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This study presents new evidence on the relationship between capacity utilization and inflation in order to provide a proper framework for understanding the complexities involved. Because empirical results suggest that capacity utilization and changes in inflation are jointly endogenous, much of the previous work in this area may suffer from simultaneity bias. Using a two-equation structural model, the author finds support for a "steady-state" rate of capacity utilization of about 81.5 percent. While that figure is in line with previous estimates, this model does not suffer from simultaneity bias and appears to be stable over time.

The relationship between the performance of the financial sector and economic activity has received increasing attention from economists during the past decade. Empirical studies generally support the view that financial structure and stress can have real economic effects. This paper explores the impact of financial restructuring on economic activity by using an alternative data set that in some respects more completely measures change in the local banking sector than do data used in previous research. Results suggest that the deaths of midsized banks have a negative but short-lived impact on economic activity.

Currently, the Federal Reserve is being urged to adopt price stability or an explicit price-index target as its primary long-term monetary policy objective. The purpose of this paper is to ascertain the short- and long-term implications of an inflation policy for real output. Inflation policy is defined here in terms of a series of innovations that exclusively determine trend inflation. To estimate this series, the authors adopt a recently developed method that allows structural interpretation of a simple vector-autoregression and apply it to a macroeconomic system that includes real output and inflation. Results suggest that the benefits of a monetary policy aimed at achieving gradual disinflation would probably outweigh the costs.

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Coordinating Economist:
Randall W. Eberts

Editors: Tess Ferg
Robin Ratliff
Design: Michael Galka
Typography: Liz Hanna

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A Reexamination of the Relationship between Capacity Utilization and Inflation

by Paul W. Bauer

Paul W. Bauer is an economist at the Federal Reserve Bank of Cleveland. The author would like to thank John B. Carlson, Randall W. Eberts, Stephen E. Haynes, and John A. Tatom for their helpful comments. Thomas Kluth provided valuable research assistance at the beginning of this project.

Introduction

The Federal Reserve places a high priority on controlling inflation and ensuring full employment of economic resources. Thus, empirical relationships that can better inform policymakers about the prospects of these two key economic variables are eagerly sought. One such relationship that has received attention over the years is that between capacity utilization and inflation (McElhattan [1978, 1985], Tatom [1979], and Gittings [1989]). Although these authors have employed various theoretical and empirical methods, in general they all have found evidence for a "steady state" or "natural rate" of capacity utilization of about 80 percent to 82 percent. Deviations from this rate are directly related to changes in the inflation rate.

Most of these models posit a single equation in which capacity utilization is assumed to be an exogenous variable that explains changes in the inflation rate. However, economic theory and Granger causality tests suggest that both capacity utilization and inflation are endogenous. This empirical finding is incorporated here by construction of a two-equation structural model based on the work of Haynes and Stone (1985). Though parsimonious, the model imposes no explicit macroeconomic world view and yet provides a reasonable fit for the movements of capacity and inflation in the U.S. economy. In addition, the full sample period from 1953 to 1989 can be employed, since no evidence of structural change is found. Although the resulting estimate of the steady-state capacity utilization rate is consistent with previous research, the structural approach permits us to examine the dynamics of the relationship, as revealed through simulated aggregate demand and supply shocks.

I. Theoretical Foundations

Capacity utilization is defined as actual production divided by capacity (section II briefly reviews the problem of adequately defining these terms). The belief that high capacity utilization levels lead to an accelerated rate of inflation is based on the assumption that high capacity utilization levels are related to increasing marginal costs of production in the short run. In the long run, high capacity utilization may prompt new investment, thereby expanding capacity and relieving price pressures.

McElhattan (1978) was the first to develop a model linking inflation and capacity utilization. The model is composed of two basic structural equations, one that relates prices to a markup on unit labor costs, and another that relates wage changes to labor-market excess demand and expected inflation. The markup equation can be written as

$$(1) \quad IR(t) = a_{12}\, W(t) - a_{13}\, T(t) + f\,[CU(t)],$$

where IR is the inflation rate, W is the rate of change of nominal wages, T is the growth rate of labor productivity, and $f\,[CU(t)] = [b_0 + b_1 CU(t)]$ is a measure of excess aggregate demand that is an increasing function of capacity utilization (CU).

The second equation relates the rate of change in nominal wages to the expected inflation rate (IR^*), the growth rate of labor productivity $[T(t)]$, and the excess demand in the labor markets $\{h\,[u(t)]\}$. It can be written as

$$(2) \quad W(t) = a_{21}\, IR^*(t) + a_{23}\, T(t) - h\,[u(t)],$$

where $h(u)$ is a decreasing function in the unemployment rate (u). With this specification, inflation-adjusted wage changes ($W - a_{21} IR^*$) rise in proportion to labor productivity for a given level of unemployment.

Substituting equation (2) into equation (1) and simplifying yields

$$(3) \quad IR(t) = a_{12}\, a_{21}\, IR^*(t) + (a_{12}\, a_{23} - a_{13})\; T(t) - a_{12}\, h\,[u(t)] + f\,[CU(t)].$$

A number of restrictions are imposed by McElhattan in order to estimate equation (3). First, given the high correlation between unemployment and capacity utilization rates, only one of these variables is included.[1] McElhattan argues that retaining the capacity utilization rate is preferable because the natural rate of unemployment may be affected by demographic changes, whereas capacity utilization is not.

Next, a formulation for inflation expectations must be imposed. As is common in price-markup models, inflation expectations are modeled as a weighted average of past inflation. McElhattan finds that only the one-year lag is statistically significant, and that its estimated coefficient is close to one. Because this coefficient must equal one for there to be no long-run Phillips curve type of trade-off between inflation and capacity utilization, this constraint is imposed as well.

Given these restrictions (and a few other minor ones), the reduced-form equation in McElhattan's model can be rewritten as

$$(4) \quad CIR(t) = a\,[CU(t) - CU^e] + v(t), \quad a > 0,$$

where $CIR(t)$ is the change in the inflation rate, $CU(t)$ is the capacity utilization rate, CU^e is the natural rate of capacity utilization, and $v(t)$ is statistical noise. With this formulation, it is easy to see that when $CU(t)$ is larger (smaller) than CU^e, the inflation rate will increase (decrease), and that when $CU(t)$ is equal to CU^e, the inflation rate will remain unchanged. This can be viewed as an output-gap model, with capacity utilization playing the role usually reserved for the unemployment rate.[2]

Gittings (1989) basically follows McElhattan's approach, but argues informally that there are two reasons for the existence of inflationary pressures when capacity utilization is high. First, as capacity constraints are reached, firms are better able to increase their prices in the face of strong demand; however, these same firms' customers may find themselves in a similar position. The second argument is that aggregate-demand growth raises the demand and prices for new capital goods, along with the costs of financing those goods, relatively more when there is less idle capital to employ. Thus, over the business cycle, the rental price of capital rises relative to that of labor.

An entirely different approach is taken by Tatom (1979), who sets up a partial adjustment model in which changes in capacity utilization are the result of monetary surprises. This relationship can be written as

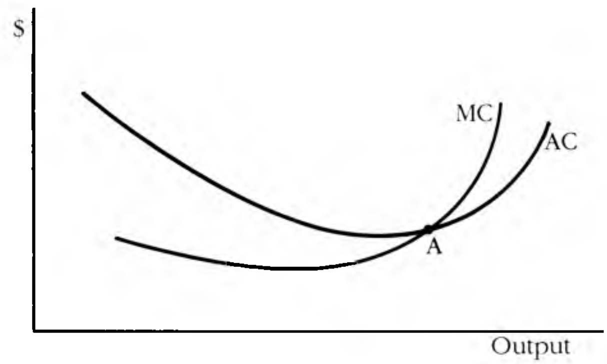
$$(5) \quad dCU(t) = a\,[CU^e - CU(t\text{-}1)] + b\,\{\,m(t) - E_{t\text{-}1}\,[m(t)]\,\},$$

■ 1 The correlation between unemployment and capacity utilization is 0.875 using yearly data from 1953 to 1989.

■ 2 See appendix B in Hallman, Porter, and Small (1989) for an overview of output-gap and price-gap models.

FIGURE 1

Short-Run Cost Curves



SOURCE: Author's calculations.

where $dCU(t)$ is the change in the capacity utilization rate, $m(t)$ is the actual rate of monetary growth, and $E_{t-1}[m(t)]$ is the anticipated rate of monetary growth in the previous period. Here, capacity utilization adjusts to its equilibrium level (CU^e) with a lag, and departures from CU^e occur as a result of monetary surprises.

This model is fundamentally different from those underlying McElhattan's and Gittings's work. In Tatom's model, money causes inflation, and only monetary surprises cause changes in capacity utilization. There is no structural link between capacity utilization and inflation, and the natural rate of capacity utilization is achieved as a result of an absence of shocks.

II. Definition and Measurement of Capacity Utilization

At first glance, the concepts of capacity and capacity utilization are easily defined.[3] Capacity is the potential output that an economic unit (for example, a plant, a firm, an industry, or an economy) can produce during a given period, and capacity utilization is simply actual output divided by potential output. However, these seemingly straightforward definitions gloss over a number of problems, the greatest of which is that they fail to take account of operating costs as output varies. Output can be increased by employing workers and machines for longer hours, but this results in overtime and higher maintenance costs.

One alternative is to define capacity as the level of output at which short-run average cost (AC), total cost divided by output, is minimized (point A in figure 1).[4] This definition has the somewhat peculiar property that an economic unit might produce at a rate greater than "capacity," but it does result in a much more informative measure of capacity utilization. At output levels below capacity (to the left of point A in figure 1), output can be increased without a significant increase in marginal cost (MC), the extra cost incurred to produce one more unit of output. However, when output exceeds capacity (to the right of point A in figure 1), increases in output are associated with more rapid increases in MC. This definition of capacity links capacity utilization with MC, and thus is one conceivable microeconomic foundation for the belief that a connection between price movements and capacity utilization exists.

Unfortunately, economic data do not fall like manna from heaven, but must be painstakingly compiled. In the case of capacity (and hence capacity utilization), the usual data collection and aggregation problems are aggravated because capacity is essentially unobserved—unlike actual output. Another complicating factor is the lack of a generally accepted definition of capacity, as noted above.

Most studies that attempt to relate inflation to capacity utilization employ the Federal Reserve's capacity utilization series. In light of the empirical results presented below, it would be useful to have at least a cursory understanding of how this series is constructed. (For a more detailed discussion, see Raddock [1985, 1990].)

The Federal Reserve's goal is to provide capacity utilization estimates that reflect the same degree of "tightness" over time for a given rate. No primary data is collected, as the Federal Reserve relies instead on annual surveys produced by McGraw-Hill and the Census Bureau, and on various industry sources. Strangely, McGraw-Hill offers no definition of capacity utilization to its survey respondents. The Census Bureau offers definitions for its two measures of

■ 3 For a general overview of the problem of defining capacity, see Bauer and Deily (June 15, 1988). For a more detailed treatment, see Klein and Long (1973), Rasche and Tatom (1977), and Berndt, Morrison, and Wood (1983).

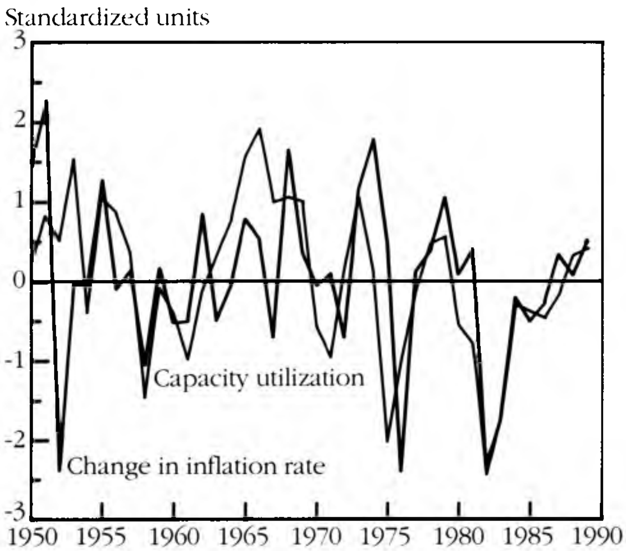
■ 4 Other authors (for example, Tatom [1979] and Rasche and Tatom [1977]) advocate defining capacity utilization as the level of output at which short-run and long-run average costs are equal, so that a firm's demand for any fixed inputs just equals the amount it actually possesses. Although this definition has some theoretical advantages, no organization that produces capacity estimates uses it. Thus, the relative merits of alternative definitions are mentioned only briefly here.

capacity, but most respondents apparently ignore them (see Bauer and Deily [July 1, 1988]).

After a preliminary end-of-year index of industrial capacity is calculated, data are adjusted to remove apparently excessive fluctuations and short-term peak capacity. As a result, capacity figures reflect maximum "sustainable" capacity. Monthly and quarterly estimates are generally straight-line interpolations from past end-of-year estimates and are based on projections of capacity growth for the current year. Because capacity is unlikely to grow at a constant rate throughout the year, the monthly and quarterly estimates should be treated with more caution than the yearly estimates.[5] For this reason, only annual data are analyzed here.

FIGURE 2

Change in Inflation and Capacity Utilization Rates, 1950-88



NOTE: Series were standardized to have a mean of zero and a variance of one.
SOURCES: U.S. Department of Labor, Bureau of Labor Statistics, and Board of Governors of the Federal Reserve System.

TABLE 1

Correlation between the Change in U.S. Inflation and Capacity Utilization Rates and Lagged Capacity Utilization

CIR (t)	*CU (t)*	*CU (t -1)*
1980-89	0.885	0.530
1970-79	0.407	0.825
1960-69	0.461	0.090
1950-59	0.315	–0.439

SOURCE: Author's calculations.

