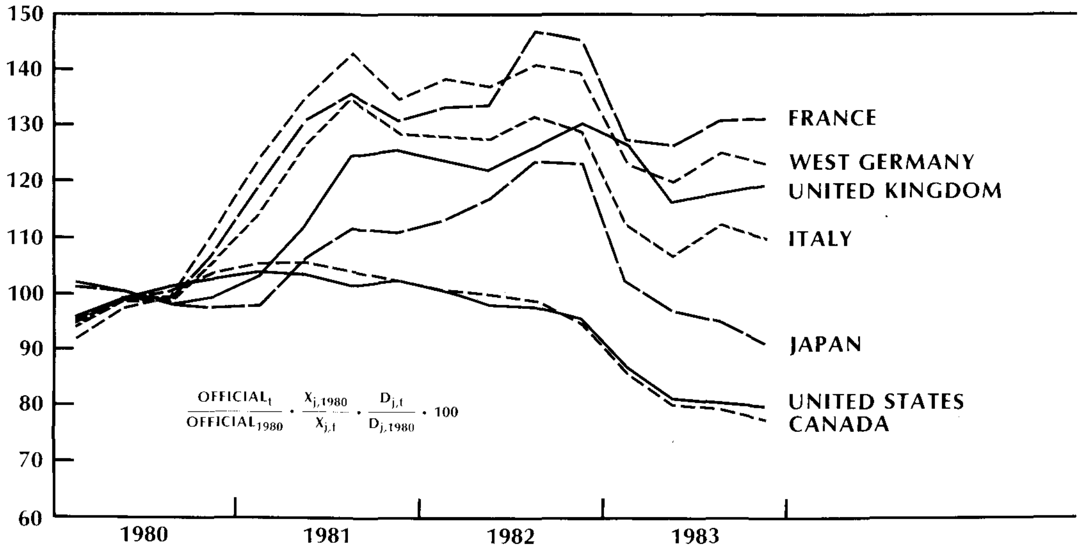
Because sufficient data on internal product prices

1. For estimating convenience it is assumed that the aggregate elasticity of oil demand with respect to GNP is 1. Given the short period over which demand is estimated, this assumption is probably not critical. There is some evidence, however, that the GNP elasticity of oil demand for the United States varies over the business cycle. See Stephen P. A. Brown, "A U.S. Economic Recovery Could Solidify World Oil Prices," *Energy Highlights*, Federal Reserve Bank of Dallas, April 1983.

2. The partial-adjustment model chosen for estimation imposes a Koyck lag structure on the crude oil prices, in which the coefficient of $PRICE_{j,t-k}$ for any period k preceding t is $\beta_j \gamma_j^k$. This functional form implies that there is a geometrically declining rate at which past prices influence the current consumption-to-GNP ratio. See G. S. Maddala, *Econometrics* (New York: McGraw-Hill Book Company, 1977), 142–43.

Real Prices of Mideastern Light Crude Oil

(1980 AVERAGE = 100)



SOURCES OF PRIMARY DATA: Federal Reserve Bank of New York
International Monetary Fund
Petroleum & Energy Intelligence Weekly, Inc.

were not readily available, an exchange-rate-adjusted crude price was used as a proxy for internal product prices in the estimation of oil demand. The price of crude oil faced by a country is only one component of the country's internal product prices. Other important components include taxes, refining costs, and transportation costs. Ignoring these other components of product prices implies that there is a constant relationship between the exchange-rate-adjusted price of crude oil and internal product prices.

Adjusted for inflation, the official U.S. dollar price of Mideastern light crude oil fell 17.5 percent from 1980 to 1983, as shown in the accompanying chart. During the same period the real price of officially priced Mideastern light crude oil, measured in each country's respective currency, decreased 21.2 percent for Canada and 3.3 percent for Japan. It increased 10.7 percent for Italy, 20.5 percent for the United Kingdom, 23.3 percent for West Germany, and 29.6 percent for France. Differences in movements in the price of oil across these seven countries reflect changes in real exchange rates.

The real price of crude oil faced by a country, as denominated in the country's currency, is a product of the country's exchange rate, inflation in that country, and the dollar price of oil:

$$(4) \qquad PRICE_{j,t} = X_{j,t} \cdot \left(\frac{D_{j,0}}{D_{j,t}}\right) OFFICIAL_t,$$

in which

$X_{j,t}$ = nominal exchange rate of country j's currency for U.S. dollars

$D_{j,0}$ = country j's implicit price deflator in the base period

$D_{j,t}$ = country j's implicit price deflator for period t

$OFFICIAL_t$ = official nominal U.S. dollar price of Mideastern light crude oil in period t.[3]

3. Spot prices for a raw material might be expected to be more indicative of the state of the market for its products. The spot

Alternatively, the price of crude oil faced by a country can be expressed as a product of the real U.S. dollar price of crude oil and the real exchange rate between the country's currency and the U.S. dollar:

$$(5) \qquad PRICE_{j,t} = E_{j,t}\, PRICE_{\$,t},$$

in which

$E_{j,t}$ = the real exchange rate, or $X_{j,t} \cdot (D_{j,0}/D_{j,t})(D_{\$,t}/D_{\$,0})$

$PRICE_{\$,t}$ = the real dollar price of oil, or $(D_{\$,0}/D_{\$,t})OFFICIAL_t$.

Estimating oil demand

Using the price series described in the preceding section, equation 3 was estimated for each of the seven countries with quarterly data from the first quarter of 1975 through the fourth quarter of 1983.[4] Oil consumption data proved to be the factor limiting the time period for which the model was estimated. For dates before 1975, data that were consistent with the post-1974 series could not be found.

crude oil market is very thin, however, and there is little statistical relationship between product prices and spot prices. There is a strong relationship between product prices and the official price of Mideastern light crude oil. See S. P. A. Brown and Keith R. Phillips, "Exchange Rates and the Estimation of World Oil Demand" (Work in progress, Federal Reserve Bank of Dallas).

4. Monthly oil consumption data for all seven countries were obtained from the U.S. Department of Energy. The data were transformed to quarterly values of average barrels per day and then seasonally adjusted with the X11 procedure contained in the Statistical Analysis System (SAS).

The quarterly GNP series were obtained or estimated with data from the International Monetary Fund, the Federal Reserve Bank of New York, and the economic staff of the British Embassy in the United States. For all seven countries except the United Kingdom, seasonally adjusted data were provided. The X11 procedure contained in SAS was used in seasonally adjusting the GNP series for the United Kingdom.

