# THE DIFFERENTIAL REGIONAL EFFECTS OF MONETARY POLICY

Gerald Carlino and Robert DeFina*

*Abstract*—This paper examines whether monetary policy has similar effects across regions in the United States. Impulse response functions from an estimated structural vector autoregression reveal a core of regions—New England, Mideast, Plains, Southeast, and the Far West—that respond to monetary policy changes in ways that closely approximate the U.S. average response. Of the three noncore regions, one (Great Lakes) is noticeably more sensitive to monetary policy changes, and two (Southwest and Rocky Mountains) are found to be much less sensitive. A state-level version of the model is estimated and used to provide evidence on the channels for monetary policy.

IN SIMPLE textbook descriptions, monetary policy actions have a single, uniform national effect. In reality, the nation is composed of diverse regions that are linked but which might respond differently to aggregate economic shocks. For example, the large declines in crude oil prices that occurred in the mid-1980s affected energy-producing regions very differently from energy-consuming regions. The notions of a "rolling recovery" and of a "bicoastal recession" have already entered the business vocabulary and suggest that the timing and perhaps the magnitude of cycles in economic activity vary across regions. The idea that monetary policy can have varied effects across regions is a short and logical next step.

There is currently little evidence on whether and to what extent monetary policy actions have differential effects on regional economic activity. Economic theory suggests a number of ways in which monetary policy could affect regions differently. These include a traditional interest rate channel due to differing interest rate elasticities of industries coupled with differing geographical concentrations of industries. In addition, new views on credit channels for monetary policy also imply differing regional responses to policy actions. Recent research indicates that the degree to which firms are dependent on banks for credit (Bernanke and Blinder (1988) and Kashyap et al. (1993)) and the ease with which banks can adjust their balance sheets (Kashyap and Stein (1995)) might each help to determine the precise impacts of a monetary policy action.[1] If bank-dependent borrowers or small firms are regionally concentrated, the effects of federal policy may vary systematically across regions.

This paper uses a quarterly structural vector autoregression (VAR) to examine whether monetary policy shocks have symmetric effects across the eight Bureau of Economic Analysis (BEA) regions in the United States. The VAR includes the growth rates of real personal income in the BEA regions, the relative price of energy, and a monetary policy variable. The estimation period is 1958:1 to 1992:4. Impulse response functions from estimated models reveal a core of regions—New England, Mideast, Plains, Southeast, and the Far West—that respond to monetary policy shocks in ways that closely approximate the U.S. average response. Of the three noncore regions, the Great Lakes is found to be the most sensitive to monetary policy changes, while two regions (Southwest and Rocky Mountains) are found to be the least sensitive.[2] The core and noncore results are robust with respect to alternative measures of monetary policy (federal funds rate, nonborrowed reserves, and a narrative measure developed by Boschen and Mills (1995)), measures of economic activity (real personal income growth and employment growth), and model specification (variables expressed as levels and as growth rates).

This paper also provides evidence on the reasons for differential responses to monetary policy. The evidence is based on additional estimates of state-level structural VARs. State-level VARs provide 48 estimated impulse responses as opposed to only eight from the regional VARs. We find that the size of a state's response to a monetary policy shock is positively related to the share of manufacturing in the state's gross product (evidence for an interest rate channel). We find some evidence that states containing a relatively larger concentration of small firms tend be more responsive to monetary policy shocks than states with smaller concentrations of small firms (evidence for a broad credit channel); however, these findings are mixed.

## II. Sources of Regional Differences in the Effects of Monetary Policy

The literature on the monetary transmission mechanism suggests several reasons why Fed policy actions might have differential regional effects. These include regional differences in the mix of interest-sensitive industries, regional differences in the mix of large versus small firms, and regional differences in the abilities of banks to alter their balance sheets.

### A. *Regional Differences in the Mix of Interest-Sensitive Industries*

It is well known that interest rate elasticities differ across industries. The different interest sensitivities across industries may interact with differing industry mixes across regions and provide a way for monetary policy to have differential regional effects. As table 1 shows, industry mix

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* Federal Reserve Bank of Philadelphia; and Villanova University, respectively.

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[1] See Hubbard (1995) for a critical review of the credit channel view of monetary policy.

[2] The noncore regions accounted for 32% of total gross state product in the United States in 1980 and for 30% of U.S. population.

TABLE 1.—PERCENT OF REGIONAL GROSS STATE PRODUCT ACCOUNTED FOR BY MAJOR INDUSTRY (1985–90)[a]

| Region | Agri | Mining | Const | Mfg | T&PU | Trade | FIRE | Service | Govt | Pop (% of US) |
|---|---|---|---|---|---|---|---|---|---|---|
| NE | 0.97 | 0.10 | 5.17 | 21.0 | 6.86 | 16.9 | 20.4 | 19.3 | 9.3 | 5.44 |
| ME | 0.75 | 0.27 | 4.69 | 16.8 | 9.0 | 16.0 | 21.0 | 19.9 | 11.7 | 18.6 |
| GL | 1.58 | 0.64 | 3.94 | 27.0 | 8.91 | 16.4 | 16.4 | 15.6 | 9.51 | 18.4 |
| PL | 4.80 | 0.88 | 4.07 | 19.6 | 10.0 | 17.3 | 16.7 | 15.3 | 11.3 | 7.57 |
| SE | 2.04 | 2.44 | 4.91 | 19.9 | 9.53 | 17.2 | 15.6 | 15.2 | 13.2 | 23.3 |
| SW | 1.77 | 7.02 | 4.73 | 15.2 | 10.4 | 17.0 | 16.2 | 15.6 | 12.1 | 9.42 |
| RM | 2.94 | 5.50 | 4.66 | 12.6 | 11.0 | 16.2 | 17.0 | 16.2 | 12.7 | 2.90 |
| FW | 2.33 | 1.62 | 4.88 | 15.8 | 8.1 | 16.7 | 19.3 | 19.1 | 12.2 | 14.4 |
| US | 1.88 | 1.85 | 4.63 | 19.2 | 9.06 | 16.7 | 17.9 | 17.2 | 11.7 | — |

Notes: NE = New England, ME = Mideast, GL = Great Lakes, PL = Plains, SE = Southeast, SW = Southwest, RM = Rocky Mountains, FW = Far West.
[a] The percent by industry averaged over the 1985–90 period.
Source: Compiled form BEA data.

differs widely across regions. For example, manufacturing, which is thought to be an interest-sensitive sector, accounted for 27% of real gross state product (GSP) in the Great Lakes region, on average, during the 1985–1990 period, but less than 13% of the Rocky Mountains region's real GSP. The share of real GSP attributable to construction, another interest-sensitive sector, also varied across regions during this period, although to a lesser extent than in manufacturing. Compounding these differences are interregional input–output relationships, which can transmit localized responses differently across regions.