III. Empirical Findings: Puzzles and Possible Explanations

Figure 2 plots changes in the inflation and the capacity utilization rates from 1950 to 1988. It appears that the two series are related—at least indirectly—as manifested by the way in which they tend to move together. Gittings, who only reports results for the 1971 to 1988 period, asserts that there is a fairly uniform one-year lag between the two series. Although this appears to be true for most of the 1970s, it does not seem to apply to the 1980s. Simple correlation among changes in the inflation and the capacity utilization rates confirms this (table 1). Note that during the 1950s there was actually a negative correlation between changes in the inflation rate and the lagged value of capacity utilization. This simple analysis reveals that, whatever the relationship between capacity utilization and inflation, it appears to vary over time.

Figure 3 plots the change in the inflation rate against the capacity utilization rate. Although a straight line appears to fit these data well, clearly a great deal of noise exists in the relationship. Table 2 reports estimates of the McElhattan model (equation [3]) using various sample periods.

Although the point estimate of the noninflationary rate of capacity utilization ranges only from 80.0 percent to 81.9 percent, when the 95 percent confidence interval can be computed, it suggests a range of from approximately 78 percent to 83 percent. The extent to which a change in inflation is associated with a given deviation of capacity utilization from the natural rate (the b coefficient) also appears to vary over time. The "penalty" for a divergence from the equilibrium rate of capacity utilization was 71 percent

■ 5 Gittings (1989) attributes the failure to find any correlation between the change in inflation rates and capacity utilization in the monthly series to noise in the price series. However, the failure could also be a result of problems in the monthly estimates of capacity utilization.

TABLE 2

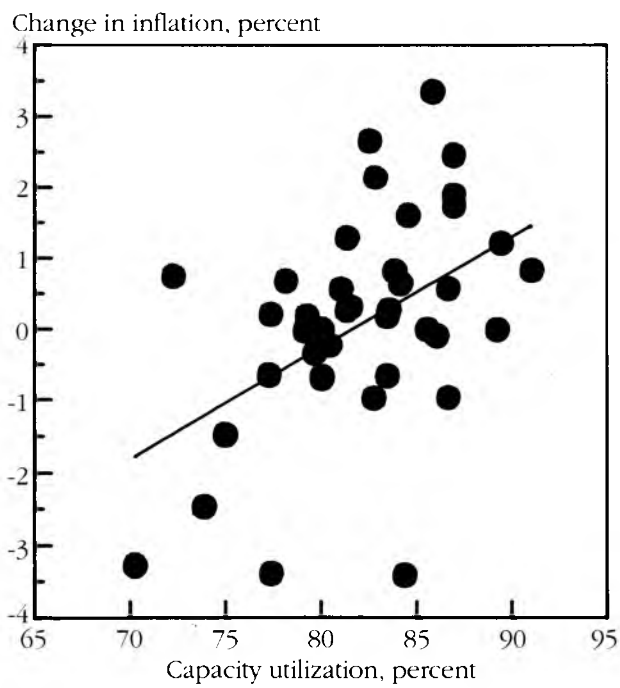
OLS Results for the McElhattan Model Using Various Time Periods

Variable	Full Sample	1950-69	1960-79	1970-89
CU	0.00154	0.00127	0.00106	0.00217
	(3.52)	(1.60)	(1.75)	(3.55)
Constant	-0.125	-0.104	-0.848	-0.174
	(-3.48)	(-1.56)	(-1.68)	(-3.56)
CU^e	81.4	81.9	80.0	80.3
	[78.7, 83.6][a]	—[b]	—[b]	[78.2, 82.5][a]
R^2	0.246	0.125	0.145	0.411

a. 95 percent confidence intervals computed following McElhattan (1978).
b. The procedure fails when both parameter estimates are not statistically significant at the confidence level selected (see Scadding [1973]).
NOTE: $CIR(t) = a + b\,CU(t) + e(t)$. T-statistics are indicated in parentheses.
SOURCE: Author's calculations.

FIGURE 3

Change in Inflation Rate vs. Capacity Utilization



NOTE: Yearly data.
SOURCES: U.S. Department of Labor, Bureau of Labor Statistics, and Board of Governors of the Federal Reserve System.

higher in the 1970s and 1980s than in the 1950s and 1960s, although the difference does not appear to be statistically significant.

If capacity utilization is a true measure of capacity constraints that result in higher inflation, then one would expect the Federal Reserve's capacity utilization series (which covers only manufacturing, mining, and utilities) to predict more accurately the price pressure for goods than for goods and services or just services. The empirical evidence for this conjecture is mixed, however (see table 3).

As expected, the worst fit is found between capacity utilization and the change in the inflation rate for services, although indirect effects are observed. The better fit (at least as measured by the R^2 coefficient and the statistical significance of the coefficients) between capacity utilization and changes in the overall implicit price deflator relative to goods only is unexpected, however. Capacity utilization should be more directly related to changes in the price of goods than to changes in the prices for all goods and services. This finding could be a result of the variance of the goods implicit price deflator being four times larger than the one for services.

At this point, it is appropriate to ask whether capacity utilization should be treated as an exogenous variable. Models can easily be developed in which both capacity utilization and changes in inflation are jointly endogenous. Granger causality tests can then be employed to examine this view, although the results must be interpreted cautiously (see Conway, Swamy, and Yanagida [1983]).

TABLE 3

OLS Estimates of McElhattan's Model Using Various Measures of Inflation

Variable	Goods and Services	Goods	Services
CU	0.172	0.199	0.071
	(4.24)	(3.21)	(1.91)
Constant	-14.0	-16.2	-5.76
	(-4.21)	(-3.21)	(-1.90)
CU^e	81.4	81.7	81.3
	[79.2, 83.4][a]	[78.8, 84.8][a]	[70.6, 89.3][a]
R^2	0.360	0.249	0.105

a. 95 percent confidence interval.
NOTE: Implicit price deflator = 31. T-statistics are indicated in parentheses.
SOURCE: Author's calculations.

TABLE 4

Granger Causality Tests

H_0: Capacity utilization does not Granger-cause changes in inflation

Lag Lengths	lnL_u	lnL_c	Likelihood Ratio[a]
1	-50.34	-56.21	11.74
2	-46.88	-52.99	12.22
3	-42.43	-49.34	13.82
4	-38.47	-45.17	13.40
5	-32.39	-40.55	16.32
6	-26.35	-35.91	19.12

H_0: Changes in inflation do not Granger-cause capacity utilization

Lag Lengths	lnL_u	lnL_c	Likelihood Ratio[b]
1	-88.80	-91.24	4.88
2	-82.39	-84.84	4.90
3	-73.74	-77.25	7.07
4	-65.63	-69.77	8.28
5	-58.12	-64.11	11.98
6	-53.13	-58.99	11.72

a. Statistically significant at the alpha = 0.01 level.
b. Statistically significant at the alpha = 0.1 level.
SOURCE: Author's calculations.

For this study, the tests are performed as suggested by Guilkey and Salemi (1982), with both a time trend and an equal number of lags of the two right-side variables. The number of lags was varied in order to check for robustness (see table 4). Within the framework of the Granger causality test, a variable x does not Granger-cause y if the coefficients on the lags of x are all not statistically different from zero. This hypothesis can be easily examined using a likelihood-ratio test.[6]

The second and third columns indicate the value of the likelihood function of the unconstrained model and the constrained model, respectively (for the latter, coefficients of the lag of the nondependent variable are constrained to equal zero). The last column lists the value of the test statistics and indicates whether the hypothesis of no unidirectional Granger causality can be rejected. These results suggest that there is bidirectional causality between the two time series, although the link from capacity utilization to changes in the inflation rate appears to be stronger (or at least easier to confirm statistically).

Bidirectional causality between capacity utilization and changes in the inflation rate is more consistent with Tatom's approach than with those of McElhattan and Gittings. However, a more complete analysis of the relationship can be obtained through use of a more fully specified structural model that relates the two time series explicitly. The work of Haynes and Stone (1985) provides the basis for one such model.

IV. A Structural Approach

Haynes and Stone construct a model of aggregate demand and supply that is identified by its dynamics: In the short run, quantity sold is demand determined, but price is supply determined. Given this assumption, shifts in aggregate demand trace out aggregate supply, affecting output before prices and leading to an inverse relationship between inflation and lagged unemployment. Haynes and Stone's aggregate supply equation can be written as

$$(6) \quad I(t) = -1/a\, U(t-i) - b/a\, dI(t) + e(t), \quad i > 0,$$

■ **6** The test statistic $2(ln_u - lnL_c)$ is distributed chi-squared with k degrees of freedom under the null hypothesis, where k is the number of constraints placed on the model (the number of lags set to zero).

TABLE 5

Estimates of Aggregate Supply and Demand

Parameter	Estimate	Standard Error	T-Statistic
a	-0.108	0.0306	-3.52
b	0.866	0.0703	12.32
c	0.0013	0.0004	3.48
d	0.151	0.0582	2.61
A	21.7	9.07	2.39
B	-89.0	17.2	-5.17
C	0.716	0.106	6.79
$D1$	104.4	20.0	5.22
$D2$	-33.2	19.6	-1.70
$RHO\,1$[a]	-0.265	0.172	-1.54
$RHO\,2$[a]	-0.308	0.168	-1.83

a. $RHO\,1$ and $RHO\,2$ are the autocorrelation coefficients in equations (8) and (9).
NOTE: Value of the likelihood function = 30.05183.
SOURCE: Author's calculations.

where $I(t)$ is the inflation rate, U is the unemployment rate, d is the difference operator, e is an error term, and a and b are positive constants. $U(t-i)$ is a generalized delay of i periods, and $dI(t)$ is an adaptive representation of inflationary expectations. Supply shocks enter the system through $e(t)$.

Similarly, short-run shifts in aggregate supply trace out aggregate demand. These shifts affect prices before output, leading to a direct relationship between unemployment and lagged inflation. Haynes and Stone model aggregate demand as

$$(7)\quad U(t) = c + fI(t-j) + gU(t-i) + v(t),\ j > 0,$$

where c, f, and g are parameters, and demand shocks enter the system through $v(t)$. The authors assume that the response of unemployment to a supply shock occurs after the price response, so unemployment is related only to lagged inflation.

The Haynes-Stone framework is modified here to use the capacity utilization rate rather than the unemployment rate. It is also augmented to include the lagged value of M2 growth as an explanatory variable in each equation. Inclusion of M2 growth allows for both a varying monetary policy, certainly an important influence on the inflation rate, and links between money and capacity utilization and inflation, along the lines of Tatom (1979).[7]

In both equations, lag lengths of one year for capacity utilization and inflation yield the lowest mean-squared error (the specification criterion employed by Haynes and Stone). In the case of M2 growth, allowing for a lag length of two years in the supply equation and of both one and two years in the demand equation yields the lowest mean-squared error. Equations (8) and (9) are estimated with an allowance for autocorrelation and cross-equation correlations. Results are presented in table 5.

$$(8)\quad ir(t) = a + b\,ir(t-1) + c\,cu(t-1) + d\,gm2(t-2) + e(t)$$

$$(9)\quad cu(t) = A + B\,ir(t-1) + C\,cu(t-1) + D1\,gm2(t-1) + D2\,gm2(t-2) + v(t)$$

Given the dynamic assumptions that identify the model, the coefficient on lagged capacity utilization should be positive in the supply equation (8), and the coefficient on lagged inflation should be negative in the demand equation (9). Both coefficients are of the expected sign and are statistically significant, so reasonable supply and demand relationships appear to have been estimated.

In the short run, faster M2 growth leads to higher inflation and higher capacity utilization (possibly by stimulating aggregate demand, but perhaps through monetary shocks, as suggested by Tatom). However, solving this two-equation system for the long-run steady state indicates that the gain in steady-state capacity utilization is quite modest. In fact, the system can be reestimated with the constraint that M2 growth not affect steady-state capacity utilization by setting

$$(10)\quad D1 + D2 = -Bd/(1-b).$$

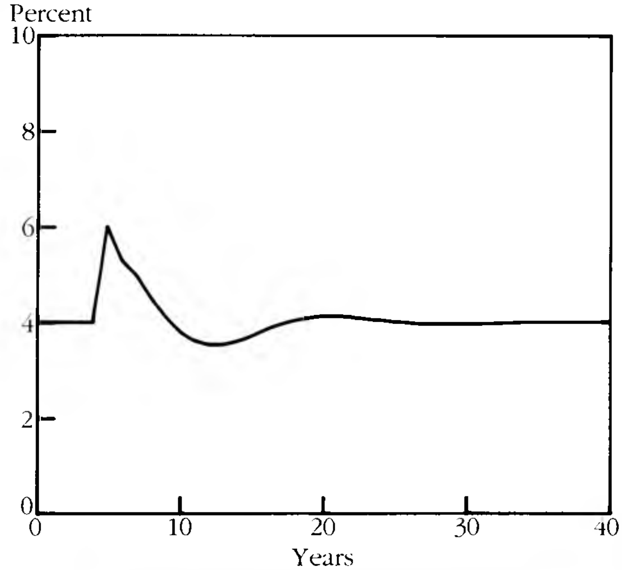
This constraint cannot be rejected at any reasonable confidence level, and thus provides evidence of a natural-rate hypothesis for capacity utilization.[8] This suggests that the Federal Reserve is fairly successful in ensuring that a given capac-

7 Although it is difficult to conceive of a rationale for including M2 growth in the supply equation, the coefficient is statistically significant. Perhaps it influences the real cost of financing in the short run. Alternatively, this two-equation system could be reinterpreted as a VAR model.