Exchange rate series were obtained from the International Monetary Fund. Implicit price deflators were obtained or estimated with data from the International Monetary Fund and the Federal Reserve Bank of New York. Oil price data are from the *Petroleum Intelligence Weekly* series for Mideastern light crude.

Using Durbin's H statistic, autocorrelation was judged to be present in original estimates for the models representing Canada, France, Italy, and West Germany. The autocorrelation was corrected in those cases.

As shown by Table 1, five of the seven models performed well. The Italian model was judged inappropriate because it exhibited a poor fit and the hypothesis of a zero coefficient on price could not be rejected with statistical confidence. Despite its good fit, the Canadian model was judged inappropriate because the hypothesis of a unitary coefficient on the lagged dependent variable could not be rejected with statistical confidence.[5]

The poor results obtained with the models representing oil demand for Canada and Italy are indicative of the problems inherent in using crude oil prices as a proxy for product prices. Government regulation and changes in oil product taxation over the period of estimation altered the relationship between crude oil prices and the oil product prices faced by consumers in these countries. The five other countries modeled also regulated their internal energy markets and changed their taxes to some extent during the period. Nevertheless, their internal product prices, unlike those in Canada and Italy, showed a strong relationship to exchange-rate-adjusted real crude oil prices.[6]

For the five acceptable models, short-run (one-quarter) price elasticities were estimated to range from −0.047 to −0.125, indicating a relatively low response of consumption to changes in price. The estimated rate of adjustment toward the long run varied from the Japanese low of 8.7 percent of the remaining gap per quarter to the French high of 37.2 percent of the remaining gap per quarter. With the estimated adjustment rates, 90 percent of the long-run adjustment to a price change would require 4 quarters for France, 25 quarters for Japan, 9 quarters for the United Kingdom, 24 quarters for the United States, and 12 quarters for West Germany. Long-run price elasticities were estimated to extend from −0.380 to −0.634, indicating some—but not a

5. A coefficient on the lagged dependent variable equal to or greater than 1 implies that the model is unstable.

6. See Brown and Phillips, "Exchange Rates and the Estimation of World Oil Demand."

Table 1

ESTIMATED PARAMETERS OF CRUDE OIL DEMAND

Country	Short-run price elasticity $\hat{\beta}_j$	Coefficient on lagged dependent variable $\hat{\gamma}_j$	Long-run price elasticity $\hat{\eta}_j$	R^2	$\bar{R}^2$	Durbin's H
Canada	-.022	.999	-40.688	.9294	.9250	*
(t)	(-1.612)	(21.683)				
(t, H_0: $\gamma_j = 1$)		(-.012)				
France	-.157	.628	-.421	.8786	.8710	*
(t)	(-3.370)	(5.102)				
(t, H_0: $\gamma_j = 1$)		(-3.018)				
Italy	-.026	.413	-.044	.1837	.1327	*
(t)	(-1.343)	(2.095)				
(t, H_0: $\gamma_j = 1$)		(-2.972)				
Japan	-.055	.913	-.634	.9746	.9730	-1.75
(t)	(-2.265)	(17.111)				
(t, H_0: $\gamma_j = 1$)		(-1.639)				
United Kingdom	-.125	.783	-.575	.9254	.9255	-.19
(t)	(-2.467)	(8.056)				
(t, H_0: $\gamma_j = 1$)		(-1.892)				
United States	-.047	.910	-.527	.9737	.9721	-1.14
(t)	(-2.360)	(17.275)				
(t, H_0: $\gamma_j = 1$)		(-1.708)				
West Germany	-.103	.728	-.380	.9477	.9444	*
(t)	(-3.668)	(8.797)				
(t, H_0: $\gamma_j = 1$)		(-3.279)				

* Corrected for autocorrelation using a computational routine that adjusts the variance-covariance matrix of the estimates for the presence of a lagged dependent variable. See Phoebus J. Dhrymes, *Distributed Lags: Problems of Estimation and Formulation*, 2d ed. (Amsterdam, New York, and Oxford: North-Holland Publishing Company, 1981), 199 (Theorem 7.1). Reported R^2 is based on the Durbin equation.

NOTE: $\bar{R}^2$ is the coefficient of determination adjusted for degrees of freedom. There are 32 degrees of freedom for each regression.

(t) is the standard t statistic for the null hypothesis that the true coefficient is zero.

(t, H_0: $\gamma_j = 1$) is the t statistic for the null hypothesis that the true value of γ_j is equal to 1.

Table 2

ESTIMATED EFFECT ON 1983 OIL CONSUMPTION OF CONTINUED ADJUSTMENT TO THE 1980 EXCHANGE-RATE-ADJUSTED OIL PRICE

	France	Japan	United Kingdom	United States	West Germany
	Thousands of barrels per day				
Simulated 1983 oil consumption					
Assuming exchange-rate-adjusted price held constant at 1980 level[1]	1,860	4,306	1,435	14,998	2,230
Assuming consumption–GNP ratio set at 1980 level[2]	1,965	5,218	1,479	17,577	2,411
Change in oil consumption attributable to continued adjustment to 1980 price	-105	-912	-44	-2,579	-181

1. Entry equals:

$$GNP_{j,1983}\left(\frac{C_{j,1980}}{GNP_{j,1980}}\right)^{\gamma_j^{12}} \cdot \prod_{k=0}^{11}\left(e^{\alpha_j} \cdot PRICE_{j,1980}{}^{\beta_j}\right)^{\gamma_j^k}$$

2. Entry equals: $GNP_{j,1983}(C_{j,1980}/GNP_{j,1980})$.

Table 3

ESTIMATED EFFECT ON 1983 OIL CONSUMPTION OF CHANGES IN THE EXCHANGE-RATE-ADJUSTED OIL PRICE OCCURRING BETWEEN 1980 AND 1983

	France	Japan	United Kingdom	United States	West Germany
	Thousands of barrels per day				
Simulated 1983 oil consumption					
Using fitted value from regression[1]	1,640	4,193	1,307	15,270	2,033
Assuming exchange-rate-adjusted price held constant at 1980 level	1,860	4,306	1,435	14,998	2,230
Change in oil consumption attributable to changes in adjusted oil price	-220	-113	-128	272	-197

1. The fitted value of consumption is equal to the simulated value of consumption obtained under the assumption that the adjusted price of oil took its actual values from 1980 to 1983.