### B. *Regional Differences in the Mix of Large versus Small Firms*

Regional differences in the proportion of large and small firms and the sources of credit available to each also could lead to different regional responses to monetary policy. According to the "credit view" of monetary policy discussed in Bernanke and Blinder (1988), Bernanke (1993), and Gertler and Gilchrist (1993), monetary policy affects economic activity by directly affecting banks' abilities to provide loans. Moreover, information costs and transaction costs often require small firms to deal with financial intermediaries, primarily banks, to meet their credit needs. Large firms, by contrast, usually have greater access to external, nonbank sources of funds. Consequently, activity in a region that has a high concentration of small firms could be especially sensitive to Fed policy.

Gertler and Gilchrist (1993) and Oliner and Rudebusch (1995) discuss an alternative way by which monetary policy changes can differentially affect large versus small firms. This avenue, referred to as the "broad credit view," concerns the total credit (bank loans, trade credit, commercial paper, etc.) available to different-sized borrowers. According to Oliner and Rudebusch (1995), the "broad credit channel emphasizes that information asymmetries between borrowers and lenders may increase the costs of all forms of debt after a [contractionary] monetary policy shock. Given the relative severity of information problems for small firms, the increase in the cost of external finance for these firms will likely be particularly sharp." Oliner and Rudebusch (1995) provide evidence that a monetary contraction shifts financing of all types from small firms to large firms.

As table 2 shows, the percentage of small firms (defined as regional firms with fewer than 250 employees) varies widely across regions. It ranges from a low of 66% to 67% in the New England, Mideast, and Great Lakes regions to a high of about 82% in the Rocky Mountains region.

### C. *Regional Differences in the Abilities of Banks to Alter their Balance Sheets*

Kashyap and Stein (1995) have suggested that Fed policy actions can have varied effects on different banks' abilities to make loans. During periods of tight monetary policy when bank reserves are restricted, some banks can find alternative sources of funding for deposits and loans (by issuing large denomination CDs, for example) more cheaply and easily than others. Such lending by banks will be less sensitive to monetary policy changes. Kashyap and Stein (1995) propose that bank size largely explains differences in financing abilities, with large banks having more funding options available than small banks. Thus regions in which a disproportionately large share of bank loans are made by small banks might respond more to monetary policy shifts than regions in which a large share of loans are made by the nation's large banks.

Kashyap and Stein (1995) define small banks as those with total assets at or below a given percentile—they use,

TABLE 2.—THE SHARE OF TOTAL REGIONAL EMPLOYMENT ACCOUNTED FOR BY A REGION'S SMALL FIRMS[a]

| | Percent Small Firms |
|---|---|
| New England | 66.2 |
| Mideast | 67.0 |
| Great Lakes | 66.5 |
| Plains | 77.1 |
| Southeast | 73.3 |
| Southwest | 76.9 |
| Rocky Mountains | 82.4 |
| Far West | 77.9 |

Notes: [a] Small firm are those with fewer than 250 employees in 1981.
Source: Compiled from County Business Patterns.

TABLE 3.—SHARE OF TOTAL OF LOANS MADE BY A REGION'S SMALL BANKS, DECEMBER 31, 1994[a]

| | 75th Percentile | | 90th Percentile | |
|---|---|---|---|---|
| | Total[b] | No Holding[c] | Total[b] | No Holding[c] |
| New England | 3.3 | 3.1 | 8.4 | 7.7 |
| Mideast | 1.5 | 1.4 | 4.4 | 3.5 |
| Great Lakes | 10.9 | 7.0 | 21.4 | 11.8 |
| Plains | 31.6 | 20.2 | 44.3 | 38.5 |
| Southeast | 12.4 | 8.8 | 21.7 | 14.1 |
| Southwest | 16.3 | 13.1 | 26.2 | 20.0 |
| Rocky Mountains | 19.2 | 12.0 | 33.9 | 18.7 |
| Far West | 4.0 | 3.8 | 8.4 | 7.6 |

Notes: [a] The percentage of the loans made by a region's banks that are at or below the 75th percentile and 90th percentile terms of total assets in the nation.
[b] The percentage of total loans make by all banks in the 75th and 90th percentile.
[c] The percent of loans made by banks in the 75th and 90th percentile excluding banks that are members of a bank holding company.
Source: Compiled from Call Reports.

alternatively, the 75th, 90th, 95th, or 98th percentile. Boyd and Gertler (1994), in contrast, classify a bank as small if its assets are less than $300 million. Table 3 shows the regional distribution of loans for the nation's banks that are at or below the 75th percentile and the 90th percentile in terms of total assets at the end of 1994. These breakdowns are constructed using Federal Reserve Call Report data. Since the asset size of the 90th percentile was just under $300 million in 1994, this grouping is also equivalent to Boyd and Gertler's (1994) definition.[3] Whether we look at all small banks or only small banks that are *not* members of a bank holding company, the regional distribution of loans by small banks appears highly unequal, suggesting that monetary policy could have differential regional effects for the reasons proposed by Kashyap and Stein (1995).[4]

The effect of the differences in regions' reliance on small banks will be diluted if bank-dependent borrowers can obtain credit from sources outside their own regions. However, there is evidence that banking markets tend to be segmented along regional lines. Moore and Hill (1982) note that since banks can identify and monitor local investment projects more efficiently than banks and investors in other regions, it will be less costly for households and small firms to borrow from local banks. Local banks also tend to offer favorable loan terms to local firms in anticipation of future deposit business. Hanson and Waller (1996) look at the relationship between growth in state real personal income and growth in a variety of state bank credit measures, where all variables are measured relative to national counterparts. They find a significant relationship between bank lending and real growth at the state level. They conclude that financial markets are not well integrated at the subnational level, implying that a bank credit channel may exist. While shedding light on a possible channel of monetary policy, Hanson and Waller's study does not examine whether such policy has differential effects on economic activity across states. Indeed, little direct evidence exists on the differential regional effects of monetary policy.