8 This hypothesis was tested using a likelihood-ratio test that is distributed chi-squared with one degree of freedom. The test statistic was 0.62, with a 1 percent critical value of 6.645 (the critical value for a 5 percent test was 3.84).

FIGURE 4

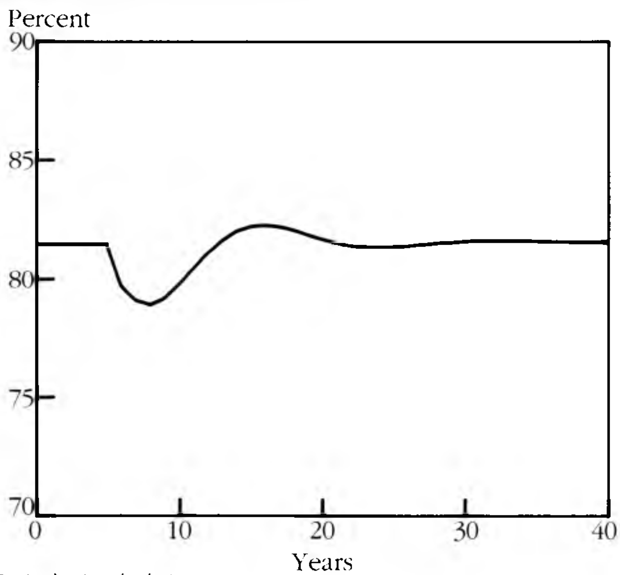
Effect of a Supply-Side Shock on Inflation



SOURCE: Author's calculations.

FIGURE 5

Effect of a Supply-Side Shock on Capacity Utilization



SOURCE: Author's calculations.

ity utilization rate does reflect the same degree of "tightness" over time (unlike earlier research based on models such as equation [3], where the "penalty" for deviations from the steady-state capacity utilization rate varied over time).

Given that M2 velocity is roughly constant in the long run, we would also expect that the steady-state inflation rate would mirror increases in M2 growth.[9] This property is equivalent to imposing

$$(11) \quad c = (1 - C)(1 - b - d)/(B + D1 + D2)$$

A likelihood-ratio test fails to reject this null hypothesis at any reasonable level of significance (the value of the test statistic is 0.32), confirming expectations.[10]

One advantage of this model over the single-equation type represented by equation (4) is that no significant structural change seems to occur over the sample period. This hypothesis was tested by dividing the sample into two periods, 1953-71 and 1972-89, and reestimating the model for each. A likelihood-ratio test was then performed to see whether the null hypothesis of no structural change could be rejected. The chi-squared test statistic with 13 degrees of freedom was 25.0, with a 1 percent critical value of 27.7. At this level of significance, it was found that the null hypothesis cannot be rejected.[11]

The long-term behavior of this two-equation system can now be investigated by solving for its steady-state solution. The steady-state capacity utilization rate is 81.5 percent when M2 grows at a yearly rate of 7 percent—an estimate that is very close to those reported for the single-equation model (tables 2 and 3).[12] The steady-state inflation rate is determined by the growth rate of M2.

The dynamics of the two-equation model can be illustrated through an examination of two simulations. The first introduces a 2 percent supply shock to the supply equation in the tenth period. (This could represent a sudden increase in the price of oil, for example.) The effects of this shock on inflation and capacity utilization are illustrated in figures 4 and 5. Initially, inflation increases while capacity utilization remains unaffected (because only the lags of variables

9 In the steady-state reduced form, inflation is equal to the M2 growth rate minus the growth rate of real output. The model's estimate of average real GNP growth over this period is 2.2 percent—a little less than its 2.9 percent average annualized growth rate estimate.

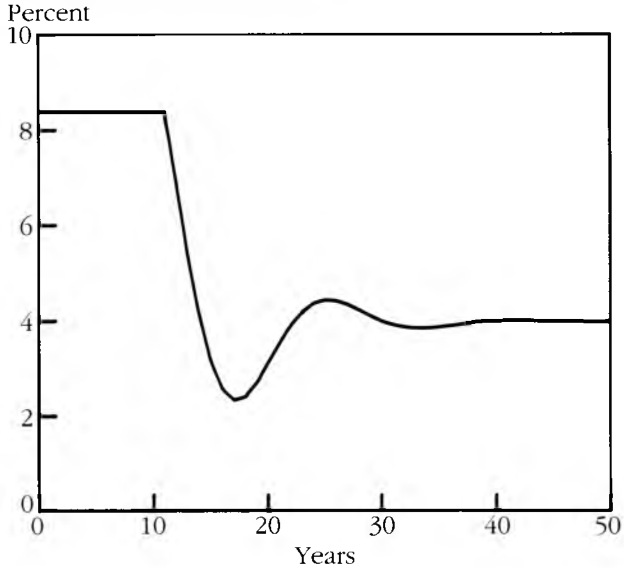
10 An unfortunate feature of this model is that it is impossible to impose simultaneously the constraints that 1) steady-state capacity utilization not depend on M2 growth and 2) steady-state inflation increase in tandem with M2 growth.

11 Although the null hypothesis would be rejected at the 5 percent level (the critical value here is only 22.4), given the large number of parameters that are allowed to vary and the extremely limited number of observations, a relatively tight level of significance is justified.

12 See also McElhattan (1978, 1985), Tatom (1979), and Gittings (1989).

FIGURE 6

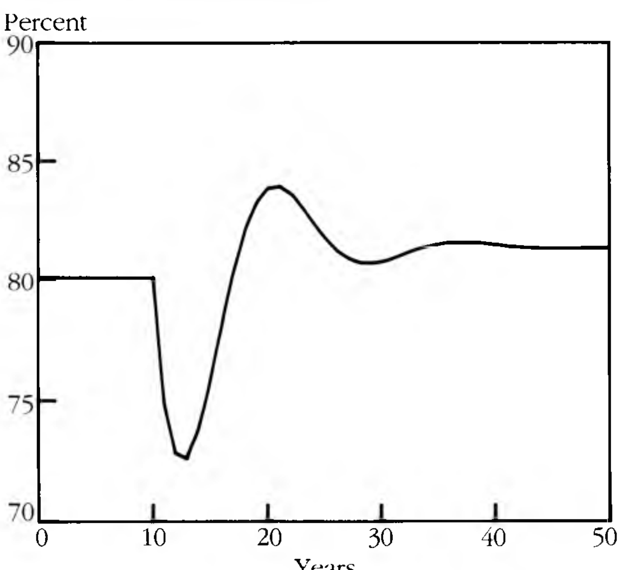
Effect of a Reduction in M2 Growth on Inflation



SOURCE: Author's calculations.

FIGURE 7

Effect of a Reduction in M2 Growth on Capacity Utilization



SOURCE: Author's calculations.

appear on the right side of equations [8] and [9]). In the next period, inflation begins to decline toward its long-run steady state (in part because the shock is no longer present), and capacity utilization decreases because the preceding period's higher inflation reduces demand. Capacity utilization continues to decline for three periods and then rebounds toward its long-run steady state. Although the model does experience some "overshooting," the system is more highly damped than many large macroeconomic models.[13]

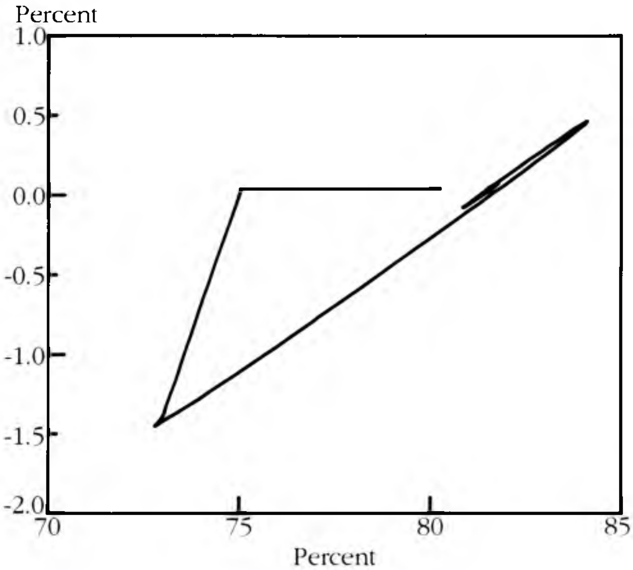
In the second simulation, inflation and capacity utilization rates are examined as the rate of M2 growth is reduced from 12 percent to 7 percent in the tenth year (see figures 6 and 7). At the beginning, the steady-state inflation and capacity utilization rates are 8.4 percent and 80.1 percent, respectively. Given the lag structures in equations (7) and (8), the initial effect is felt as a reduction in capacity utilization the following year, a decline that continues over the next three years. Inflation remains unaffected until the second year after the policy change, and then falls throughout each of the next six years. Even though the system is highly damped, there is still some overshooting in both the inflation and capacity utilization rates. Ultimately, the system reaches a new steady state with nearly the same capacity utilization rate (81.5 percent—not a statistically significant difference), but with an inflation rate of only 4.0 percent.

Figure 8 provides some insight into why the McElhattan-type "misspecified" model (at least in reference to the current one) yields reasonable results despite the apparent structural change over time. The figure plots the change in the inflation rate against the capacity utilization rate as the system returns to equilibrium following a reduction in the growth rate of M2. Clearly, a world with many such shocks could easily generate a plot similar to that of figure 3. A direct relationship between capacity utilization and changes in inflation would always be found, but depending on the latest shocks to the system, the actual estimated parameter

■ **13** Judd and Trehan's (1989) approach also finds relatively damped cycles, and is similar in spirit to this study in that it does not subscribe to any particular macroeconomic theory. The authors identify supply and demand shocks for a five-variable VAR system (unemployment rate, real GNP, nominal interest rate, labor supply, and foreign trade) using rather uncontroversial restrictions. Their approach yields even shorter cycles that exhibit much less overshooting. This could be the result of more detailed modeling (the inclusion of five variables) or of the use of quarterly rather than annual data. Because their study includes unemployment data, it is not limited by the capacity utilization problems discussed earlier.

FIGURE 8

Simulated McElhattan-Type Plot



SOURCE: Author's calculations.

values of the regression line would change. This is consistent with results presented earlier.

In short, as a rough approximation, this relatively simple model tracks the U.S. economy's response to the major economic events of the 1970s and 1980s reasonably well. It also provides some insight into why the basic relationship between capacity utilization and inflation appears to vary over time.

V. Conclusion

This paper examines the theoretical and empirical relationship between capacity utilization and inflation. Although there clearly is a connection between these two time series, earlier models suggest that structural changes occurred in the relationship over the 1953-89 period. Granger causality tests appear to confirm the suspicion that there is bidirectional causality between capacity utilization and a change in the inflation rate. One implication of this finding is that alternative models that treat both variables as endogenous should be employed.

A relatively simple two-equation structural model is developed here that is sufficient to explain the relationship. The dynamics of supply and demand relationships are employed to identify the system following Haynes and Stone (1985), with the model treating both capacity utilization and inflation as endogenous variables. No evidence of structural change is found from 1953 to 1989, but because of the relatively small number of yearly observations, only two subperiods are investigated.

The natural rate of capacity utilization is found to be about 81.5 percent and independent of the growth rate of M2. Although faster monetary growth increases both capacity utilization and inflation in the short run, only inflation is increased in the long run (moving in tandem with M2 growth). As a rough approximation, the model appears to track the real economy's reaction to supply and monetary shocks reasonably well. However, development of the proper framework for examining the endogenous relationship between capacity utilization and inflation is the most important contribution of this study.

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Financial Restructuring and Regional Economic Activity

by Brian A. Cromwell

Brian A. Cromwell is an economist at the Federal Reserve Bank of Cleveland. The author wishes to thank Randall Eberts, Joseph Haubrich, Louis Jacobson, Elizabeth Laderman, Katherine Samolyk, and Gary Whalen for helpful comments and suggestions.

Introduction

The relationship between the performance of the financial sector and economic activity has received increasing attention from economists in the past decade. The "credit view" holds that variation in the supply of financial services not captured in monetary aggregates can help to explain real economic activity. Empirical studies of the importance of the banking sector have been conducted at both the macro and the regional level and in general support the view that financial structure and stress can have real effects.

Interest in this view coincides with the dramatic restructuring of the financial sector that has accompanied both the advent of deregulation and interstate banking and the significant increase in bank failures in the 1980s. By the end of 1988, all but three states permitted some form of interstate acquisition of their banks, 14,600 offices of banking organizations existed outside of the organizations' home state, and more than half of these were permitted to offer all banking services.[1] Bank failures, which averaged 75 per year in the 1970s, rose from less than 50 per year in the early 1980s to more than 200 per year by 1987. Financial-sector restructuring in the form of bank mergers, takeovers, or failures can affect investment and consumption decisions by disrupting the links between borrowers and creditors.

This paper explores the impact of financial restructuring on economic activity using an alternative data set that in some respects more completely measures change in the local banking sector than do data used in previous research. Restructuring in the local banking sector is measured by the birth, expansion, contraction, and death of banks within a standard metropolitan statistical area (SMSA), as estimated by the Small Business Administration using Dun and Bradstreet files.

These data, known as the U.S. Establishment and Longitudinal Microdata (USELM), attempt to

■ 1 These figures come from a recent comprehensive review of interstate banking by King, Tschinkel, and Whitehead (1989). Earlier surveys include Whitehead (1983a, 1983b, and 1985), and Amel and Keane (1986).

record the location and employment levels of all establishments in all industries. For multi-establishment firms, ultimate ownership of each establishment is tracked. With respect to banking, an establishment equals a bank, a bank subsidiary, or branch office.[2] (Although the USELM establishment framework does not account for the variations in bank organizations across states, its advantage is that it is applied consistently.)