Table 4
ESTIMATED EFFECT ON 1983 OIL CONSUMPTION OF CHANGES IN EXCHANGE RATES OCCURRING BETWEEN 1980 AND 1983

	France	Japan	United Kingdom	United States	West Germany
	Thousands of barrels per day				
Simulated 1983 oil consumption					
Using fitted value from regression	1,640	4,193	1,307	15,270	2,033
Assuming exchange rate held constant at 1980 level[1]	1,959	4,396	1,502	15,270	2,315
Change in oil consumption attributable to changes in exchange rates	-319	-203	-195	0	-282

1. Entry equals:

$$GNP_{i,1983}\left(\frac{C_{i,1980}}{GNP_{i,1980}}\right)^{\gamma_i^{12}} \cdot \prod_{k=0}^{11}\left(e^{\alpha_i} \cdot E_{i,1980}^{\beta_i} \cdot PRICE_{\$,k}^{\beta_i}\right)^{\gamma_i^k}.$$

Table 5
ESTIMATED DIFFERENCE BETWEEN 1983 OIL CONSUMPTION AND CONSUMPTION CONSISTENT WITH 1983 GNP, OIL PRICES, AND EXCHANGE RATES OVER THE LONG RUN

	France	Japan	United Kingdom	United States	West Germany
	Thousands of barrels per day				
Oil consumption consistent with 1983 GNP, oil prices, and exchange rates over the long run[1]	1,669	3,995	1,287	15,394	2,056
Average 1983 fitted oil consumption	1,640	4,193	1,307	15,270	2,033
Gap remaining[2]	29	-198	-20	124	23

1. Entry is the level of consumption that would eventually prevail if real gross national product, the real dollar oil price, and the real exchange rate were to remain at 1983 levels thereafter.
2. The gap is indicative of the direction in which consumption would change if GNP, oil prices, and exchange rates were to remain at their 1983 levels.

the oil price shock occurring in 1979. For the countries examined, appreciation of the dollar also had a strong effect in reducing oil consumption, more than offsetting a downturn in the real dollar price of oil. Because appreciation of the dollar preceded downward movement in the price of oil, the estimates provide some support for the notion that appreciation of the dollar against other major currencies has reduced oil demand and added to the downward pressure on the real dollar price of oil.

It appears that world oil consumption is poised to increase. For the economies examined, adjustment to the sharp oil price increase of 1979 appears nearly complete. Continuing or emerging economic growth of these economies can be expected to stimulate oil consumption further. Any decline in the foreign exchange value of the dollar will reduce crude prices faced by foreign economies and further stimulate world oil consumption.

Farmers and Economic Shocks: Ranking Texas Agricultural Production Regions

*By Hilary H. Smith**

Agriculture in Texas, as in the rest of the country, is prone to various kinds of economic turmoil: unfavorable weather, surges of pestilence, prices below the cost of production, and political maneuvers either here or abroad. In Texas during 1982 and 1983, there were massive hailstorms, a damaging drought, and inopportune freezing temperatures. Are farmers in some parts of Texas better able to deal with such economic disruptions than farmers elsewhere in the state?

To answer this question, Texas is divided into production regions, and various gauges of farmers' shock-bearing capacity are examined. The results show the western part of Texas, where about half the farm households are located and where about half the agricultural cash receipts are generated, is the most vulnerable. This area is most likely to suffer weather-related agricultural shocks and is the least capable of coping with the agricultural income disturbances.

Shocks and the Texas agricultural economy

Agricultural shocks affect farmers, farm workers, agribusinesses, and the general economy. This paper examines only the relative capacity among Texas farm households to handle agricultural income disruptions but acknowledges that other sectors of the agricultural and general state economy must absorb the disturbances as well. Agriculture is important to the state economy, generating around $10 billion of cash receipts and perhaps four times that much related economic activity. Moreover, farmers and ranchers in Texas are a powerful interest group, so their ability to handle agricultural income disruptions is of political significance.

The relative capacity of farm households to withstand economic shocks is analyzed here by comparing farmers in different regions. Texas is divided into 10 agricultural production regions, based on crop reporting districts as fitted to 1980 Census county groups. Map 1 shows these regions, their 1982 agricultural cash receipts, and the number of farm households in 1980. This article examines possible ways for farm households to ride out agricultural income shocks. These ways include diversified income

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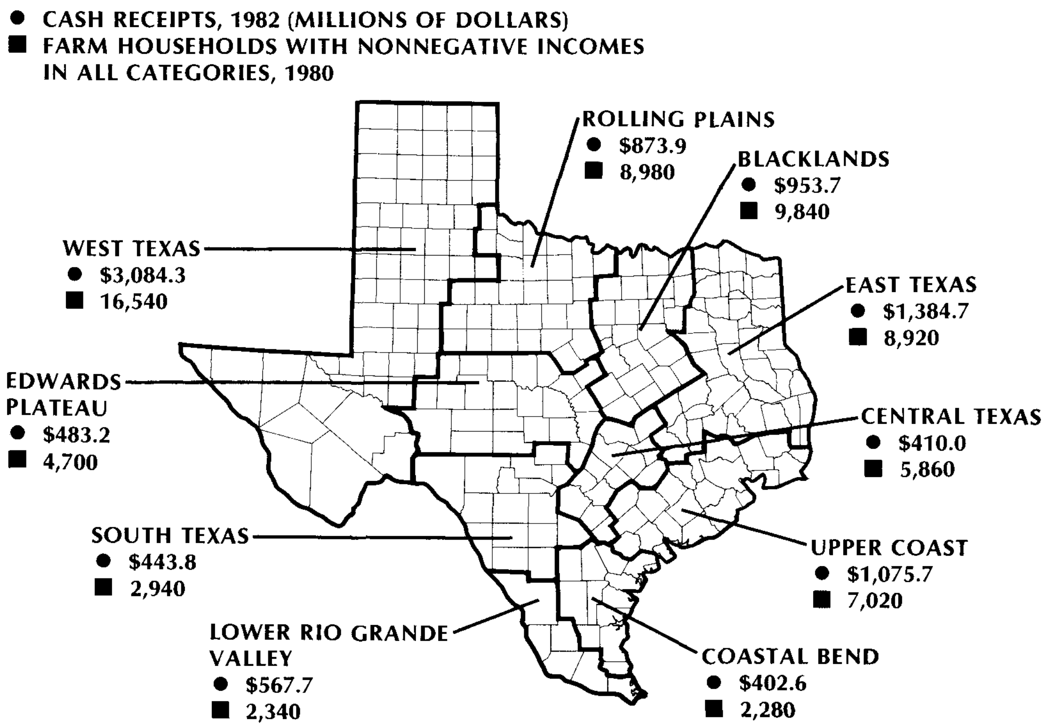
Map 1

Agricultural Cash Receipts and Farm Households in Texas Agricultural Production Regions



SOURCES OF PRIMARY DATA: Texas Department of Agriculture and U.S. Department of Agriculture.
U.S. Bureau of the Census.

sources, savings and investments, temporary or alternative employment, and, to some extent, government programs.