## III. Literature Review

Some researchers have investigated the effects of monetary policy on interregional banking flows as opposed to economic activity. Studies by Miller (1978) and Bias (1992) have found that Fed policy actions do affect regional banking flows differentially. More typical of earlier studies is the use of a reduced-form, St. Louis-type equation that regresses personal income, earnings or employment on the high-employment federal government revenues, high-employment federal government expenditures, and the national money supply. These models are applied at the regional level to test the monetarist proposition that monetary policy has an important impact on nominal income (Toal (1977), Garrison and Chang (1979), Beare (1976), Mathur and Stein (1980), and Garrison and Kort (1983)).

The Garrison and Chang (1979) and Toal (1977) studies are most closely related to ours, in that they look at the effect of monetary policy on income variables for the eight BEA regions of the United States. Garrison and Chang (1979) study regional manufacturing earnings during the 1969–1976 period and find that monetary policy has differential effects across regions, with an especially large impact in the Great Lakes region and a rather small impact in the Rocky Mountain region. Like Garrison and Chang (1979), Toal (1977) concludes that differences in regional responses to monetary policy changes existed in the 1952–1975 period, with relatively larger responses in the Mideast, Great Lakes, and Southeast regions, and relatively weak responses in the Rocky Mountains and New England regions. Beare (1976) uses data for the predominately agrarian Canadian prairie provinces during the 1956–1971 period and finds that different provinces respond differently to money supply changes. Garrison and Kort (1983) investigate the impacts of monetary policy on state-level employment. For the 1960–1978 period, they find that the states comprising the Great Lakes region are generally the most responsive to money supply changes, while states in the Rocky Mountain region are least responsive.[5]

A shortcoming with existing studies is their attempt to measure monetary policy impacts region by region without accounting for feedback effects among regions (i.e., monetary policy directly affects region $i$, and through trade with region $j$, monetary policy indirectly affects region $j$, and vice versa).[6] Carlino and DeFina (1995) use VARs to document

[3] Data in table 3 are presented for all banks as well as for only those banks that are *not* members of a multibank holding company.

[4] As a member of a bank holding company, a small bank can issue large denomination (uninsured) CDs at more favorable rates because it can rely on the financial strength of the larger bank holding company. Note also that although the data indicate the location of the lending bank, they do not specify the locations of the borrowers. One reason for focusing on the lending patterns of small banks is that they tend to specialize in loans to local customers. Large banks tend to make loans outside their local market.

[5] Mathur and Stein (1980) question the usefulness of reduced-form, St. Louis-type equations for analyzing the effects of policy shocks.

[6] One exception is the paper by Taylor and Yücel (1996) in that they use VAR analysis to estimate the effects of both monetary and fiscal policy for

the importance of feedback effects among U.S. regions. It seems that a system technique, such as VAR, is better suited for gathering evidence on business-cycle dynamics.

## IV. Empirical Approach

### A. The Model

The analysis focuses on the dynamic behavior of an $n \times 1$ covariance-stationary vector,

$$Z_t = [\Delta y_{1,t}, \Delta y_{2,t}, \ldots, \Delta y_{n-1,t}, \Delta p_{e,t}, m_t]'$$

where $\Delta y_{i,t}$ is the growth rate of real personal income in region $i$ at time $t$, $\Delta p_{e,t}$ is the growth rate of the relative price of energy, and $m_t$ is a stationary variable measuring monetary policy actions at time $t$. The dynamics of $Z_t$ are represented by a VAR,

$$AZ_t = B(L)Z_{t-1} + e_t \tag{1}$$

where $A$ is an $n \times n$ matrix of coefficients describing the contemporaneous correlations among the variables; $B(L)$ is an $n \times n$ matrix of polynomials in the lag operator $L$; and $e_t = [\epsilon_{1,t}, \epsilon_{2,t}, \ldots, \epsilon_{n-1,t}, \epsilon_{p,t}, \epsilon_{m,t}]$ is an $n \times 1$ vector of structural disturbances.

Solving for $Z_t$ produces the following reduced-form system:

$$Z_t = C(L)Z_{t-1} + u_t \tag{2}$$

where $C(L) = A^{-1}B(L)$ is an infinite-order lag polynomial, and $u_t = A^{-1}e_t$ describes the relationship between the model's reduced-form residuals and the model's structural residuals.[7]

### B. Data

The study employs BEA quarterly data on real personal incomes by major region for the period 1958:1 to 1992:4 (see appendix A for regional definitions).[8,9] Empirical research on the regional effects of Fed actions requires an explicit indicator of monetary policy. Following Bernanke and Blinder (1992) and Cook and Hahn (1988), among others, we use the federal funds rate. To gauge the robustness of the results to the choice of policy indicator, we also use nonborrowed reserves and a narrative indicator of policy (Boschen and Mills (1995)).

Finally, to account for aggregate supply shocks, an energy price variable is included in the system. This variable is calculated as the producer price index for fuels and related products, and power relative to the total producer price index. It is especially important to account for oil price shocks because the Fed tightened when the relative price of oil increased in 1973 and in 1979 and eased when the relative price of oil declined in 1986.

TABLE 4.—AUGMENTED DICKEY-FULLER TESTS OF SYSTEM VARIABLES

| REGION | Level[a] | Growth Rate[b] |
|---|---|---|
| New England | −1.279 | −5.420** |
| Mideast | −1.590 | −7.046** |
| Great Lakes | −1.655 | −6.647** |
| Plains | −1.523 | −8.908** |
| Southeast | −0.506 | −7.010** |
| Southwest | −0.955 | −6.978** |
| Rocky Mountains | −0.006 | −8.013** |
| Far West | −1.149 | −7.200** |
| United States | −0.496 | −5.490** |
| Relative Price of Oil | −1.724 | −9.157** |
| Fed Funds Rate | −2.440 | −9.769** |

[a] Equations include an intercept and time trend.
[b] Equations include an intercept term.
** Indicates significance at the 1% level. Critical values are found in Fuller (1976).

### C. Unit-Root Tests

The variables used in the estimation must be stationary so that standard statistical theory applies. Table 4 reports the results of the augmented Dickey–Fuller (ADF) unit-root test applied to the levels and first-differences of the system's variables. Data on regional and U.S. real personal incomes, the relative energy price, and nonborrowed reserves are in logs. As table 4 shows, the unit-root null cannot be rejected for any of the data series, although stationarity is achieved by first-differencing. Thus the VARs to be estimated include the stationary first differences of regional personal income, the relative price of energy, and the federal funds rate.