The USELM data are aptly suited for examining the disruption of credit relationships, since Dun and Bradstreet collects these data for the purpose of recording the creditworthiness of firms. A "death" is recorded if a firm fails or is taken over by management sufficiently different from existing management to warrant a reexamination of the firm's credit. To the extent that takeovers, management changes, and branch closings affect local credit relations, the employment effects from bank deaths can potentially measure the disruption of credit links between borrowers and lenders.

The empirical analysis presented here uses employment changes resulting from the birth, death, expansion, and contraction of small, midsized, and large banks as a proxy for restructuring in the banking industry. These measures are linked to the local economic performance of 217 SMSAs in the periods 1980-82 and 1984-86. If the credit view is supported, the impact of employment changes from a bank death should be negative and significantly greater in magnitude than the impact of a bank contraction, since a bank closing should be more disruptive to credit relations than simply a reduction in the bank's staff.

The results suggest that the deaths of midsized banks—those employing between 100 and 500 employees—have a negative but short-lived impact on economic activity. Exploration of the channels for this impact indicates that bank deaths affect employment in other midsized firms that presumably rely principally on local banking markets and that are the most likely customers of midsized banks. The results control for overall financial restructuring and lagged economic activity, and are robust across several specifications.

I. Local Economic Effects of Financial Stress

Restructuring due to financial stress—reflected in bank failures, closings, and mergers—is potentially detrimental to local economic growth.[3] In the case of a bank failure, several types of economic agents may be affected. Bank shareholders and uninsured depositors, for example, may suffer declines in wealth. For a local economy, however, any wealth effect is likely to be small. If the failed bank is merged with another bank, as is commonly the case, uninsured depositors may suffer no losses. Moreover, even if a failed bank is closed, these depositors generally recover (over time) a high percentage of their funds when the bank's assets are liquidated.

Another effect on local economic activity takes place through reductions in bank employment when banks fail or are taken over. In addition to this direct consequence for local employment, unemployed bank workers suffering a loss of personal income will also likely reduce their consumption expenditures, sending a further negative ripple (or multiplier) effect through the economy. The following section presents direct measures of the employment losses caused by bank deaths and looks for employment effects in nonbank sectors. More important, it examines a direct measure of employment losses in banking due to bank contractions to see if the spillovers from bank employment losses are different for failures than for contractions.

Credit Disruptions

In addition to these two potential effects, bank closings can disrupt credit relationships. The credit-view literature holds that the principal channel of a bank failure's economic impact is the disruption of borrower-lender relationships. Each lender is assumed to have more information on its existing borrowers than do other potential lenders.[4] When bank failure results in closure rather than reorganization, borrowers are forced to seek credit from new sources. During the period in which a new long-term credit

■ 2 In practice, Dun and Bradstreet tracks all banking establishments listed in telephone directories, including branch offices.

■ 3 This section in part follows Gilbert and Kochin's (1990) presentation.

■ 4 Gertler (1988) surveys the literature on credit and aggregate economic activity. Articles on the theory of financial intermediation include Diamond (1984) and Campbell and Kracaw (1980).

relationship is established, borrowers are likely to face higher costs of credit or credit rationing.

Even if the failed bank is merged into a surviving bank, borrowers may encounter new loan policies and new senior management if they apply for extensions of their credit. Again, the terms of credit are likely to be less favorable than those offered by the previous management.

The USELM data count bank takeovers and mergers as bank deaths *only* if there is a change in operating management sufficient for Dun and Bradstreet to reexamine the new organization's credit rating. To the extent that takeovers, management changes, and branch closings affect local credit relations, USELM's measures of change in bank employment due to bank deaths can be used to estimate the disruption of credit links between borrowers and lenders. Because we can also measure employment changes due to bank contractions, we use the difference between the impact of bank contractions and bank deaths to distinguish between the direct employment effects of a bank death and the impact of credit disruptions.[5]

Empirical Evidence

Empirical studies of the importance of the banking sector at both the macro and the regional level generally support the view that financial structure and stress can affect economic activity.

In a study of the macro effects of financial stress, Bernanke (1983) argues that extensive bank runs and defaults in the 1930-33 financial crisis reduced the efficiency of the financial sector in performing its intermediation function, and that this had adverse effects on real output through other than monetary channels . He examines the effect of the real value of the change in deposits of failed banks on the growth rate of industrial production. Using regression analysis with monthly data for the years 1919 through 1941, Bernanke finds that bank failures have a negative and statistically significant effect on industrial output. Samolyk (1988) conducts a similar test on British data, using corporate and noncorporate insolvencies as proxies for the health of the financial sector, and also finds that credit factors matter empirically in explaining output. Using Canadian data, Haubrich (1990) also determines that the credit disruptions resulting from bank failures, as opposed to expansions or contractions, affect economic activity.

The credit view would also predict an impact of stress in local banking markets on local economies. Calomiris, Hubbard, and Stock (1986) examine the impact of bank failures on real farm output. Using annual state data for farm output, they find that the number of bank failures lagged one year has a negative and statistically significant effect. Gilbert and Kochin (1990) test the hypothesis that bank failures have adverse effects on sales subject to sales tax and on county employment using rural county-level data. They find that bank closings have a negative impact on local sales and on nonagricultural employment.

II. Alternative Data on Financial Restructuring: The USELM File

Financial restructuring at the local level is analyzed here by the birth, expansion, contraction, and death of banking establishments at the SMSA level, as measured in the USELM data for the periods 1980-82 and 1984-86. The longitudinal establishment data files of the U.S. Establishment and Enterprise Microdata (USEEM) were constructed primarily from data in the Dun and Bradstreet Duns Market Identifier Files (DMI). The Small Business Administration then assembled more than 16 million establishment records contained in the DMI files to construct the USELM file. A team from the Brookings Institution merged DMI data longitudinally and associated each establishment with its owners. The longitudinal detail in the data makes it possible to measure employment change of establishments in a given size class. It also allows employment change to be decomposed by establishment birth or death, or by the growth (or contraction) of continuing establishments, and allows for the tracking of mergers, acquisitions, and divestitures.

The data base includes employment figures and industry classifications for all establishments and enterprises, sales data for all enterprises and subsidiaries, age of all nonbranch establishments, and organizational status and geographic data for each establishment. In principle, every establishment in the United States is covered, except for federal agencies.

Dun and Bradstreet's principal business is provision of credit ratings, which must be

■ 5 We maintain the assumption that the local economic effect of a bank contraction results solely from the multiplier effect of reduced employment, rather than from credit disruptions.

updated to reflect discontinuances of business management. Among other activities, the firm tracks court proceedings in bankruptcy cases in order to record the creditworthiness of establishments. A death is recorded if the establishment is closed or is taken over by management that is sufficiently different to warrant reexamining the firm's credit.[6]

In the case of banking, a death could represent the takeover of a bank by another bank (recorded simultaneously as a death and an expansion), a major change in ownership and management (recorded simultaneously as a death and a birth), or a true failure (recorded solely as a death). As such, it is an imperfect measure of liquidations of financial institutions. To the extent that takeovers and management changes affect local credit relations, the USELM measure of bank deaths can proxy for the disruption of credit links between borrowers and lenders.

Advantages of the USELM Data

In an exploration of the impact of bank structure on regional development, Bauer and Cromwell (1989) show that the private banking sector appears to be systematically related to firm births. This study, however, is limited by the inadequacy of data on the location of financial institutions. Bank data were obtained from the Federal Financial Institutions Examination Council's Reports on Condition and Income, known as call reports, for 1980. For some financial measures such as total loans, however, it is not possible to determine where the loans were made, even though their dollar value is known. For example, loans made by an Ohio bank to firms in Florida and Ohio are counted in the same way. An additional measurement problem is that a call report for a consolidated banking unit may include data for branches not located in the SMSA. In states that allow branch banking, activity at the branches may be reported solely in the headquarter's SMSA.

In principle, the USELM data report the location and employment levels of banking establishments with a greater degree of accuracy than the call report data, which record bank statistics at the firm (headquarters) level for many establishments (subsidiaries or branches). For example, a bank headquartered in Cincinnati could report data for its Columbus and Dayton branches in its call report, resulting in a distortion of the measured banking activity in Cincinnati. For purposes of the USELM data, however, branches in Columbus and Dayton are recorded as establishments in those SMSAs. Out-of-state ownership of establishments is also recorded, allowing examination of the impact of interstate banking.

Disadvantages of the USELM Data

The USELM data set, while having advantages for this analysis, is not problem free. The major reasons to question the validity of statistics derived from DMI files stem from three characteristics of the underlying data-collection effort. First, the employment figures are self-reported by establishments, usually in telephone interviews. Second, employment figures are not routinely updated; updates are primarily a result of requests for credit checks. Jacobson (1985) reports that substantial lags can occur between the date the file is extracted and the last time the firm was surveyed. Employment statistics are often more than two years out of date, which may lead to infrequent reports for smaller and slower-growing establishments with less need for credit checks, and to delays in picking up shutdowns and status changes for these firms. Third, there are delays in recognizing the creation of new establishments. A business may be in existence three to five years before it is recorded as a birth in the USELM data.

To deal with some of these shortcomings, the Small Business Administration has modified the Dun and Bradstreet data in several ways in constructing the USELM file.[7] Establishments for which employment data were missing were assigned the state-level median employment for organizations in their standard industrial classification (SIC) code. Some 8.7 percent of establishments in finance, insurance, and real estate (FIRE) received estimated employment figures in

■ 6 In practice, a death is identified when a firm appearing in the Dun and Bradstreet file in an initial year does not appear in an end-year file. The reason could be either that the firm was actually closed or that its identifying number was changed as a result of new management. Similarly, a birth is identified by the appearance of a firm in the end-year file. The Small Business Administration uses two-year intervals to track firms.

■ 7 See Harris (1983) and Armington and Odle (1983, 1984).

TABLE 1

Bank Employment Share: Small, Midsized, and Large Banks

Bank Size	1980	1982	1984	1986
Type 1: 0 to 100 employees	0.247 (0.277)	0.247 (0.269)	0.256 (0.260)	0.229 (0.224)
Type 2: 100 to 500 employees	0.237 (0.282)	0.232 (0.271)	0.188 (0.206)	0.190 (0.196)
Type 3: More than 500 employees, in-state HQ	0.480 (0.351)	0.486 (0.344)	0.511 (0.308)	0.519 (0.294)
Type 4: More than 500 employees, out-of-state HQ	0.035 (0.127)	0.035 (0.120)	0.046 (0.119)	0.062 (0.145)

NOTE: Numbers are expressed as SMSA means. Standard deviations are in parentheses.
SOURCE: Author's calculations based on USELM data.

1980 (Harris [1983]). Furthermore, when total firm-level employment exceeded the aggregate establishment-level employment data, it was assumed that the number of branches had been underreported. Branches were then imputed from this excess employment according to the average branch size for the particular industry-size classification. In 1980, 17 percent of the establishments in FIRE were imputed. The imputed branches lack geographic information beyond the state, however.

Finally, records that were not updated during the sample period were removed from the data set. The remaining updated records were then weighted to reflect the underlying population. Weights were assigned on the basis of the level of reporting in firm categories based on size, organization type, and industry. Reporting problems within the banking sector led analysts at the Brookings Institution to separate commercial banking from the rest of FIRE, resulting in a much more accurately weighted population in both parts of this sector. The level of reporting problems is higher outside SMSAs than within them (Armington and Odle [1983]), which does not present a problem for the analysis where the SMSA is the unit of observation.

In general, geographic errors from nonreporting of branches appear to be more severe for large firms with several establishments. Errors resulting from records that are not routinely updated and from inaccurate geographic distribution of weighting are more severe for small firms that have infrequent credit checks under the Dun and Bradstreet system. The employment statistics and location of independent midsized firms with few establishments or branches however, appear to be more accurately measured. The midsized firms undergo more frequent credit checks than small businesses and are less likely to be widely distributed geographically. The following estimation presents plausible and significant results for the economic impact of midsized banks on midsized firms, but weaker results for small and large banks. Whether these results are due to the true importance of midsized banks or to measurement error is uncertain, so the findings should be interpreted with caution.

III. Model Specification

This paper assumes that banking employment losses due to bank deaths, as measured in the USELM data, are a reasonable proxy of credit disruptions during restructuring in local banking markets. To control for the direct effects of the job losses of bank employees and the credit effects of restructuring, I use the difference between the impact of bank contractions and bank deaths on local economic activity to identify real economic effects resulting from disruption of credit channels. Data were collected for 217 SMSAs for the periods 1980-82 and 1984-86.[8] The sample was limited to those SMSAs for which complete information was available.

Four types of establishments are identified through an extract of the USELM data base. Type 1 establishments belong to independent firms with fewer than 100 employees—typically single-establishment small businesses. Type 2 establishments belong to independent firms with 100 to 500 employees—midsized firms that may have more than one establishment. Type 3 establishments belong to firms with greater than 500 employees headquartered within the same state. Type 4 establishments belong to firms with greater than 500 employees headquartered out of state.

The percentage of bank employees in the four types of firm categories used in this study (averaged across SMSAs) is given in table 1. The

8 The Small Business Administration's time-series construction of the files compels the use of two-year periods.

TABLE 2

Distribution of Bank Employment by Asset Category, 1980

Asset Class ($ millions)	Number of Banks	Number of Employees	Employees per Bank
0 to 5	874	3,113	3.6
5 to 10	1,937	16,282	8.4
10 to 25	4,662	80,604	17.3
25 to 50	3,553	125,868	35.4
50 to 100	1,972	144,710	73.4
100 to 300	1,156	220,047	190.4
300 to 500	198	79,216	400.1
500 to 1,000	158	112,331	711.0
1,000 to 5,000	157	280,302	1,785.4
5,000 and more	37	419,908	11,348.9

SOURCE: Federal Deposit Insurance Corporation, *Annual Report*, 1980.