Farmers are not able to take advantage of all the means available to cushion income disruptions. Most agricultural production is by farmers working full-time. These farmers are considered to be small businessmen. It is assumed that the full-time farmer is sufficiently engaged on the farm or ranch to preclude taking a job elsewhere. The part-time farmer or rancher is more diversified and, consequently, better able to withstand an agricultural income disturbance. To the extent that a region is composed of part-time farmers, it is more insulated from agricultural shocks. But for full-time farmers, temporary wage employment is not an alternative.

Agricultural emergency programs of the Federal Government do help farmers ride out economic disruptions, but most of the aid is available as loans to farmers who cannot qualify for commercial credit. Many farmers do not seek emergency government loans, as shown by Texas farmers' slight interest in Farmers Home Administration loans during last year's drought. In addition, agricultural programs are relatively uniform across the state, so regional differences because of government programs are small. In the event of an agricultural income disturbance, farmers frequently cannot rely on the option of outside employment nor on government agricultural emergency programs.

Absolute versus proportional shocks

The capacity to withstand a disruption in agricultural income depends on a farm household's income diversification, its level of income, and, in some cases, diversification of agricultural enterprises. The last category is assumed to be of minimal importance because the disturbances are assumed to be agriculture-wide. Drought, for example, affects most farm and ranch enterprises. As long as the income flows are not perfectly correlated, income diversification can help.[1]

Shocks to agricultural income are generally of two types: proportional and absolute. If cotton farmers in West Texas suffer a 10-percent yield loss because of drought, that is a proportional loss. An example might be two farmers who suffer a 10-percent yield loss in their cotton crops, one with 200 acres and the other with 2,000 acres. Each has suffered the same proportional loss of cotton income. An example of absolute loss would be spotty hail or insect damage. Considering the same pair of farmers, suppose that the first suffered a $10,000 "hail-out" while the second had $10,000 of weevil damage. In this case, the amount of loss is the same, but the loss as a percentage of gross income from farming is very different.

For a discussion of proportional versus absolute losses, see the accompanying box. In the event of a proportional loss, the farm household with more diversified income sources is better off than a household with income largely from farming. For absolute losses, the wealthier the household, the better.

Most agricultural income losses are probably a mixture of both, with the proportional type predominating. This article primarily examines the capacity of farmers in different regions to withstand proportional agricultural income shocks, although some wealth comparisons are made.

The data

Household income data for analyzing the Texas regional differences in capacity to withstand agricultural income shocks were provided by the Public-Use Microdata Sample of the 1980 Census of Population and Housing.[2] The income of all members of the household was summed by the seven income categories provided by the Census data. The income categories are wages and salaries; nonfarm self-employment; farm self-employment; interest, dividends, royalties, and rental payments (hereafter referred to as interest and dividends); social security; public assistance; and all other income. These income sources were also calculated as percentages of total household income.

The original subset sample totaled 4,132 households. Two sets of deletions were made. One set was households without any income in any category, of which there were 42. These households could not be affected by income losses. Another 619 households were deleted because they had negative income in at least one category. This second set of deletions was necessary to rule out "tax farmers"—individuals showing losses from farming year after year—and to exclude individuals having a single bad year. Every year, some genuine farmers lose money, but the composition of this group constantly changes.[3]

1. The risk to a household's total income can be reduced if the income flows do not have a correlation coefficient of 1. Differences in the income categories make it clear that such perfect correlation does not occur here. Social security payments, interest and dividends, nonfarm wages and salaries, and farm self-employment income should not show high positive correlation; for example, a large drop-off in farm income will not be associated with a large fall in social security income.

2. Of the three available Census samples, the 5-percent A Sample was used. On the basis of the county groups identified in that sample, Texas was divided into production regions. The 5-percent sample was subset by selecting only households that contained at least one member with farmer occupation code 473, "farmers, except horticultural." The income reported in the 1980 Census was earned in 1979, one of the best years overall for farm income in Texas. Because all regions did not report record cash receipts that year, using 1979 created some bias. For example, West Texas had a greater share of Texas agricultural receipts in 1979. Thus, income comparisons will be more dramatic than if other years were used. Overall though, every region had a good year, so the use of 1979 income data should not distort the results substantially.

3. There were methodological reasons for exclusion. Including negative income categories naturally reduces total household income, so calculating category income as a percentage of total household income can result in shares over 100 percent or shares that are negative. To make the sample more typical, households with negative income categories were dropped.

The model

The purpose of this article is to determine whether regional differences exist in the capacity of households to withstand agricultural disruptions, not to investigate the causes of those differences. To quantify the differences, regression analysis could be used or a very simple linear one-factor analysis of variance (ANOVA) model. Regression analysis would employ the dependent variable regressed on a constant and a series of dummy variables representing the regions. Analysis of variance is, in a sense, a subset of regression analysis but is formulated specifically to deal with the types of comparisons undertaken here, making the ease of use considerable.

Results equivalent to those from using regression were obtained from a simple ANOVA model:

(1) $$y_{ij} = \mu + \tau_j + \varepsilon_{ij},$$

where y is the observation. The different observation classes used were income category, income category as a percentage of household income, and household income. In this model the observation is composed of three parts: the Texas mean, or grand mean, for the variable is μ, the "treatment effect" stemming from regional differences from the grand mean is τ, and the random component is ε. Because the treatment effects are modeled as deviations from the grand mean, they sum to zero. The subscript i is the observation, while the subscript j is the region.

Regional means (μ_j) are the sum of the grand mean and the treatment effects:

(2) $$\mu_j = \mu + \tau_j.$$

Substituting equation 2 into equation 1 gives the estimating equation:

(3) $$y_{ij} = \mu_j + \varepsilon_{ij}.$$

Variation within a region is caused by all the unexamined factors that make up the random error term. Variation between regions is the result of any differences in regional means *and* random variation.