### D. Empirical Estimates

The impact of Fed actions are summarized using impulse response functions.[10] Assuming the primitive innovations $e_t$ are identified, impulse response functions $Z_t$ are calculated directly from equation (2) as

$$Z_t = [I - C(L)L]^{-1}A^{-1}e_t = \Theta(L)e_t \tag{3}$$

four large states during the 1982–1995 period. The Taylor and Yücel (1996) study is limited in that it looks at only four states.

[7] The problem of identifying the structural shocks $e_t$ from the VAR reduced-form residuals $u_t$ and their variances is taken up later. The solution depends on identification restrictions placed on the $A$ matrix and on the variance–covariance matrix of structural errors.

[8] Real incomes are calculated by deflating each region's nominal income with the national consumer price index (CPI). Ideally, regional incomes should be deflated using regional price deflators. However, such deflators are not available. Consumer price indexes do exist for many of the metropolitan areas in the various regions. We found a high degree of correlation in consumer price inflation across these metropolitan areas during the 1958:1–1986:4 period.

[9] An alternative variable for measuring regional economic activity is GSP. GSP data are published annually and, as such, are not as well suited to the analysis of regional business cycles.

[10] See Sims (1980) for a discussion.

TABLE 5.—IMPULSE RESPONSE OF REGIONAL PERSONAL INCOME GROWTH*,**

| Period | NE | ME | GL | PL | SE | SW | RM | FW |
|---|---|---|---|---|---|---|---|---|
| 1 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 |
| 2 | −0.0818[a] | −0.1039[a] | −0.0698[b] | −0.0907 | −0.0638[a] | −0.0357 | 0.0301 | 0.0815[a] |
| 3 | −0.2297[a] | −0.2156[a] | −0.3239[a] | −0.2440[a] | −0.3132[a] | −0.2012[a] | −0.2216[a] | −0.2593[a] |
| 4 | −0.0924[b] | −0.0878 | −0.2409[a] | −0.1930[b] | −0.0894[a] | −0.0242 | −0.1359[b] | −0.0763[b] |
| 5 | −0.1316[a] | −0.1063[b] | −0.1748[a] | −0.2202[b] | −0.1343[a] | −0.0601 | −0.0600 | −0.1223[a] |
| 6 | −0.1527[a] | −0.1525[a] | −0.1886[a] | −0.0950 | −0.1404[a] | −0.0791[b] | −0.0717 | −0.1649[a] |
| 7 | −0.0520 | −0.0260 | −0.1252[a] | −0.1309 | −0.0700[b] | −0.0271 | −0.0729 | −0.0857[b] |
| 8 | −0.0058 | −0.0449 | −0.0525 | −0.0854 | −0.0504 | −0.0228 | −0.0526 | −0.0579 |
| 9 | −0.0401 | −0.0339 | −0.0661 | −0.0553 | −0.0521 | −0.0198 | −0.0091 | −0.0627[b] |
| 10 | −0.0203 | −0.0217 | −0.0406 | −0.0077 | −0.0195 | 0.0055 | 0.0044 | −0.0351 |
| 12 | −0.0003 | −0.0063 | 0.0006 | −0.0073 | −0.0018 | −0.0013 | 0.0149 | −0.0112 |
| 14 | 0.0113 | 0.0077 | 0.0132 | 0.0238 | 0.0221 | 0.0153 | 0.0225 | 0.0023 |
| 16 | 0.0061 | 0.0041 | 0.0098 | 0.0154 | 0.0120 | 0.0041 | 0.0133 | 0.0028 |
| 20 | 0.0018 | 0.0024 | 0.0076 | 0.0067 | 0.0071 | 0.0054 | 0.0034 | 0.0043 |

Notes: * Change in log levels of a region's real personal income resulting from a one-standard-deviation increase in the federal funds rate change.
** NE = New England, ME = Mideast, GL = Great Lakes, PL = Plains, SE = Southeast, SW = Southwest, RM = Rocky Mountains, and FW = Far West.
[a] Indicates significance at the 1 percent level; [b] Indicates significance at the 5 percent level.

where

$$\Theta(L) = \sum_{l=0}^{L} \Theta_l L^l \tag{4}$$

and $\Theta_l$ is a $k \times k$ matrix of structural parameters.

## V. Empirical Results

The elements of the $B(L)$ and $A$ matrices are estimated using the two-step procedure described in Bernanke (1986). In the first step, ordinary least squares (OLS) is used to obtain estimates of the reduced-form errors $u_t = A^{-1}e_t$ from the dynamic simultaneous equation model (2). Next, sufficient restrictions are placed on the variance–covariance matrix of structural errors (i.e., orthogonality and normalization restrictions) and on the matrix of contemporaneous correlations $A$ to achieve identification. Then, given estimates of $A$, estimates of $B(L)$ are derived from the relationship $C(L) = A^{-1}B(L)$ where $C(L)$ comes from the estimated reduced-form equation (2). Estimates of $A$ also allow estimates of the structural errors $e_t$, as implied by the relationship $u_t = A^{-1}e_t$.

Three sets of restrictions are placed on the matrix $A$. First, we follow Carlino and DeFina (1995) by assuming there is some period for which a region-specific shock affects only the region of origin, although eventually it may spill over into other regions. That is, a shock to a region's real personal income growth affects other regions' growth only after a one-period lag. Second, Fed policy actions are assumed to affect regional income growth no sooner than with a one-period lag. This assumption conforms with widely accepted views about monetary "impact lags." Similarly, shocks to the relative price of energy are assumed to have no contemporaneous effect on regional incomes. Third, we assume that neither regional income growth nor policy actions affect relative energy prices contemporaneously.

Taken together, these restrictions imply particular values for the $A$ matrix in equation (1). Let $i$ and $j$ index regions, let $m$ index the monetary policy variable, and let $p$ index the relative price of energy. Then $a_{ij} = 1$ for $i = j,p,m$; and $a_{ij} = 0$ for $i \neq j,p$. There are no restrictions on the elements $a_{mj}$ (allowing the Fed to respond contemporaneously to region-specific shocks) and $a_{mp}$ for $m \neq j,p$. Given the resulting $A$ matrix, the elements of $e_t$ are (over)identified. Essentially regions are subject to three types of shocks in the model: monetary policy shocks, relative energy price shocks, and an aggregate of all other shocks. The latter incorporates aggregate supply shocks other than energy price shocks, aggregate demand shocks other than monetary policy shocks, and region-specific demand and supply shocks. By construction, the primitive regional shocks are orthogonal.[11] Four lags of each variable are used in the estimation, a sufficient number to eliminate serial correlation in the errors.[12] Given these estimates, impulse responses are calculated using equation (3).