TABLE 3

Bank Employment by Asset Category over Time

Asset Class ($ millions)	Employees per Bank 1980	Employees per Bank 1986	Percent Change
0 to 25	13.4	11.3	13.2
25 to 100	49.0	33.3	32.1
100 to 1,000	272.2	150.1	44.9
1,000 and more	3,609.3	2,710.3	24.9

SOURCES: Federal Deposit Insurance Corporation, *Annual Report*, 1980, and *Statistics on Banking*, 1986.

impact of financial restructuring is suggested by changes in these employment categories over time. Employment in small banks declined from an average 24.7 percent in 1980 to 22.9 percent in 1986; employment in midsized banks declined from 23.7 percent to 19.0 percent. Employment in large in-state banks, however, increased from 48.0 percent to 51.9 percent, while employment in out-of-state banks almost doubled, from 3.5 percent to 6.2 percent, reflecting the growing importance of interstate banking.

Bank Employment as a Proxy for Financial Structure

Before concluding that changes in bank employment represent changes in financial structure, one must assess the validity of bank employment as a proxy for bank size, the impact of labor productivity in the banking sector, and changes in the size distribution of banks.

Table 2 reports the distribution of bank employment in 1980 across banks categorized by asset size. In general, the standard definition of a midsized firm as having 100 to 500 employees matches up well with the standard definition of a midsized bank as having assets between $100 million and $1 billion. Average employment per bank ranges from 73 employees for banks in the $50 million to $100 million category to 711 employees for banks in the $500 million to $1 billion category. Similarly, our measures of small and large firms match up with standard definitions of small and large banks.

The data suggest large improvements in labor productivity (measured as employees per bank, controlling for assets) for all bank size categories over the 1980-86 period. As shown in table 3, average employment per bank declined over the period for banks in all asset categories. Decreases ranged from 13 and 32 percent for small banks in the $0 to $25 million and $25 million to $100 million categories, respectively, to 45 percent for midsized banks in the $100 million to $1 billion category. Productivity gains for large banks with assets of greater than $1 billion averaged only 25 percent. Again, our definition of a midsized firm is consistent with the definition of midsized banks, which averaged 150 employees per bank in 1986.

Table 4 reports shifts in the relative importance of small, midsized, and large banks over the 1980-86 period. In general, small banks declined in both number and relative importance, midsized banks were little changed, and large banks increased in importance.

Small banks with less than $25 million in assets constituted 50.8 percent of all banks in 1980 but held only 5.1 percent of all assets. By 1986, they had declined to 34 percent of all banks and their share of assets stood at 2.4 percent. Declines in the share of assets also occurred in banks in the $25 million to $50 million and $50 million to $100 million categories. The number and share of assets of midsized banks were little changed over the 1980-86 period. As a percentage of all banks, those in the $100 million to $300 million category increased from 7.9 percent to 13.4

TABLE 4

Distribution of Banks by Asset Category

Asset Class ($ millions)	Number of Banks	Percent of Banks	Percent of Assets
1980:			
0 to 25	7,473	50.8	5.1
25 to 50	3,553	24.2	6.8
50 to 100	1,972	13.4	7.3
100 to 300	1,156	7.9	9.8
300 to 500	198	1.3	4.1
500 to 1,000	158	1.1	5.8
1,000 and more	194	1.3	61.1
Total	14,704	100.0	100.0
1986:			
0 to 25	4,823	34.0	2.4
25 to 50	3,685	26.0	4.5
50 to 100	2,899	20.4	6.8
100 to 300	1,903	13.4	10.4
300 to 500	334	2.4	4.3
500 to 1,000	216	1.5	5.1
1,000 and more	340	2.4	66.5
Total	14,200	100.0	100.0

SOURCES: Federal Deposit Insurance Corporation, *Annual Report*, 1980, and *Statistics on Banking*, 1986.

TABLE 5

Nonbank Employment Rates: Small, Midsized, and Large Firms

Firm Size	1980	1982	1984	1986
Total employment	0.385 (0.064)	0.385 (0.066)	0.390 (0.065)	0.411 (0.072)
Type 1: 0 to 100 employees	0.133 (0.027)	0.138 (0.030)	0.130 (0.026)	0.136 (0.025)
Type 2: 100 to 500 employees	0.056 (0.017)	0.053 (0.016)	0.053 (0.015)	0.053 (0.014)
Type 3: More than 500 employees, in-state HQ	0.080 (0.045)	0.077 (0.044)	0.077 (0.044)	0.080 (0.048)
Type 4: More than 500 employees, out-of-state HQ	0.119 (0.050)	0.112 (0.046)	0.110 (0.046)	0.114 (0.046)

NOTE: Numbers are expressed as SMSA means. Standard deviations are in parentheses.
SOURCE: Author's calculations based on USELM data.

percent, but their share of assets rose only from 9.8 percent to 10.4 percent. Banks in the $300 million to $500 million category increased their share of assets slightly, while those in the $500 million to $1 billion range showed a small decline in asset share. Banks with assets of greater than $1 billion constituted only 1.3 percent of all banks in 1980 but accounted for 61.1 percent of assets. By 1986, the asset share of large banks rose to 66.5 percent.

In general, the changes in bank structure suggested by the employment shifts in table 1 reflect transformations observed in the distribution of banks by asset size. Small banks declined in importance, while large banks gained. The drop of employment in the midsized banks, however, is most likely the result of strong labor productivity gains, which exceeded those of both the small and large banks, rather than a decline in their importance in terms of assets. Bank employment at any particular time, however, does appear to track closely with asset size. Thus, the use of bank employment losses due to bank deaths appears to be a reasonable proxy of credit disruptions. In addition, the definitions of small, midsized, and large firms used here correspond with standard definitions of small, midsized, and large banks.

Dependent Variable and Specification

County-level employment data from the Bureau of Labor Statistics aggregated to the SMSA level are used to measure local economic activity. Bank employment, as reported in the USELM data, is subtracted from the aggregate employment. An alternative proxy for output (personal income) yielded qualitatively similar results to those reported here, but in order to compare total employment rates with the employment rates in firms of various size classes, it is not used. Thus, specifications are also estimated with employment in small, midsized, and large firms (nonbank) as dependent variables. These average employment rates (employment divided by population) are reported in table 5. Total employment rose over the period, which began in recession, from 38.5 percent in 1980 to 41.1 percent in 1986. Employment in small and midsized firms was essentially flat. Employment in large firms headquartered within the state changed little over the period, as did employment in out-of-state firms.

The effects of bank deaths on local economic activity are estimated using regression

TABLE 6

Summary Statistics: Independent Variables

WAGE	9.256 (1.875)	*CONTRACT* 1	0.011 (0.035)
TAX	0.404 (0.038)	*CONTRACT* 2	0.014 (0.047)
EXPAND 1	0.033 (0.073)	*CONTRACT* 3	0.034 (0.075)
EXPAND 2	0.020 (0.055)	*CONTRACT* 4	0.005 (0.034)
EXPAND 3	0.057 (0.151)	*DEATH* 1	0.020 (0.056)
EXPAND 4	0.009 (0.101)	*DEATH* 2	0.011 (0.046)
BIRTH 1	0.015 (0.032)	*DEATH* 3	0.040 (0.110)
BIRTH 2	0.025 (0.068)	*DEATH* 4	0.001 (0.004)
BIRTH 3	0.099 (0.302)		
BIRTH 4	0.024 (0.173)		

NOTE: Numbers are expressed as SMSA means. Standard deviations are in parentheses.
SOURCE: Author's calculations based on USELM data.

analysis. The dependent variable is the employment rate at the end of the period. (Using employment levels and including population as an independent variable yielded similar results, as did using changes in employment and changes in employment rates.)

The following measures potentially affecting employment growth are included as independent variables:

1) *LWAGE* = average (log) wage of production workers in the SMSA as measured by the Census of Manufacturers.

2) *TAX* = effective corporate tax rate for the state.

3) *DUM* 86 = a dummy variable equaling 1 for the 1984-86 period.

4) *BIRTH* 1-4 = percent change in banking employment due to the births of banks of types 1 through 4.

5) *EXPAND* 1-4 = percent change in banking employment due to the expansion of banks of types 1 through 4.

6) *CONTRACT* 1-4 = percent change in banking employment due to the contraction of banks of types 1 through 4.

7) *DEATH* 1-4 = percent change in banking employment due to the deaths of banks of types 1 through 4.

8) lagged employment rates.

The means and standard deviations for these variables are given in table 6. Financial restructuring is measured by the percent change in bank employment due to births, expansions, contractions, and deaths (*BIRTHi*, *EXPANDi*, *CONTRACTi*, and *DEATHi*; $i = 1, ...4$) for banks of types 1 through 4. Note that the credit-disruption hypothesis cannot explain why an expansion or birth of a bank would have an effect on nonbank employment. All components of change in banking are included for completeness, however. In particular, the expansion and contraction variables appear separately in case the multiplier effect of changes in bank employment on local economies is not symmetric. Differences in the estimated coefficients of *CONTRACT* and *DEATH* variables are meant to measure the impact of credit disruptions. On average, 2.0, 1.1, 4.0, and 0.1 percent of SMSA bank employment is lost over a two-year period, due to deaths of bank types 1 through 4, respectively.

As suggested in the literature on firm location, wages and tax rates are included to control for their impact on economic growth. A dummy variable for the 1984-86 period controls for any fixed effect associated with this period of economic expansion.

Following previous empirical studies, the specification includes lagged values of the dependent variables as independent variables to control for the possibility of a spurious relationship between bank deaths and employment. Suppose the causality between bank deaths and employment actually runs from employment to bank deaths. Banks tend to close after periods of relatively slow regional economic growth. If the lagged values of the dependent variables were not included as independent variables, the coefficients on the bank death variables would tend to be negative and significant even if bank deaths had no true effect on employment. The regression results thus indicate whether lagged bank deaths explain employment after accounting for employment in the past year.

Timing problems in the data make it impossible to entirely discount a spurious correlation.

The dependent variables are 1) overall employment data for the SMSA—an average rate for the ending year—and 2) employment rates for various-sized firms reported in the USELM data, based on end-of-calendar-year employment. The bank death data are employment changes due to deaths that occur in the two-year period from December of the beginning year to December of the ending year. The lagged dependent variables are employment rates at the beginning and middle of the two-year period. Bank deaths at the end of the two-year period (part of our independent variable) are thus potentially caused by adverse economic conditions at the end of the period (our dependent variable).

To explore the simultaneity problem further, however, we examine the effect of bank deaths on economic activity in the year following the two-year period. We also examine the impact of economic conditions at the beginning of the period on bank deaths. Finally, we test whether our results are driven by adverse shocks occurring in the oil states. The results, while not conclusive, suggest that our measure of the impact of banks deaths on regional activity is not being driven by simultaneous-equation bias.

IV. Estimation Results

The model was estimated on the pooled sample of 434 observations using ordinary least squares, which are presented in table 7.

Model 1 (in column 1), which uses the total employment rate as the dependent variable, suggests that high wages have a negative impact on total employment, while taxes have no significant effect. Lagged employment rates have a significant effect, as does the 1984-86 dummy variable.

The expansion and births of large in-state banks (*EXPAND* 3 and *BIRTH* 3) have a small but statistically significant effect on nonbank employment. The coefficient on *EXPAND* 3 is 0.012 and is significantly different from zero at the 95 percent confidence level, which indicates that a 10 percent increase in bank employment from the expansion of large banks raises the nonbank employment rate by 0.12 percentage point. A 10 percent increase from the birth of large banks raises the employment rate by 0.07 percentage point.

The contraction of bank employment (*CONTRACT* 1 through *CONTRACT* 4) does not have a statistically significant effect on nonbank employment for any of the bank types. The death of midsized banks (*DEATH* 2), however, has a statistically significant negative effect on the nonbank employment rate. The estimated coefficient is -0.053 with a standard error of 0.020, suggesting that a 10 percent decrease in bank employment from the death of midsized banks reduces the nonbank employment rate by 0.53 percentage point. Given that the average employment rate in the sample is 39.8 percent, this represents a drop in nonbank employment of 1.3 percent. The estimated coefficient for midsized bank contraction (*CONTRACT* 2) is -0.004 and is statistically not significantly different from zero. The difference between the estimated coefficients of *DEATH* 2 and *CONTRACT* 2 suggests that the effect of midsized bank deaths on employment is almost entirely due to credit disruptions as opposed to the multiplier effect of lost bank jobs. An F-test rejects the hypothesis that the coefficients are identical at the 90 percent confidence level.

The coefficients for *DEATH* 3 and *DEATH* 4 are negative, and the coefficient for *DEATH* 1 is positive, but all are statistically insignificant. These insignificant effects for deaths of types 1, 2, and 3 banks could result from the relative importance of midsized banks to local economies, from a difference in the type of deaths they represent (for example, small-bank deaths being takeovers rather than true failures), or from the relative accuracy of the midsized bank data previously discussed. Results should thus be taken as positive evidence for the importance of midsized banks, rather than as evidence for the lack of importance of small and large banks.