The first hypothesis tested, using an F statistic, was whether there are any differences between regional means.[4] If there are no differences in mean income levels by category, category shares of household income, or mean household income, then the regions are equally capable of withstanding a disturbance in their agricultural income. The null hypothesis was

(4) $$H_0: \mu_1 = \mu_2 = \mu_3 = \ldots = \mu_{10},$$

against the alternative that the regional means are not all equal. Recalling equation 2, this is equivalent to testing whether the individual regional effects (τ_j) are significantly different from region to region.

Further comparisons were made between individual regional means using t tests.[5] The mean of every region was compared with the mean of each of the other regions. Such null hypotheses were

(5) $$H_0: \mu_j = \mu_k \quad \text{for all } j \neq k, \quad j,k = 1,10,$$

against the alternative sets of hypotheses that the

4. An F statistic can be constructed from a ratio of the mean square of the variation between regions (which includes the variation from random sources and variation from different treatment effects) to the mean square of variation within all the regions (which results from only the random error term):

$$\frac{\sum_{j=1}^{10} n_j(\mu_j - \mu)^2}{\sum_{j=1}^{10} \sum_{i=1}^{n_j} (y_{ij} - \mu_j)^2},$$

where the summation j equals 1 to 10 is over regions and the summation i equals 1 to n_j is over observations. The number o observations in the jth region is n_j.

The value of the F statistic depends on the variation from differences in regional means or treatment effects. If the regional means are the same, the F statistic will be close to 1, and the null hypothesis cannot be rejected. Conversely, if the treatment effects are significantly different, the between-region variation will be larger than the purely random within-region variation, and the F statistic will be significantly larger than 1.

5. A t statistic can be calculated to test whether the difference is significant:

$$\frac{\text{abs}(\mu_j - \mu_l)}{\text{RMSE}(1/n_j + 1/n_l)^{1/2}},$$

where abs is the absolute value, RMSE is the root mean square error, and n_j and n_l are the number of observations for regions j and l.

means are different from each other.[6]

Finally, use of the *F* and *t* tests through the ANOVA technique required that the random error terms be normally distributed and the variances in each region be the same. The samples tested out to be normally distributed, and the variances were shown to be similar for most comparisons.[7]

Texas farm households

Wealth and diversification are two criteria for measuring the capacity of Texas farm households to withstand agricultural shocks. Table 1 shows that there is considerable dispersion in terms of categorical income distribution, as well as shares of total household income, and in the variance in aggregate household income. For example, farm self-employment income ranges from an average high of $16,732 in West Texas to a low of $7,995 in East Texas. On average, Texas farm households earn less than half, 47 percent, of their money from farming. This proportion varies widely by region, from a high of 58 percent in West Texas to a low of 38 percent in East Texas and the Blacklands. Total household income varies from a low of $20,013 in Central Texas to almost 50 percent more in the Coastal Bend. Overall, if all Texas farm households had the same income distribution as did the average Texas farm household in Table 1, the impact of agricultural income stream disruptions would be far less.

The categories accounting for three-quarters or more of the farm household incomes are farm self-employment, wages and salaries, interest and dividends, and social security payments. The analysis presented here concentrates on these four categories and ignores the minor categories of non-farm self-employment, public assistance, and "other income."

Are the magnitudes of income and categorical percentages statistically significant for the regions? If so, how were these regions ranked? Following the ANOVA test procedures, both *F* tests and *t* tests were performed for each of the major income categories. Presentation of the *F*-test results is straightforward because an *F*-statistic value shows for a particular income category whether regional means as a group differ from the mean for Texas.

The *t* test elaborates on the group results shown by the *F* test.[8] For the four income categories, *t* tests were performed between all possible regional pairings within each category: every region was tested against nine others. The results of these *t* tests were then aggregated and linked. For example, if one region proved to have a significantly greater mean in one income category than all the other regions, it was given a score of +9. Similarly, if one region had a significantly lower mean than five other regions, it would score a −5. For a particular income category, therefore, a region might score +2/3/−4, which means that it had a higher mean than two other regions, was not statistically different from three regions, and had a lower mean than four others.

Taking the diversification issue first, it is contended above that a diversified region is better able to deal with proportional agricultural income shocks

6. These simple *t* tests are not as reliable as in regression analysis. Given that there are 10 regions, to test each region against all the other regions involves 45 comparisons. In single *t* tests, as in a regression, the possibility of error is not of as much concern. For example, at the 95-percent significance level, the *t* test will reject a false null hypothesis 95 times out of 100. For a single *t* test, those are good odds that the *t* test is indicating significance of the null hypothesis correctly. If multiple *t* tests are conducted, as between all pairs of regions, then the probability that the *t* test will reject the true null hypothesis—a Type I error—increases as the number of *t* tests conducted goes up. For the 45 comparisons the probability of making at least one Type I error is close to .90.

7. All the samples were normally distributed according to a modified version of the Kolmogorov-Smirnov *D* statistic, which was calculated for each income and percentage category for each region. See M. A. Stephens, "Use of the Kolmogorov-Smirnov, Cramér-Von Mises and Related Statistics Without Extensive Tables," *Journal of the Royal Statistical Society*, ser. B, 32, no. 1 (1970): 115–22.

   The regional variances for each major income category were tested against the variance of the entire Texas sample. It was found that 55 percent of the sample statistics could not be rejected and over 71 percent were within 110 percent of the critical value. Most of the statistics that did reject the null hypothesis were from the mean income categories, rather than the more critical categories based on the percentage of total income. Although compliance with the equal variance requirement was not strict, the bias that is introduced should not do much violence to the aggregated results.

8. The *t* test, in a strict sense, does not need to be consistent with the *F* test. That is, even though the *F* test may reject the hypothesis that the regional means as a group are different from the state mean, the *t* test can still pick out significant differences between individual regional means.