### A. *Impulse Response Functions*

Table 5 presents the point estimates for the impulse responses of regional personal income growth resulting

[11] Alternative identification schemes exist. Two commonly used approaches are a Cholesky decomposition of the variance–covariance matrix and the imposition of long-run neutrality restrictions such as those used in Blanchard and Quah (1989), Shapiro and Watson (1988), and Gali (1992). Each has shortcomings. The Cholesky decomposition requires explicit assumptions about the ordering of equations. The system to be estimated contains over 3.5 million possible orderings, and there is no definitive approach to choosing which is best. See Cooley and Leroy (1985) for a detailed critique of the Cholesky approach. Alternatively, long-run restrictions are defended as theoretically appealing and more "credible" than "ad hoc" contemporaneous restrictions. Faust and Leeper (1997), however, detail several potential drawbacks, such as that long-run restrictions cannot ensure reasonable short-run dynamics.

[12] Ljung-Box $Q$ test statistics indicate that the null hypothesis of white noise errors cannot be rejected at the 5% level of significance for any of the system's equations. The choice of lag length was also addressed in a restricted way using the Akaike and Schwartz information criterion. That is, the number of lags of all variables in a particular equation was sequentially varied from one to eight. These criterion suggested that an optimal lag length was in the neighborhood of two to five quarters, depending on the equation. Thus the choice of four lags appears appropriate on several grounds.

from a one-standard-deviation (83 basis point) unanticipated increase in the federal funds rate change.[13] Real regional personal income growth generally declines during the first five to six quarters following the policy shock and generally become statistically insignificant thereafter.[14] The impulse responses indicate that unanticipated monetary policy shocks typically have their maximum impact on real personal income growth after three quarters. For example, a policy innovation results in a 0.0818% decrease in real personal income growth in the New England region in the second quarter. The effect of the policy shock then builds to a maximum of 0.2297 in quarter three and oscillates during quarters four through six. In no instance does monetary policy have a significant influence on regional personal income growth after nine quarters.[15]

Figure 1 displays the cumulative impulse responses derived from table 5. The cumulative impulse response of U.S. aggregate personal income is included as a benchmark for comparison. We found that an unexpected one-standard-deviation increase in the federal funds rate reduces real *growth* temporarily and, thus, leaves the *level* of real personal income below what it otherwise would have been for about two years. Subsequently, the impact on the level of income generally becomes statistically insignificant. The model treats tightening and easing of the federal funds rate symmetrically, so that an unexpected cut in the funds rate temporarily raises real personal income relative to what it would have been otherwise.

Interestingly, not all regions respond by the same magnitude. Several regions generally respond to monetary policy surprises with a magnitude and timing similar to those of the national economy. Specifically, the responses of income in five regions (New England, Mideast, Plains, Southeast, and Far West), called "core" regions, mirror the national response.[16] Core regions accounted for 68% of aggregate 1980 GSP in the United States and for 70% of the total U.S. population.

Other regions (Great Lakes, Southwest, and Rocky Mountains), called noncore regions, show magnitudes of monetary policy effects quite different from the magnitudes for the national economy. The noncore regions accounted for 32% of total 1980 GSP in the United States and for 30% of the U.S. population. In the Great Lakes region, personal income is more responsive to monetary policy shocks than is the U.S. average. The Great Lakes region accounts for 18% of total GSP. In two other regions (Rocky Mountains and Southwest) personal income is much less responsive to monetary policy shocks than is the U.S. average. Together these two regions account for 14% of aggregate GSP and 12% of the U.S. population. The Rocky Mountains region is the smallest, accounting for only 3% of aggregate GSP produced and for only 3% of the nation's population.

The core versus noncore dichotomy is confirmed by formal statistical tests. Monte Carlo simulations (500 replications) were used to calculate *t*-statistics for the difference between each region's estimated cumulative response and the estimated cumulative response for the nation. The *t*-statistics indicate that the cumulative responses for the five regions comprising the core are not significantly different (at the 5% level) from the national cumulative response at the 5, 10, 15, and 20 quarter horizons. In contrast, the cumulative responses of the noncore regions are generally found to be significantly different (at the 5% level) from the cumulative response of the nation. The only exceptions occur for the cumulative response of the Rocky Mountains region, which is insignificantly different from national cumulative response at the 15 and 20 quarter horizons.

### B. *Robustness Tests*

Alternative specifications were estimated to examine the sensitivity of the results to alternative measures of monetary policy, measures of regional economic activity, and methods for making the data stationary. These alternative specifications are compared to the specification employing the change in the federal funds rate as the monetary policy variable, referred to as the "baseline" system.

*Measures of Monetary Policy Actions:* To check the robustness of the results to the choice of monetary policy indicator, we substituted the stationary first-difference of logged nonborrowed reserves for the federal funds rate change in the VAR. Figure 2 shows the effect of a one-standard-deviation positive shock (2.7%) to the growth rate of nonborrowed reserves. With the exception of the New England region, the core–noncore distinction found for the federal funds rate arises for nonborrowed reserves as well. With nonborrowed reserves as the policy variable, the New England region behaves more like the Great Lakes region, a noncore region.

One problem with traditional money-market measures of monetary policy, such as the federal funds rate and nonborrowed reserves, is that observed changes in these variables reflect forces other than the decisions of the monetary authority, and that the changes can have different interpretations depending on the operating procedure in place. To minimize these difficulties, some researchers have relied on an alternative "narrative approach" that attempts to identify monetary policy shocks by looking at evidence derived from