To investigate further the local economic impact of bank deaths, the model was reestimated with the small-, midsized-, and large-firm employment rates as dependent variables in models 2, 3, and 4, respectively (shown in columns 2, 3, and 4 of table 7). In model 2, the expansion of midsized banks has a positive effect on small-business employment, while the contraction of small banks has a negative effect. Bank deaths do not have a significant effect on small businesses, which is contrary to the common view that small firms are the first to suffer the effects of a credit crunch. It is possible, however, that these firms, many of which are relatively new or small "mom-and-pop" operations, rely more on informal sources of capital—such as loans from friends and relatives, retained earnings, and personal savings—than on funds from commercial banks. This would make small firms less likely to be affected by bank deaths. It

TABLE 7

Regression Results: Impact of Financial Restructuring on Employment Rates

Coefficient	(1) Total Employment Rate	(2) Small-Firm Employment Rate	(3) Midsized-Firm Employment Rate	(4) Large-Firm Employment Rate
LWAGE	-0.013[a]	0.001	-0.001	0.005
	(0.005)	(0.001)	(0.001)	(0.004)
TAX	0.029	0.003	0.0002	0.038[a]
	(0.024)	(0.007)	(0.0060)	(0.018)
EXPAND 1	0.009	0.0002	-0.005	0.001
	(0.013)	(0.0038)	(0.003)	(0.010)
EXPAND 2	0.002	0.010[a]	-0.003	0.011
	(0.017)	(0.005)	(0.004)	(0.013)
EXPAND 3	0.012[a]	0.002	-0.0002	0.004
	(0.006)	(0.002)	(0.0015)	(0.005)
EXPAND 4	-0.009	0.002	0.006[a]	-0.007
	(0.009)	(0.003)	(0.002)	(0.007)
BIRTH 1	-0.032	-0.006	-0.002	-0.009
	(0.031)	(0.009)	(0.008)	(0.023)
BIRTH 2	-0.001	0.003	-0.001	-0.004
	(0.014)	(0.004)	(0.003)	(0.010)
BIRTH 3	0.007[a]	0.001	0.0005	0.004[b]
	(0.003)	(0.001)	(0.0008)	(0.002)
BIRTH 4	-0.002	-0.002	0.001	0.003
	(0.005)	(0.002)	(0.001)	(0.004)
CONTRACT 1	-0.030	-0.013[b]	-0.009	-0.016
	(0.027)	(0.008)	(0.007)	(0.020)
CONTRACT 2	-0.004	-0.002	-0.004	-0.025[b]
	(0.019)	(0.006)	(0.005)	(0.014)
CONTRACT 3	0.007	0.002	-0.004	0.003
	(0.012)	(0.004)	(0.003)	(0.009)
CONTRACT 4	0.017	-0.001	-0.0003	0.011
	(0.028)	(0.008)	(0.0070)	(0.021)
DEATH 1	0.022	0.002	-0.001	0.005
	(0.017)	(0.005)	(0.004)	(0.013)
DEATH 2	-0.053[a]	-0.006	-0.017[a]	0.001
	(0.020)	(0.006)	(0.005)	(0.015)

is also possible that these results are driven by the errors in the data for small firms discussed above.

The impact of financial restructuring on midsized firms is measured in model 3. Midsized firms are more likely to rely on local commercial banks than on national credit markets (such as with large firms) or on informal sources of capital (such as with small firms); see Elliehausen and Woken (1990). They are thus more likely to be affected by stress in the financial sector. The results support this conclusion. The estimated coefficient for *DEATH* 2 is -0.017 with a t-statistic of 3.40, indicating that a 10 percent decrease in bank employment from bank deaths reduces the employment rate of midsized firms by 0.17 percentage point. With an average employment rate of 5.3 percent, this represents a 3.2 percent drop in midsized firms' employment. Again, an F-test rejects at the 90 percent confidence level the hypothesis that the coefficient for *DEATH* 2 equals the coefficient for *CONTRACT* 2.

TABLE 7 continued

Regression Results: Impact of Financial Restructuring on Employment Rates

Coefficient	(1) Total Employment Rate	(2) Small-Firm Employment Rate	(3) Midsized-Firm Employment Rate	(4) Large-Firm Employment Rate
DEATH 3	-0.002	0.003	-0.003	-0.005
	(0.009)	(0.002)	(0.002)	(0.006)
DEATH 4	-0.008	0.020	-0.062	0.026
	(0.254)	(0.074)	(0.063)	(0.189)
CONSTANT	0.036[a]	-0.002	0.002	-0.040[a]
	(0.016)	(0.005)	(0.004)	(0.012)
DUM 86	0.010[a]	-0.002[a]	0.002[a]	0.007[a]
	(0.002)	(0.001)	(0.001)	(0.002)
EMPRTB	1.624[a]	0.308[a]	0.047[a]	0.001
	(0.088)	(0.026)	(0.022)	(0.066)
EMPRTA	-0.697[a]	-0.291[a]	-0.036	0.031
	(0.092)	(0.027)	(0.023)	(0.069)
SEMPRTA	—	0.975[a]	—	—
	—	(0.013)	—	—
MEMPRTA	—	—	0.885[a]	—
	—	—	(0.019)	—
GEMPRTA	—	—	—	0.970[a]
	—	—	—	(0.017)
Log likelihood	1122.7	1661.8	1726.3	1250.9
R^2	0.933	0.965	0.912	0.912
Mean of the dependent variable	0.398	0.137	0.053	0.079
Number of observations	434	434	434	434

a. Significant at the 95 percent confidence level.
b. Significant at the 90 percent confidence level.
NOTE: Numbers are expressed as estimated coefficients. Standard errors are in parentheses.
SOURCE: Author's calculations.

The impact of midsized bank deaths on midsized firms is small but statistically significant. This result is plausible because these firms are most likely to rely on local banks and because midsized commercial banks are more likely to concentrate on lending to such firms. This finding controls for lagged total employment rates and for the lagged midsized-firm employment rate. A spurious correlation is still possible, however, if these lagged measures are inadequate controls due to timing problems in the data.

Model 4 measures the impact of financial restructuring on large firms. These firms are more likely to have access to national credit markets and thus may be less affected by local restructuring. The results suggest that the expansion or birth of large banks has a positive impact on large-firm employment, while the contraction of midsized banks has a negative impact. These findings are significant at the 90 percent confidence level, but not at the 95 percent level, implying that changes in local bank structure do not have powerful effects on large firms.

To explore the robustness of the results and the potential for simultaneous-equations bias, other specifications of the model were also tested. First, the employment rate in the year following the two-year period used in the USELM file was used as a dependent variable, with lagged employment rates for the beginning, middle, and end of the period included as independent variables. The estimated coefficient on *DEATH* 2 was -0.013 (compared with an estimated coefficent of -0.053 in model 1) with a t-statistic of 1.024. This smaller (and statistically insignificant) effect suggests that either the effect of bank deaths dampens out quickly over time, or that the original result was driven by simultaneity bias. This specification, however, allows bank deaths to have an effect up to three years after they occur. Such long-term influences of credit disruptions are unlikely if banking markets are competitive.

Second, the impact of local economic conditions on bank deaths was explored by using midsized bank deaths as a dependent variable and the beginning-of-period economic conditions as an independent variable. The economic conditions (total employment rates and employment rate of midsized firms) had no explanatory effect on *DEATH* 2. (T-statistics of the economic variables in various specifications were never larger than 0.50.) This suggests that bank deaths were not statistically driven by our measures of local economic conditions.

Finally, robustness of the results was tested by including geographic dummy variables to control for effects from economic distress in oil-producing states, which had an especially large number of bank failures. In particular, we tested whether the findings were solely due to bank failures in Oklahoma, Louisiana, and Texas between 1984 and 1986. Controlling for these states did not affect the results. The coefficient on *DEATH* 2 in model 1 remained above -0.040 in the various specifications tested, and the t-statistic did not fall below 2.60.

In sum, the structure of the data in this experiment does not permit standard tests of Granger causality or simultaneity bias. The alternative specifications tested, however, suggest that the results are not being driven by the occurrence of bank deaths in economically distressed states or by any obvious feedback of economic conditions on bank deaths.

V. Conclusion

Restructuring of the financial sector in the form of bank mergers, failures, or takeovers potentially affects investment and consumption decisions by disrupting the links between borrowers and creditors. Empirical evidence at both the macro and regional levels has shown that financial structure and stress can have real economic effects.

This paper further explores the impact of financial restructuring on local economies using a data set that measures change in the local banking sector by the birth, expansion, contraction, and deaths of banks at the metropolitan level.

The empirical analysis suggests that, controlling for overall financial restructuring and lagged economic activity, the deaths of midsized banks—employing between 100 and 500 employees—have a negative but short-lived impact on local economic activity. Furthermore, employment in other midsized firms appears to be most directly affected by these deaths. These firms presumably rely on local banking markets and are the most likely customers of midsized banks. The results are robust across several specifications.

The strongest effects of bank deaths are found for midsized banks and firms, but this should not be interpreted to mean that the deaths of small and large banks have no effect. Measurement problems, which exist for all size categories, are least severe for midsized banks and firms, which can account for the statistical significance of these results and the statistical insignificance of the results for the other two types. Nonetheless, the results suggest that midsized banks are an important source of funds for midsized firms and that a disruption of this link through financial restructuring can have negative short-run local economic effects.

The effects of credit disruption appear to be short run. In particular, the impact of bank deaths on employment rates appears to die out after two years. One would expect such a result if banking markets were competitive and contestable. In such markets, firms that found their source of credit disrupted by a bank death would quickly be able to establish credit relations with another institution. Thus, the results presented here, while consistent with the credit-view theory that disruption in financial markets can have real effects, also suggest that these effects are short-lived, as one would expect in a competitive environment.

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The Short-Run Dynamics of Long-Run Inflation Policy

by John B. Carlson,
William T. Gavin,
and Katherine A. Samolyk

John B. Carlson and Katherine A. Samolyk are economists and William T. Gavin is an assistant vice-president and economist at the Federal Reserve Bank of Cleveland. The authors gratefully acknowledge helpful suggestions from David Altig, Richard Baillie, Dennis Hoffman, and Richard Jefferis.

Introduction

Currently, some economists, legislators, and policymakers are recommending that the Federal Reserve adopt price stability or an explicit price-index target as its primary long-term monetary policy objective.[1] This recommendation is based on three ideas. First, high and uncertain inflation leads to an inefficient allocation of resources. Second, inflation is ultimately a monetary phenomenon, so controlling inflation over the long term would be the sensible goal for monetary policy. Finally, a positive inflation trend provides no net long-run benefit to the economy.

The dynamic effects of monetary policy are difficult to understand for many reasons, the most important of which is simply that the term "monetary policy" can be interpreted in different ways. It may be defined in terms of the short-run interest-rate policy of the Federal Reserve, the growth rates of the monetary aggregates, or the ultimate policy objectives mandated by Congress through the Humphrey-Hawkins Act.

Our purpose in this paper is to ascertain the short- and long-term implications of an inflation policy for real output. Although a complete cost-benefit analysis is not provided, we do present one basis for assessing the short-run costs of lowering trend inflation, based on the estimated dynamic relationship between output and inflation. The simple framework developed here abstracts from issues concerning the implementation of monetary policy. We do not specify a policy reaction function, nor do we include interest rates or the money supply, variables normally associated with monetary policy. Rather, an inflation policy is defined in terms of a disturbance that exclusively determines trend inflation.

In order to isolate this permanent inflation disturbance, we adopt an identification method developed by Blanchard and Quah (1989). Using a model of output growth and unemployment, they identify two independent disturbances that they interpret as shocks to aggregate demand

■ 1 See House Joint Resolution 409, introduced by U.S. Congressional Representative Stephen L. Neal of North Carolina, and testimony in support of the resolution by Hoskins (1990).

and aggregate supply. Their identifying assumption is that only the supply shock has permanent effects on output, while neither supply nor demand disturbances affect the trend rate of unemployment. Using a model of inflation and output, we identify two independent disturbances that we interpret as innovations to inflation policy and real output. Our identifying assumption is that only the inflation shock affects trend inflation. Innovations to real output may affect the path of inflation in the short run, but the inflation shock alone determines the inflation trend.

Sims (1986) discusses how estimated disturbances can be interpreted as reflecting government policy choices in the context of a VAR model. Policy actions are associated with prediction errors in the corresponding policy variables. We interpret the shocks that drive the inflation trend as reflecting innovations accommodated by policy. We then compute the impulse-response functions and variance decompositions for real output in order to determine how these inflation-policy shocks contribute to real GNP fluctuations.

To understand our purposes, consider an economy characterized by the classical dichotomy. In such an economy, the processes driving inflation and output could be identified by restricting inflation so that it would have no effect on real output. The estimated nominal disturbances would then unambiguously reflect monetary policy actions, albeit irrelevant ones for real economic activity.

Similarly, this study attempts to disentangle the disturbances associated with inflation policy from the real disturbances driving the macroeconomy. It does not, however, require adherence to the classical dichotomy in either the short or the long run. Our specification allows both estimated shocks to influence output and inflation in the short run.

To impose our assumption that only the inflation shock determines the inflation trend, we constrain the model so that the output disturbance has only a transitory effect on inflation. Inflation shocks are interpreted as policy innovations that can affect both the short- and long-run dynamics of the system. Thus, our key identifying assumption is consistent with a large variety of economic structures, including all of the major macroeconomic theories.

In examining the dynamic consequences of these two fundamental shocks, we expect that innovations in the inflation trend will have an effect—although not a substantial one—on output. One reason for this is that the failure to index taxes on capital gains in the United States creates a situation in which raising the inflation trend would increase the marginal tax rate on capital. Thus, a higher inflation trend creates an incentive to substitute current consumption for capital accumulation.[2]

Our results indicate that inflation-policy shocks have small effects on real GNP over both long and short horizons. In the long run, the effect is negative. Thus, a policy action that would reduce the inflation trend would be associated with a long-run increase in the level of real output but with only negligible short-run costs.