Table 1
FARM HOUSEHOLD INCOME IN TEXAS, BY CATEGORY
(Averages, based on 1979 data)

Agricultural production region	Wages and salaries	Self-employment		Interest and dividends	Social security	Public assis-tance	Other income	TOTAL INCOME
		Nonfarm	Farm					
West Texas	$4,153	$1,347	$16,732	$3,734	$1,004	$ 49	$ 534	$27,554
Percent of total	16.3	4.1	57.9	11.3	7.3	.6	2.5	
Rolling Plains	4,948	909	14,417	3,362	1,427	46	873	25,982
Percent of total	17.0	3.5	52.7	11.2	12.0	.7	3.0	
Blacklands	4,876	1,077	8,838	3,133	1,793	145	1,183	21,046
Percent of total	19.9	4.0	37.9	12.5	17.4	2.0	6.5	
East Texas	5,372	1,221	7,995	2,927	1,673	122	1,096	20,407
Percent of total	21.6	4.8	38.1	11.5	16.6	2.3	5.1	
Upper Coast	5,213	1,626	11,885	4,549	1,311	110	534	25,229
Percent of total	19.8	4.5	44.9	14.5	13.1	.7	2.5	
Central Texas	4,835	1,101	8,080	2,976	1,807	120	1,094	20,013
Percent of total	17.8	3.0	39.8	14.1	18.9	1.2	5.2	
Coastal Bend	4,686	1,207	14,844	6,136	1,808	88	331	29,101
Percent of total	17.7	2.8	48.8	15.7	12.8	.4	1.9	
Lower Rio Grande Valley	4,594	1,078	8,688	5,118	1,486	393	594	21,951
Percent of total	24.3	3.0	38.8	11.1	13.3	5.8	3.7	
South Texas	4,428	1,364	9,247	4,518	1,039	88	1,373	22,056
Percent of total	19.3	3.6	45.9	16.2	8.2	.9	5.8	
Edwards Plateau	3,376	1,906	11,760	2,769	1,429	95	521	22,257
Percent of total	14.8	4.4	52.1	10.7	14.3	1.5	2.2	
TEXAS	$4,698	$1,268	$11,984	$3,609	$1,428	$101	$819	$23,908
Percent of total	18.4	3.9	47.2	12.3	13.0	1.3	3.8	

SOURCE OF PRIMARY DATA: U.S. Bureau of the Census.

Table 2
SIGNIFICANCE TESTS OF SHARES OF FARM HOUSEHOLD INCOME IN TEXAS REGIONS, BY MAJOR CATEGORIES

Agricultural production region	Wages and salaries	Farm self-employment	Interest and dividends	Social security
West Texas	0/5/−4	+9/0/0	0/5/−4	0/1/−8
Rolling Plains	0/7/−2	+6/2/−1	0/6/−3	+1/5/−3
Blacklands	+2/7/0	0/3/−6	0/9/0	+4/5/0
East Texas	+3/6/0	0/3/−6	0/7/−2	+4/5/0
Upper Coast	+2/7/0	+2/4/−3	+4/5/0	+2/4/−3
Central Texas	0/8/−1	0/5/−4	+1/8/0	+7/2/0
Coastal Bend	0/9/0	+4/4/−1	+3/6/0	+1/7/−1
Lower Rio Grande Valley	+4/5/0	0/5/−4	0/9/0	+1/7/−1
South Texas	0/9/0	+2/5/−2	+4/5/0	0/4/−5
Edwards Plateau	0/5/−4	+5/3/−1	0/6/−3	+2/6/−1
F-test statistic	2.70	19.17	2.21	11.92

NOTE: The income share, by category, for each region was tested against each of the income shares of the other nine regions by individual *t* tests within the same category. The results of the nine tests are grouped as to whether the region's categorical income share is statistically larger (the first number in each entry), no different (the second entry number), or statistically smaller (the last number in each entry) than the same categorical income shares in other regions.

than are regions with incomes more concentrated in farming or ranching. Table 2 gives the *F*-test statistic for each major income category as a percentage of total household income.

Several things are of interest in Table 2. First, all the *F* statistics are significant at the 10-percent level. Also, the regions with farm household incomes concentrated in farming are easy to spot. The *t*-test scores indicate that West Texas farm households earn a larger portion of income from farming, in a statistical significance sense, than do households of any other Texas region. In the areas of diversification out of farming—wages and salaries, interest and dividends, and social security payments—scores are markedly lower for West Texas than for many other regions. Other regions with heavy farm income concentrations are the Rolling Plains and the Edwards Plateau, with scores of +6 and +5, respectively, under farm self-employment income. Both regions are at a disadvantage in most of the diversification categories.

Regions having the lowest percentage of their household income in farm self-employment are those that obviously earn more income in the nonfarming categories and are more diversified. Table 2 reveals that the Blacklands and East Texas farm households have noticeably lower mean percentages of income from farming. Both regions have similar distribution patterns in the other income categories—no noticeable diversification in interest and dividends but strength in social security payments and some strength in wages and salaries. Central Texas and the Lower Rio Grande Valley also show signs of diversification, with low percentages of income from farming and some significant mean percentages in social security and wages and salaries. The remaining regions show a mixture of concentration and diversification.

Overall, Table 2 contains the following information. West Texas, the Rolling Plains, and the Edwards Plateau are the least diversified regions, while the Blacklands and East Texas head the list of the most diversified. This diversification is the household income attribute that largely cushions the impact of proportional agricultural income disruptions.

Table 3
SIGNIFICANCE TESTS OF MEANS OF SELECTED FARM HOUSEHOLD INCOME IN TEXAS REGIONS

Agricultural production region	Major categories: Wages and salaries	Farm self-employment	Interest and dividends	Social security	Total income
West Texas	0/8/−1	+8/1/0	0/8/−1	0/1/−8	+6/3/0
Rolling Plains	0/9/0	+7/1/−1	0/8/−1	+1/6/−2	+4/5/0
Blacklands	0/9/0	0/4/−5	0/7/−2	+5/4/0	0/5/−4
East Texas	+2/7/0	0/4/−5	0/6/−3	+3/6/0	0/5/−4
Upper Coast	0/9/0	+4/3/−2	+3/6/0	+1/4/−4	+3/6/0
Central Texas	0/9/0	0/4/−5	0/8/−1	+4/5/0	0/5/−4
Coastal Bend	0/9/0	+5/4/0	+6/3/0	+3/6/0	+6/3/0
Lower Rio Grande Valley	0/9/0	0/5/−4	+2/7/0	+1/8/0	0/7/−2
South Texas	0/9/0	0/6/−3	0/9/0	0/5/−4	0/7/−2
Edwards Plateau	0/8/−1	+3/4/−2	0/6/−3	+1/7/−1	0/6/−3
F-test statistic	1.03	19.98	2.17	7.20	7.58

NOTE: The mean of income, by category, for each region was tested against each of the means of the other nine regions by individual *t* tests for the same category. The results of the nine tests are grouped as to whether the region's mean is statistically larger (the first number in each entry), no different (the second entry number), or statistically smaller (the last number in each entry) than the means for the other regions.