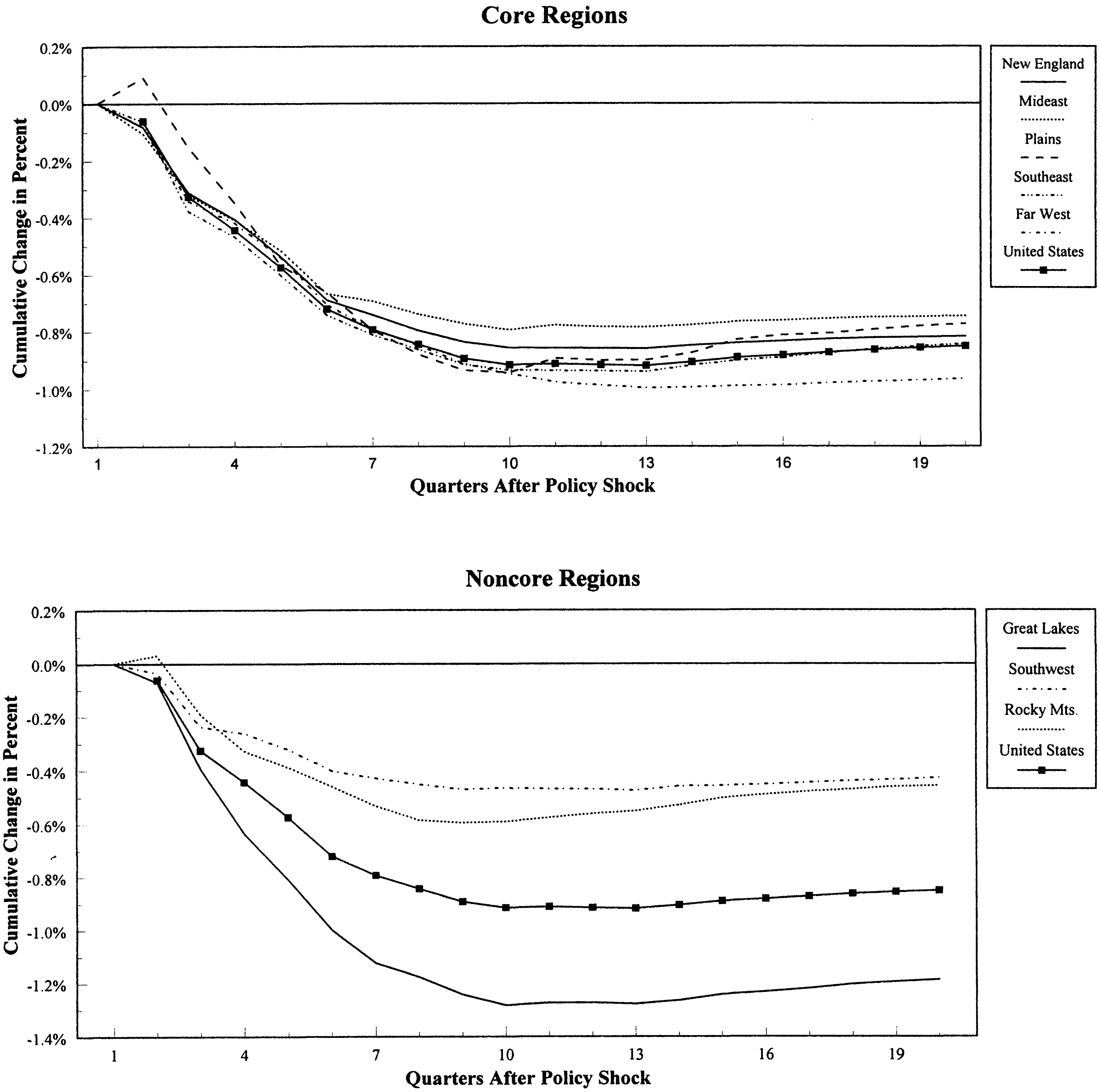
[13] Recall that the VAR includes an equation that predicts changes in the federal funds rate on the basis of a year's worth of past data for each of the following variables: the change in the federal funds rate; real personal income growth in each of the eight major regions; and the change in the relative price of energy. Unexpected changes in the federal funds rate are measured by taking the difference between the actual and predicted change. Unexpected changes in the federal funds rate are used to proxy monetary policy surprises in the policy simulations that follow. The analysis assumes that unexpected changes in the federal funds rate arise only from policy surprises. The robustness of the approach is examined in a subsequent section of this paper.

[14] Monte Carlo simulations were performed to generate standard errors.

[15] We leave to the reader a detailed examination of the results of the impulse response functions for the other regions.

[16] The terminology "core regions" is taken from Bayoumi and Eichengreen (1993), who use it in a related but somewhat different way.

FIGURE 1.—CUMULATIVE IMPULSE RESPONSE OF REAL PERSONAL INCOME TO FED FUNDS SHOCK, CORE AND NONCORE REGIONS




Notes: The estimated impulse response functions were computed from ten-variable VARs containing real personal income growth in each of the eight major BEA regions; growth in the producer price index for fuels and related products relative to the overall producer price index; and a monetary policy variable equal to the first difference of the federal funds rate. Shocks constitute one-standard-deviation unanticipated increases in the monetary policy variable.