We recognize the preliminary nature of these results and discuss some qualifications below; for example, the zero restriction that we place on the long-run impact of real output shocks on inflation is not strongly supported by the data. To investigate the sensitivity of our results to this restriction, we compare them to the findings obtained in two recursive VAR systems that do not restrict the long-run relationship between inflation and output. Even though both of these systems show that output shocks have a small positive impact on the inflation trend, the estimated effect of inflation on real output is essentially the same as in our model.

I. Framework for Identification

To identify the innovations to real output and the inflation trend, we apply an approach developed by Blanchard and Quah (1989) and Shapiro and Watson (1988) to a simple two-variable system that includes inflation and output. It is assumed that there are two fundamental disturbances in the system—e_p, an inflation shock, and e_y, an output shock—and that they are uncorrelated at all leads and lags. The system is

■ 2 In a general-equilibrium framework, we would also expect work effort to be substituted for capital in the production process. Thus, labor productivity would fall, hours worked would rise, and the net effect on output would be ambiguous. See Jarrett and Selody (1982) and Bryan (1990).

TABLE 1

Unit Root Descriptive Statistics

	Augmented Dickey–Fuller T-Statistics[a]	
Series	**With a Time Trend**	**Without a Time Trend**
y_t	-0.22	-2.16
dp_t	-3.08	-2.47
dy_t	-4.23[b]	-4.24[b]
ddp_t	-3.99[b]	-4.00[b]
Residual y	-3.17	-1.07
Residual dp	-3.12	-3.14[b]

a. The Dickey–Fuller t-statistics were calculated from a regression that included six lags of the differenced data. All regressions included a constant, and there were 144 observations. See Fuller (1976, p. 373) for a tabulation of the distribution of this statistic.

b. The null hypothesis of a unit root is rejected at the 10 percent significance level.

NOTE: The series "residual y" is the residual from a regression of y_t on a constant and dp_t. The series "residual dp" is the residual from a regression of dp_t on a constant and y_t. If both y_t and dp_t are $I(1)$ and the residual contains a unit root, then y and dp cannot be cointegrated.

SOURCE: Authors' calculations.

identified by imposing the restriction that only the inflation shock may affect the inflation rate in the long run. The output disturbance is a composite of real supply and real demand shocks that may affect inflation only in the short run.[3]

Let dp denote the inflation rate and y denote the log level of output. In vector notation, let X be (dp, y) and e be (e_p, e_y). We assume that there is some n for which X follows a stationary process, given by

$$(1) \quad (1-L)^n X(t) = A(0)\,e(t) + A(1)\,e(t-1) + \dots = A(L)\,e(t),$$

where $A(L)$ is a matrix of polynomials in lag operators.

Results of unit root tests that use the augmented Dickey–Fuller procedure are presented in table 1. These statistics do not reject the null hypothesis that the elements of X are $I(1)$; that is, the inflation rate and output each contain one unit root. Therefore, we difference the inflation and output series once before estimating the model.[4] Tests for cointegration of dp and y suggest that these two variables do not share a common trend.

Under our restriction that the long-run impact of e_y on inflation is zero, the sum of the coefficients in the upper-right polynomial in $A(L)$ must equal zero. Under our assumptions, the variance $(e) = I$ and the contemporaneous effect of e on X is given by $A(0)$. Thus, our framework allows for bidirectional causality, even though the effect of an output innovation on inflation must dissipate in the long run.

Because e is not observable, $A(L)$ cannot be estimated directly. In practice, $A(0)$ can be identified in a variety of ways.[5] We use the instrumental variables approach described by Shapiro and Watson (1988), which allows the system in equation (1) to be estimated directly in autoregressive form:

$$(2) \quad B(L)(1-L)X(t) = u(t),$$

where, in general, the u's are combinations of structural disturbances that, by construction, may be correlated contemporaneously but not across time. We estimate e_p and e_y by imposing our assumptions on this autoregressive form. Because the matrix of long-run multipliers is assumed to be lower triangular, the sum of the coefficients in the upper-right polynomial in $B(L)$ must also equal zero. In practice, this restriction is imposed by including first differences of the current value and n-1 lags of output growth as regressors:

■ 3 As Blanchard and Quah show in the appendix to their 1989 paper, there are some common identification problems in low-dimension dynamic systems. For example, if the aggregate shocks are composites of many different types of disturbances, as is the case here, then our decomposition may be invalid. We intend to address this issue in subsequent work by adding more economic structure (and more variables) to the basic framework.

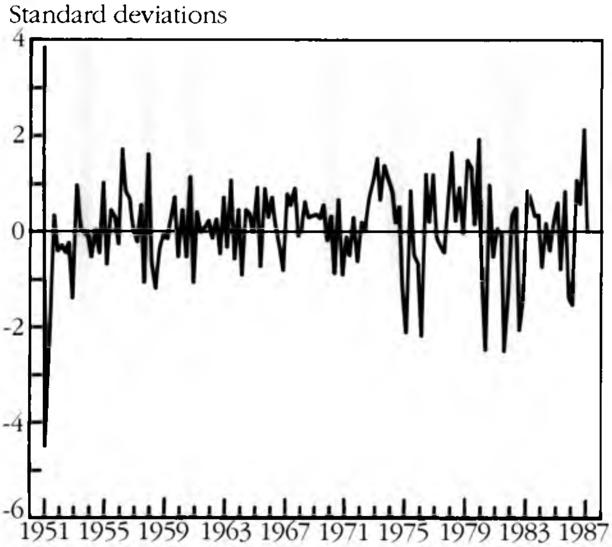
■ 4 King et al. (1989) and Shapiro and Watson (1988) find a unit root in inflation. We could not reject the null hypothesis of a unit root in inflation using the Dickey–Fuller t-statistic, but other tests, including the Dickey–Fuller normalized bias, the Phillips–Perron normalized bias, and the Phillips–Perron t-statistic, did reject the null hypothesis. See Phillips and Perron (1988), Said and Dickey (1985), and Schwert (1987) for arguments in favor of using the normalized bias t-statistic. We assume the existence of a unit root because the particular constraint that we impose to achieve identification requires that the model be specified in first differences of inflation and output. We could have specified the model in output growth and inflation, but doing so would have required policy shocks to be defined as the sole determinant of the price level. Preliminary work with this specification resulted in a time series of policy shocks that had extremely large negative effects on real GNP: An increase in the price level led to an implausibly large decline in real GNP. We suspect that any problems caused by possible overdifferencing in our model are small relative to the difficulties that would be induced by the alternative specification.

■ 5 See Blanchard and Quah (1989), Shapiro and Watson (1988), Bernanke (1986), Sims (1986), Litterman and Weiss (1985), Judd and Trehan (1989), and Boschen and Mills (1989).

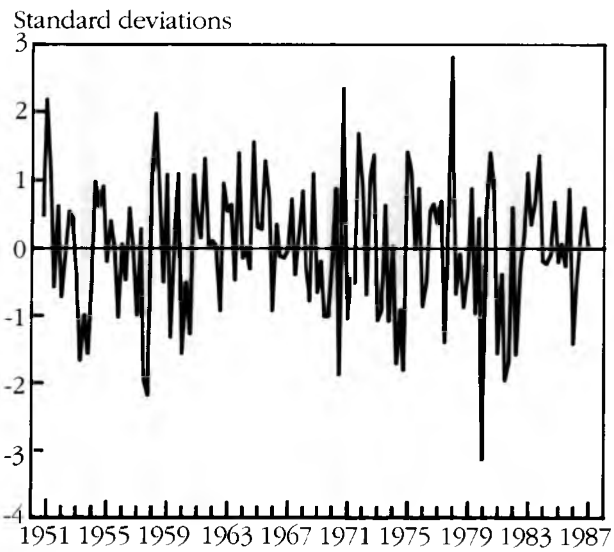
FIGURE 1

Innovations to Output and Inflation: Model Imposing Long-Run Constraint on Output Shocks

A. Estimated Innovations to the Inflation Rate



B. Estimated Innovations to Real Output



NOTE: Shaded bars indicate National Bureau of Economic Research business cycles.
SOURCE: Authors' calculations.

$$(2') \quad ddp_t = c_1 + \sum_{i=1}^{n} b_{11}^{i} ddp_{t-i} + \sum_{i=0}^{n-1} \delta_{12}^{i} ddy_{t-i} + e_{pt}$$

$$dy_t = c_2 + \sum_{i=1}^{n} b_{21}^{i} ddp_{t-i} + \sum_{i=1}^{n} b_{22}^{i} dy_{t-i} + b_{20}\hat{e}_{pt} + e_{yt},$$

where the long-run constraint has been incorporated by replacing

$$\sum_{i=0}^{n} b_{12}^{i} dy_{t-i} \quad \text{with} \quad \sum_{i=0}^{n-1} \delta_{12}^{i} ddy_{t-i}$$

in the *ddp* equation.

Contemporaneous effects of the inflation-policy shock are allowed to enter the equation for output growth.[6] Because the *ddp* equation includes a current value of the change in real output, we use an instrumental variables estimator. The instrument list includes six lagged values of *ddp* and *dy*, as well as the contemporaneous and six lagged values of the relative oil price; the price of oil is assumed to be exogenous in this model. Essentially, this two-stage procedure replaces dy_t with the ordinary least squares projection of this variable on the list of instruments. Then, by including the residual from the price equation in the output equation, the real output shock can be identified.

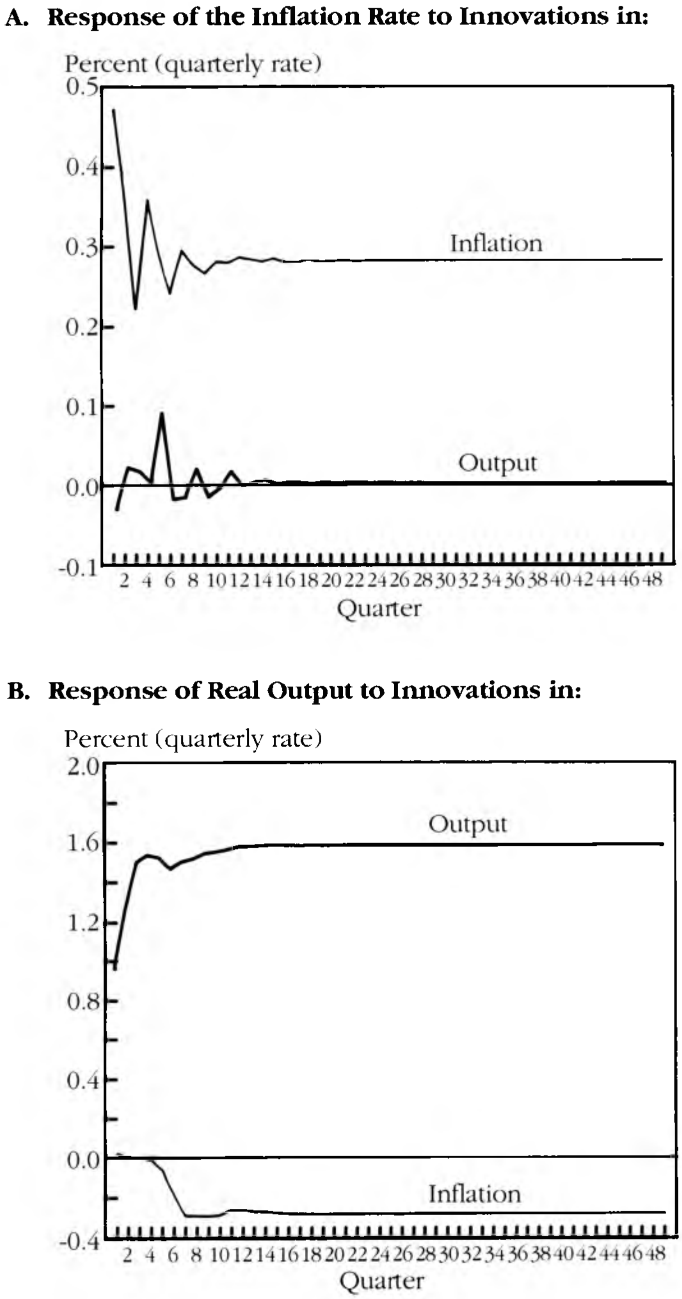
II. Results

In this paper, inflation is measured as the change in the logarithm of the Consumer Price Index (CPI), and output is measured as the logarithm of real GNP. Monthly CPI data were averaged to determine the quarterly series, and data from the sample period 1951:IQ to 1987:IIQ were used to estimate equation (2′). The estimated series for e_p and e_y are shown in figure 1. A cursory look at these two series reveals an output shock that exhibits its largest negative values in the midst of

6 This program was written in Regression Analysis Time Series (RATS) Version 3.10. We thank Mark Watson for sending us the data and program used to identify supply shocks in Shapiro and Watson (1988).

FIGURE 2

Impulse-Response Functions



SOURCE: Authors' calculations.

recessions; however, a clear cyclical pattern is not evident for the inflation-policy shock. On average, the inflation-policy shock was positive during the 1973-75 recession and negative during the 1981-82 recession.

Impulse-response functions for inflation, shown in figure 2A, indicate the response of inflation to a one-standard-deviation shock in the error vector. Taken at face value, our results suggest that the output shock has small short-run effects on inflation when the long-run effect is constrained to be zero. A standard-deviation shock to inflation (about 2 percent at an annual rate [a.r.]) raises the rate slightly more than 1 percent (a.r.) in the long run.

The impulse-response functions for output are shown in figure 2B. A one-standard-deviation shock to output results in increased output throughout the first year. In the long run, output rises by more than one and one-half times the initial shock, a gain that is nearly complete after the first year.