Farm household wealth

To complement the examination of income categories as a percentage of total household income, farm household wealth was considered. The Census does not offer data on household wealth other than the value of the house. However, the means of income categories were compared across regions, instead of comparing the categorical average shares of household income. Many farmers are land-rich and cash-poor. In that sense, the proper wealth measure of capacity to withstand agricultural income disruptions may be levels of income and returns from assets rather than equity in land and machinery. Equity-backed loans might be an alternative, but with agricultural income disturbances often creating cash-flow difficulties for farmers and ranchers, loans on equity would only threaten longer-term viability.

Table 1 reveals considerable differences in mean total household income and in the various components of income. If total household income is used as a proxy for wealth, the Coastal Bend is the wealthiest region while Central Texas is the least well off. The interest and dividends column shows a wide variation in income, from a low in the Edwards Plateau to a high in the Coastal Bend region. Similar differences exist in other categories.

To examine whether the means of the income categories are significantly different from one another on a regional basis, ANOVA results are presented in Table 3 in a manner identical with Table 2. Only the regional mean incomes from wages and salaries do not seem different from the Texas average, as judged by an insignificant F statistic at the 90-percent level. All the other income categories have regional means significantly different from the respective statewide means. In withstanding an absolute agricultural income disturbance, it is household wealth or a proxy, household income, that is important. It is also likely that interest and dividends income is proportional to some measures of wealth, so such income can be considered a proxy as well.

Comparing the means of household income in Table 3, the West Texas, Coastal Bend, and Rolling

Proportional and Absolute Losses to Farm Income

Particular assumptions about farmers' utility functions are necessary in order to use the proportional-absolute loss analysis. It is assumed that all farmers and ranchers have the same utility function. Further, so that farmers react to proportional losses the same way, it is assumed that this utility function has decreasing marginal utility of income and a constant elasticity of utility with respect to income, with a value between zero and 1. The simple utility function was

(1) $$U = Y^{\alpha},$$

where *U*, *Y*, and α are utility, income, and the elasticity of utility, respectively.

Taking the first derivative gives the marginal utility of income,

(2) $$\frac{\partial U}{\partial Y} = \alpha^{*}Y^{(\alpha - 1)}.$$

For marginal utility to be decreasing, its derivative must be negative. Because

(3) $$\frac{\partial^2 U}{\partial Y^2} = \alpha^{*}(\alpha - 1)^{*}Y^{(\alpha - 2)}$$

and α is assumed to be between zero and 1, marginal utility is decreasing as income is increasing.

Utility functions with constant elasticity of utility, coupled with decreasing marginal utility of income, give farmers the same loss of utility for a proportional loss of income. To derive the elasticity of utility (ε_Y) from the above utility function, logarithms were taken of both sides; then the derivative of log of utility with respect to log of income was taken:

(4) $$\varepsilon_Y = \frac{\partial \ln(U)}{\partial \ln(Y)} = \alpha.$$

For expositional purposes, the elasticity of utility can be expressed as

(5) $$\alpha = \frac{\Delta U/U}{\Delta Y/Y},$$

where ΔU and ΔY are discrete changes in utility and income, respectively. Rearranging terms shows that the change in utility depends on the elasticity of utility, total utility, and the percentage change in income:

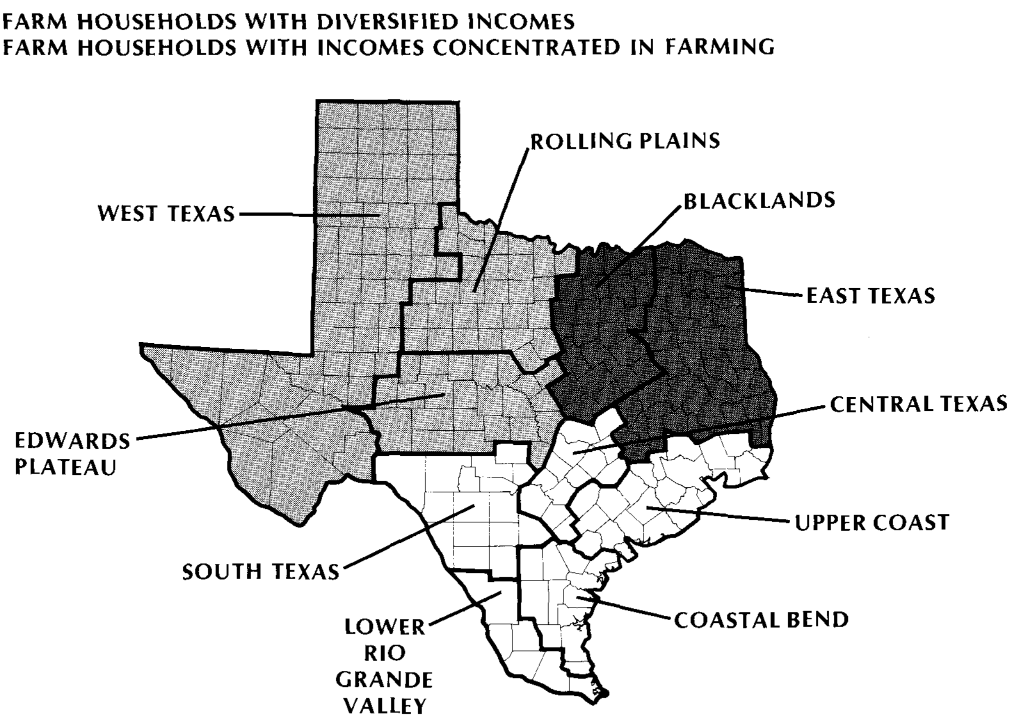
(6) $$\Delta U = \alpha^{*}U^{*}(\Delta Y/Y).$$

Assuming that total utility and elasticity of utility are the same for all farmers, then for a given proportional income change, farmers' utility will change by the same amount. The implication is that a farmer with losses of $7,000 and total income of $70,000 suffers the same reduction in utility as does the farmer with losses of $70,000 and income of $700,000. Proportional losses hurt the same.

By the same token, absolute losses give the above farmers different values of ($\Delta Y/Y$) and different reductions in utility. For example, losses of $10,000 result in a utility loss 10 times greater for the farmer with the smaller income. For absolute losses, the wealthier the farm household, the better.