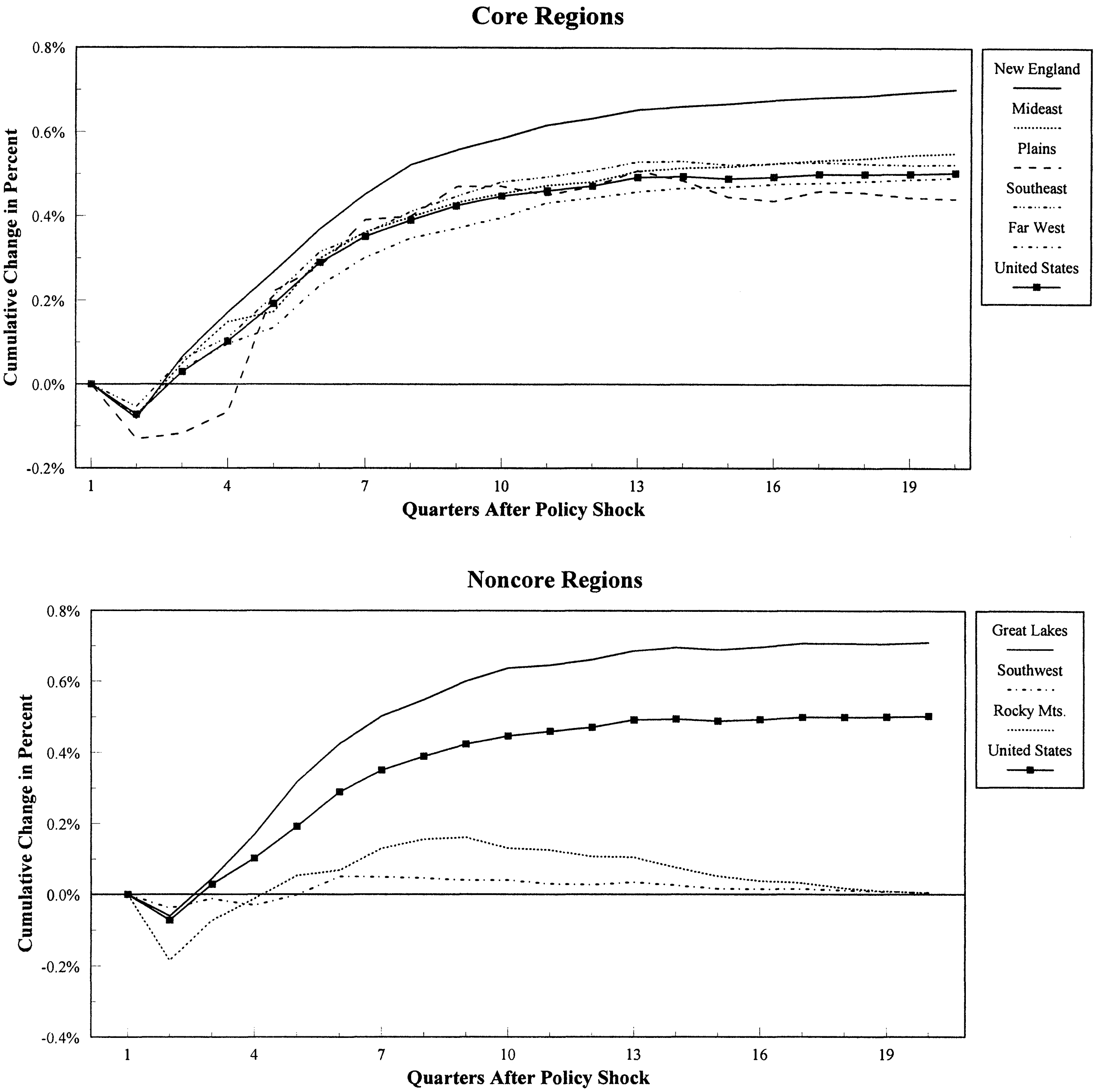
the Federal Open Market Committee (FOMC) policy directives (Romer and Romer (1989, 1990), Boschen and Mills (1995), Ramey (1993), and Ball and Croushore (1995)).[17] Several indexes exist and we employ one developed by Boschen and Mills (1995) that ranks monetary policy on a numerical scale from −2 (large emphasis on inflation control) to +2 (large emphasis on promoting real growth).[18]

Figure 3 reports the cumulative impulse response of real regional personal income resulting from a one-standard-

[17] The narrative approach has a long history, dating to the work of Friedman and Schwartz (1963).

[18] Boschen and Mills (1995) report that alternative narrative measures, such as that constructed by Romer and Romer (1989), are highly correlated with their measure. See also Hoover and Perez (1994) for a discussion of the limitations of narrative measures.

FIGURE 2.—CUMULATIVE IMPULSE RESPONSE OF REAL PERSONAL INCOME TO NONBORROWED RESERVES SHOCK, CORE AND NONCORE REGIONS



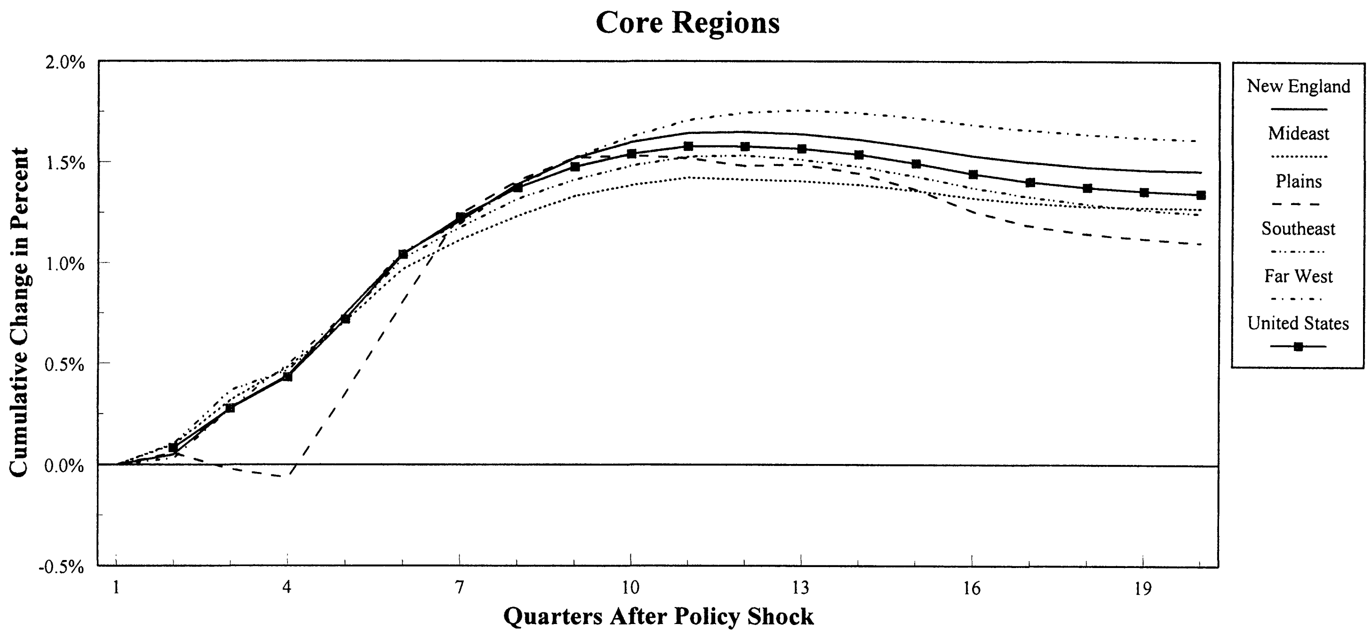

Notes: The estimated impulse response functions were computed from ten-variable VARs containing real personal income growth in each of the eight major BEA regions; growth in the producer price index for fuels and related products relative to the overall producer price index; and a monetary policy variable equal to the growth in nonborrowed reserves. Shocks constitute one-standard-deviation unanticipated increases in the monetary policy variable.