Inflation-policy shocks have a small but positive short-run effect on real GNP. After the second quarter, the sign becomes negative and remains that way. Thus, a policy action that lowers the inflation trend would initially have a negative effect on real output, but would raise the level in the long run.

Interpreting our results structurally, a decline in the inflation trend from 4 percent to zero would have a negligible damping effect on output in the first two quarters—less than 2/10ths of 1 percent. After the second quarter, the effect would become positive, and in the long run (after about two years), the output level would have increased about 2.5 percent.[7]

Another way to examine the dynamic effects of inflation shocks on output is to decompose the variance of output into the part caused by variation in the separate shocks. The variance decompositions for different time horizons are shown in the top section of table 2. Note that most of the variance in the two series, inflation and output, is explained by their own independent shocks. The output shock never explains more than 2 percent of the variance in the inflation rate, and the inflation shock explains almost none of the variance in output (in the long run, it accounts for only about 3 percent). Although raising the inflation trend reduces the

7 Note that this long-run effect is qualitatively and quantitatively similar to that produced by the failure to index capital-gains income for tax purposes. See Altig and Carlstrom (1990).

TABLE 2

Variance Decompositions

	Percent of Output Variance Explained by Shock to:		Percent of Inflation Variance Explained by Shock to:	
Quarter	Real Output	Inflation Rate	Real Output	Inflation Rate
Decomposition with Long-Run Constraint on Output Shocks				
1	100.0	0.0	0.5	99.5
4	100.0	0.0	0.3	99.7
8	98.7	1.3	1.2	98.8
12	98.0	2.0	0.9	99.1
20	97.5	2.5	0.6	99.4
36	97.2	2.8	0.4	99.6
49	97.1	2.9	0.3	99.7
Choleski Decomposition with Inflation as the Lead Equation				
1	99.7	0.3	0.0	100.0
4	99.5	0.5	3.4	96.6
8	97.3	2.7	10.6	89.4
12	96.3	3.7	12.7	87.3
20	95.7	4.3	14.9	85.1
36	95.4	4.6	16.7	83.3
49	95.2	4.8	17.4	82.6
Choleski Decomposition with Output as the Lead Equation				
1	100.0	0.0	0.3	99.7
4	99.9	0.1	2.3	97.7
8	98.5	1.5	8.5	91.5
12	97.8	2.2	10.1	89.9
20	97.4	2.6	12.0	88.0
36	97.2	2.8	13.4	86.6
49	97.1	2.9	13.9	86.1

SOURCE: Authors' calculations.

level of output, the increase explains little of the variation in the series.

Other research suggests that these results are not dependent on the small size of our model. King et al. (1989) also find that the permanent inflation shock never explains more than 3 percent of output variance over any horizon.[8]

III. Some Caveats

There are at least three potentially important caveats that may limit the validity of our findings. First, our identifying restriction—that the long-run impact of an output disturbance on inflation must be zero—is only weakly supported by the data. Second, the real shock is clearly an amalgamation of supply and demand shocks. Third, it may not be inappropriate to difference the inflation rate. Differencing may wash out some important short-term relationships between output and inflation. The second and third problems will be addressed in future research.

As noted above, our identifying restriction is not strongly rejected by the data. The likelihood-ratio statistic for our restricted model is 3.12, which implies that this hypothesis is rejected at the 7.7 percent significance level. Evidence provided below indicates that our empirical results do not depend critically on this identifying assumption.

To examine the implications of this restriction, we contrast our results against those obtained using a standard VAR approach; that is, we estimate equation (2) with $B(0)$ equal to the identity matrix. The contemporaneous relationships between output growth and the change in inflation are thereby captured in the variance-covariance matrix of the estimated residuals. These residuals are then transformed into orthogonal series in order to examine the dynamic consequences of independent disturbances to output and inflation.

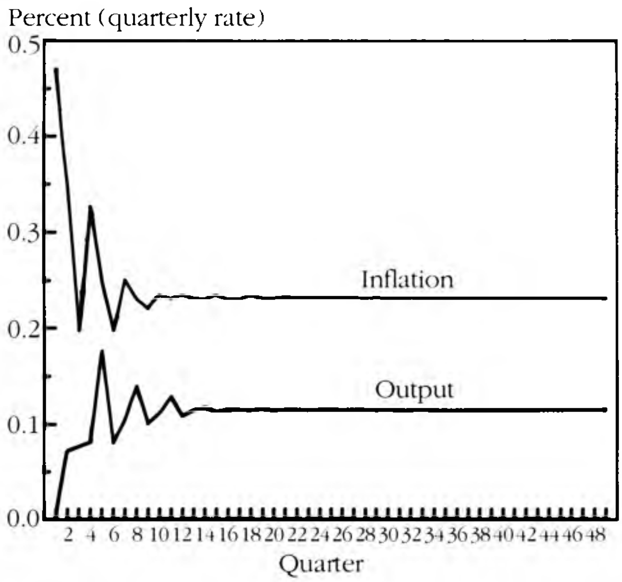
The conventional method of orthogonalization (based on the Choleski decomposition of the variance-covariance matrix) restricts the transformation matrix to be lower triangular. While this decomposition achieves an orthogonalization of the residuals, it also imposes a recursive structure on the system. In contrast to

8 See table 8b, King et al. (1989, p. 25). It should be noted that one of their identifying assumptions is that the permanent inflation shock does not affect output in the long run.

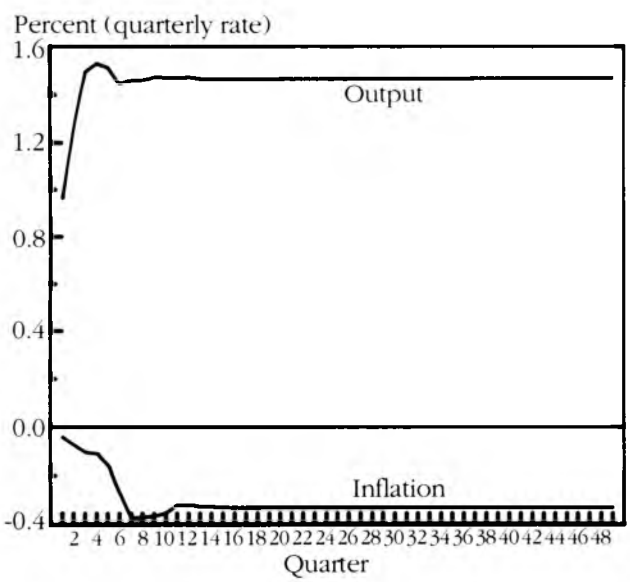
FIGURE 3

Impulse-Response Functions: Recursive System with Inflation as Exogenous Shock

A. Response of the Inflation Rate to Innovations in:



B. Response of Real Output to Innovations in:



SOURCE: Authors' calculations.

our method of identification, this decomposition places no long-run restrictions on the model. In this study, we have only two variables and hence only two potential orderings. Each alternative decomposition corresponds to an alternative ordering of the variables.

We then compare the structural implications of the two alternative recursive systems (table 2) to our restricted model. In the first specification, we assume that the first-stage residual (u_1 in matrix equation [2]) in the *ddp* equation is the structural disturbance e_p. Any correlation between u_1 and u_2 is assumed to be caused by u_1, the inflation-policy shock. The output shock, e_y, is defined as the variation in the first-stage VAR residual, u_2, that is not correlated with e_p. In the second specification, the order of the equations, and hence the assumption about the direction of causation among the contemporaneous errors, are reversed.

Figures 3 and 4 show the impulse-response functions for these two specifications. The major difference is that real shocks now affect the trend inflation rate. One explanation for this result could be that monetary policy actions are not aimed exclusively at achieving a specific inflation trend. Suppose that the Federal Reserve were following a strict money growth rule. Under such a rule, higher real output growth would result in lower inflation. However, the positive relationship in figure 3A is probably the consequence of a monetary policy that tries to smooth money market interest rates in the absence of an explicit inflation target.

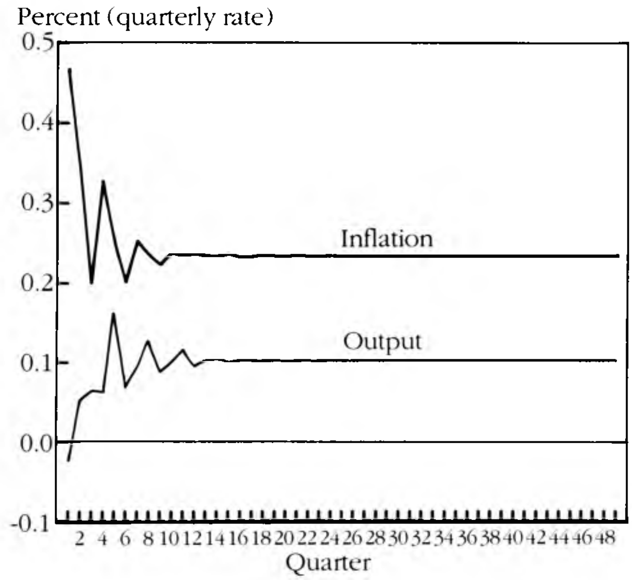
For example, whenever the investment–demand function shifts to the right, the economy experiences a transitory period of capital accumulation and relatively higher real returns. In order to prevent an increase in the federal funds rate, the Federal Reserve automatically increases the money growth rate and thereby raises the inflation rate. Unless it consciously reverses this accommodative money growth, the long-run inflation trend will be positively related to output shocks.

Although the real shock affects long-run inflation in the unrestricted model, the estimated effect of an inflation shock on output is essentially the same as in our model. Lowering the 4 percent inflation trend to zero would raise long-run output more than 3 percent in the first

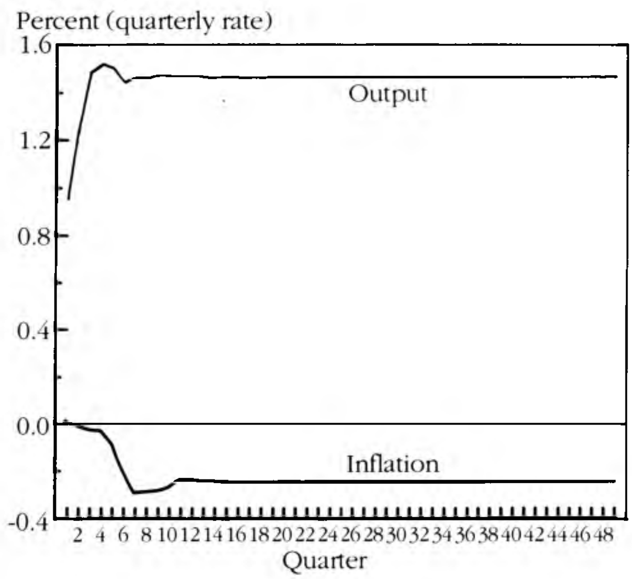
FIGURE 4

Impulse-Response Functions: Recursive System with Output as Exogenous Shock

A. Response of the Inflation Rate to Innovations in:



B. Response of Real Output to Innovations in:



SOURCE: Authors' calculations.

case and slightly more than 2 percent in the second. In both instances, the estimated long-run benefits of eliminating inflation outweigh the estimated short-run costs.

In some sense, it should not be surprising that simple time-series models of output and inflation would exhibit an inverse long-run relationship when estimated over the post-WWII period. After all, income growth was higher and inflation was lower before 1965. If, however, output has a substantial random-walk component, output growth could vary significantly by chance. Thus, the estimated inverse link between output and inflation could be spurious. Moreover, our model does not include other variables that might account for the productivity (and hence output) slowdown.

IV. Conclusion

We assume that two types of disturbances generate inflation and output dynamics—an inflation shock and an output shock—both of which are defined by identification restrictions. The inflation shock is allowed to have transitory and permanent effects on both output and inflation. Although the output shock may have only transitory effects on inflation, it may have both transitory and permanent effects on output.

We interpret the inflation shock to be a consequence of monetary policy given our restriction that it alone determines the inflation trend. Under this interpretation, the estimated policy shocks have minimal real effects. Although the results concerning the impact of inflation policy on real output are produced in a small and very simple model, we suspect that they will hold up in future extensions of this work. One indication can be found in King et al. (1989), who find similar results using a larger and theoretically richer model.

The policy implications of our findings are encouraging. Not only would a policy aimed at lowering the inflation trend raise the output level in the long run, but a structural interpretation of our VAR indicates that the short-run output loss associated with such a policy may be negligible. Our results thus suggest that there is a sequence of feasible policy actions that could lower trend inflation in such a way that the benefits would outweigh the costs. This seems to be the case on the margin—that is, where policy might succeed in offsetting inflation shocks marginally more than it did on average over the estimation period.

This study does not address the question of how such a policy would be implemented, however. Specifically, we do not consider how policymakers would control the inflation shock or to what extent they should offset it. Since the inflation-policy innovations are estimated over a period in which monetary policy largely accommodates quarterly disturbances to inflation, it is questionable whether our findings would apply in those circumstances where policy largely offsets inflation shocks. Conventional econometric evidence suggests that the underlying structure of the economy is unlikely to remain invariant to a monetary policy procedure that does not accommodate a large part of inflation shocks.

In light of this evidence, extreme policy measures could lead to greater output losses than our results suggest. Any attempt to largely offset a positive inflation shock within a quarter, however, would seem to be infeasible from a practical standpoint, if not from a technical one. On the other hand, the experience of the 1970s suggests that policymakers can also be too timid in implementing a strategy to combat inflation shocks. On balance, these results suggest that the benefits of a monetary policy aimed at achieving gradual disinflation would probably outweigh the costs. One avenue for future research would be to extend the framework presented here to include variables more closely associated with policy actions so that policy implementation issues might be investigated.

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