Map 2

Income Diversification Within Farm Households Across Texas Agricultural Production Regions



Plains regions do have higher household incomes than many other regions, and these differences are statistically significant. But West Texas and the Rolling Plains have the largest absolute incomes from farming and, hence, are in a position to lose the most in the event of a proportional agricultural shock. The Coastal Bend, with a +6 score, is the only region that stands out in the comparison of mean levels of interest and dividends, suggesting that this region's household portfolio of income-producing assets is measurably larger than for most other regions.

Table 3 also reveals which regions have significantly smaller mean household incomes. Central Texas and South Texas have lower mean household incomes, as do the Blacklands and East Texas. These four also have lower incomes from farming. The less diversified regions show only modestly lower means in the nonfarm-originating income categories.

Interestingly, the most diversified regions (the Blacklands and East Texas) seem to be the least able to withstand an absolute income disturbance because the households in those regions are less wealthy than in other regions in Texas. The more concentrated regions (West Texas and the Rolling Plains) are wealthier than the more diversified regions and, thus, are better equipped to handle an absolute agricultural income disruption.

Risk and shock-bearing capacity

From the above analysis, it is clear that West Texas, the Rolling Plains, and the Edwards Plateau are the regions most dependent on farm income, while the Blacklands and East Texas are the most diversified regions. In terms of withstanding proportional shocks to agricultural income streams, the areas that are most dependent on farm income are the least able to cope. For less likely absolute income shocks, the tables are turned. Farming-dependent households generally earn more income, implying that their wealth—and therefore their capacity to withstand an absolute shock—is larger. The other regions in the state (Central Texas, the Coastal Bend, the Valley, the Upper Coast, and South Texas) lie somewhere between the two groups above in their ability to withstand the different kinds of shocks. Map 2 shows the different regions and their capacity to withstand proportional shocks.

Given the relative capacity to withstand agricultural income disruptions, which regions are more likely to experience a disturbance in their farm income streams? Weather events, unlike many price effects or policy actions, are one type of shock that is often both specific to regions individually and important to all types of agriculture.

Texas is a borderline state in terms of climate. The eastern part of the state receives agriculturally bountiful amounts of rainfall—over 40 inches a year—while the western part is bone-dry much of the time, with as little as 8 inches of yearly rainfall. West of a line through Fort Worth, Waco, Austin, and Corpus Christi lies the portion of Texas that gets less than 30 inches of rain a year. Drought becomes a frequent, if not regular, event in these drier parts of Texas. For example, based on the Palmer Drought Severity Index, the Edwards Plateau experienced at least moderate drought in about half of the past 50 years.[9] Moreover, West Texas, the Rolling Plains, and the Edwards Plateau are prone to hailstorms, early frosts, and other weather calamities.

The results so far present one conundrum: farmers and ranchers in the western half of Texas apparently choose to work in an area with large weather-related risks and yet are less able to deal with proportional agricultural income disturbances. But farmers may have several choices in dealing with the risks imposed on West Texas agriculture. The risks can be reduced through risk spreading or neutralizing the risk or can be compensated with risk premiums.

Risk spreading occurs, probably implicitly, when a household pools the income from all its members, then allocates shares. If farming is the riskiest of all the income categories, risk spreading reduces the risks faced by the farmers in the household. Wages and salaries of household members are generally the largest source of nonfarm income for such farm households. But the western half of the state is sparsely settled, and density of nonfarm economic activity is much lower than in the rest of the state. This lower density of economic commerce precludes extensive off-farm employment of household members, making the household more dependent on agriculturally based income.

If risk spreading is not viable, the farmer has strategies for effectively neutralizing the risk. Knowing that drought, for instance, is a long-run fact of life has most likely spurred successful farmers and ranchers to formulate agricultural management plans that take into account periodic dry-outs. For example, rain-fed agriculture may use standby irrigation to augment rainfall during abnormally dry periods.

If the farmers are risk-averse, as assumed throughout this analysis,[10] then economic theory would suggest that in order to bear increased risk, farmers must receive a premium over what farmers receive in less risky regions. Such premiums may be higher returns to agricultural endeavors, or they may be nonpecuniary rewards. Examination of cash receipts and farm household self-employment income shows that the western half of Texas earns more gross farm income,[11] and part of this return could be considered a risk premium.

If the risks were not compensated by additional returns from farming, the farmers and ranchers might be expected to have an incentive to move elsewhere. Migration theory predicts that as long as the discounted benefits from moving outweigh the

9. The Palmer data were provided by Meteorologist George Bomar of the Texas Department of Water Resources.

10. In the discussion of proportional losses, the box covers the assumption of declining marginal utility of income. Declining marginal utility of income also implies risk aversion.

11. Net farm income is a preferable measure to cash receipts but is not available on a regional basis. Both net farm income and cash receipts are proxies for return on investment.

discounted costs, there is an incentive to migrate.[12] The obvious costs to a farmer in moving to a less risky region are that much farm and ranch capital may be suitable only for particular operations and a farmer's skills and knowledge are likely the products of several generations of specialized experience. These are not traded lightly for the uncertainties of a different way of farming, and the costs may be so high that there are no short-run methods of financially surviving the sure losses that initially result from a change in environment.

Recent developments in migration theory also suggest that migration takes place as a result of fixed-place amenities or compensating differentials. Incomes and rents adjust to keep utility constant over geographic areas.[13] These compensating differentials may include intangible aspects that reward farmers for staying in higher-risk farming areas. Farmers generally consider farming and ranching to be a way of life, one in which the land, tradition, and the part of the country they live in play important roles in their utility functions.

Conclusions

The major finding of this article is that different regions of Texas are not uniformly able, *on average,* to deal with shocks to agricultural income. For example, while farmers in the Lower Rio Grande Valley did suffer large losses in the Christmas freeze, the average farm household there receives only 39 percent of its income from farming and 24 percent from wages and salaries. In contrast, the drought in West Texas over the last two years has hit households that are very concentrated in farming, which accounts for 58 percent of total household income there.

Farmers in the western portions of the state face a riskier environment. Because the risks cannot be spread, farmers are likely to be compensated with higher monetary returns and nonpecuniary differential compensation factors—largely the intangible aspects of farm life that are very difficult to quantify but are highly important in a household utility function.

12. See Larry A. Sjaastad, "The Costs and Returns of Human Migration," *Journal of Political Economy* 70 (October 1962, pt. 2): 80–93.

13. See Philip E. Graves, "Migration with a Composite Amenity: The Role of Rents," *Journal of Regional Science* 23 (November 1983): 541–46.

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