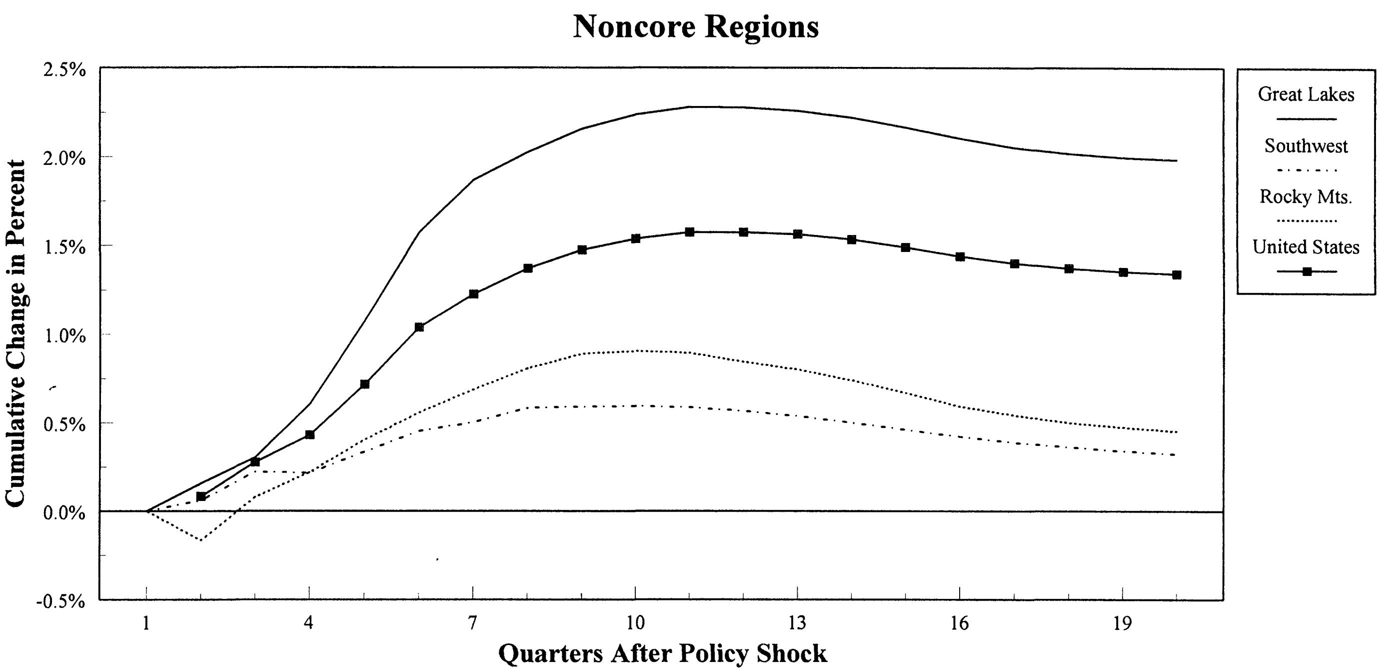
deviation change in the level of the Boschen and Mills index.[19] The estimates identify the identical core–noncore dichotomy associated with federal funds rate shocks.

[19] The level of the Boschen–Mills index is used in the growth rate specification of the VAR because this index is stationary in levels.

An alternative procedure for dealing with the issue of identifying monetary policy shocks is to use a financial variable, like the federal funds rate, but to add variables to the model that the Fed includes in its policy information set. Doing so increases the likelihood that any measured policy shocks are truly exogenous, and not endogenous reactions to omitted variables. Following this logic, two economywide

FIGURE 3.—CUMULATIVE IMPULSE RESPONSE OF REAL PERSONAL INCOME TO B/M SHOCK, CORE AND NONCORE REGIONS



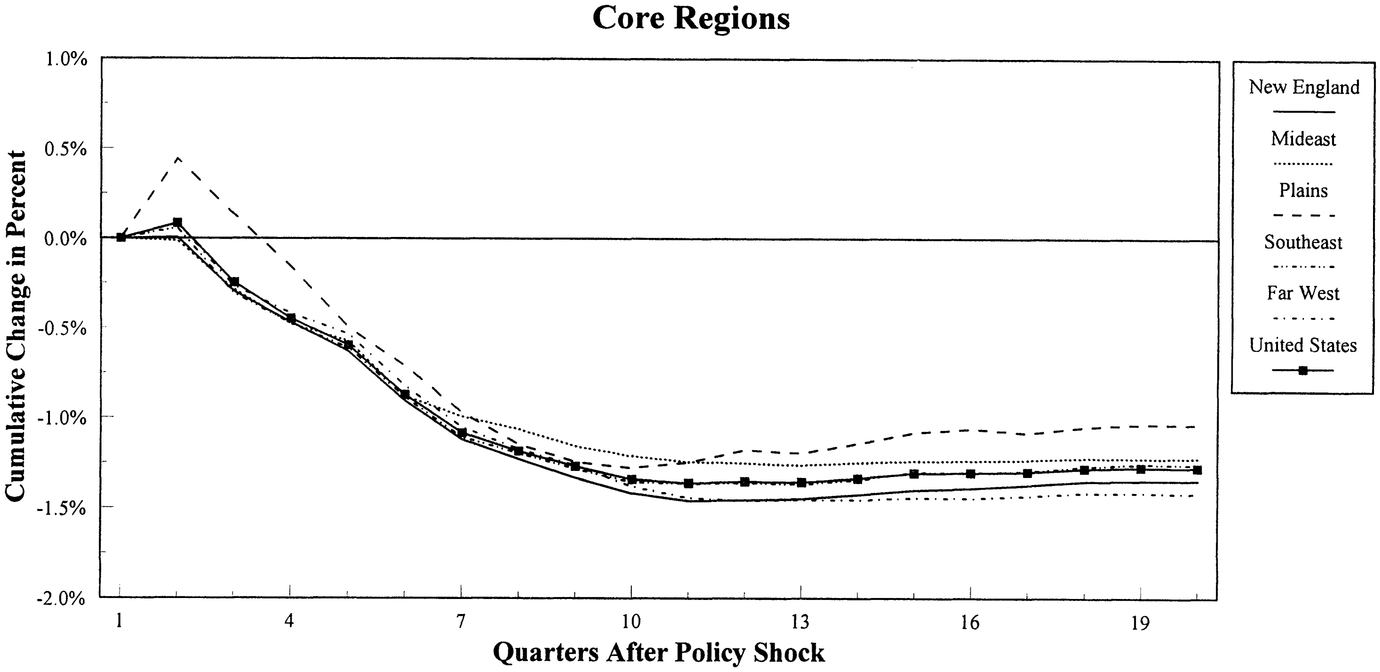



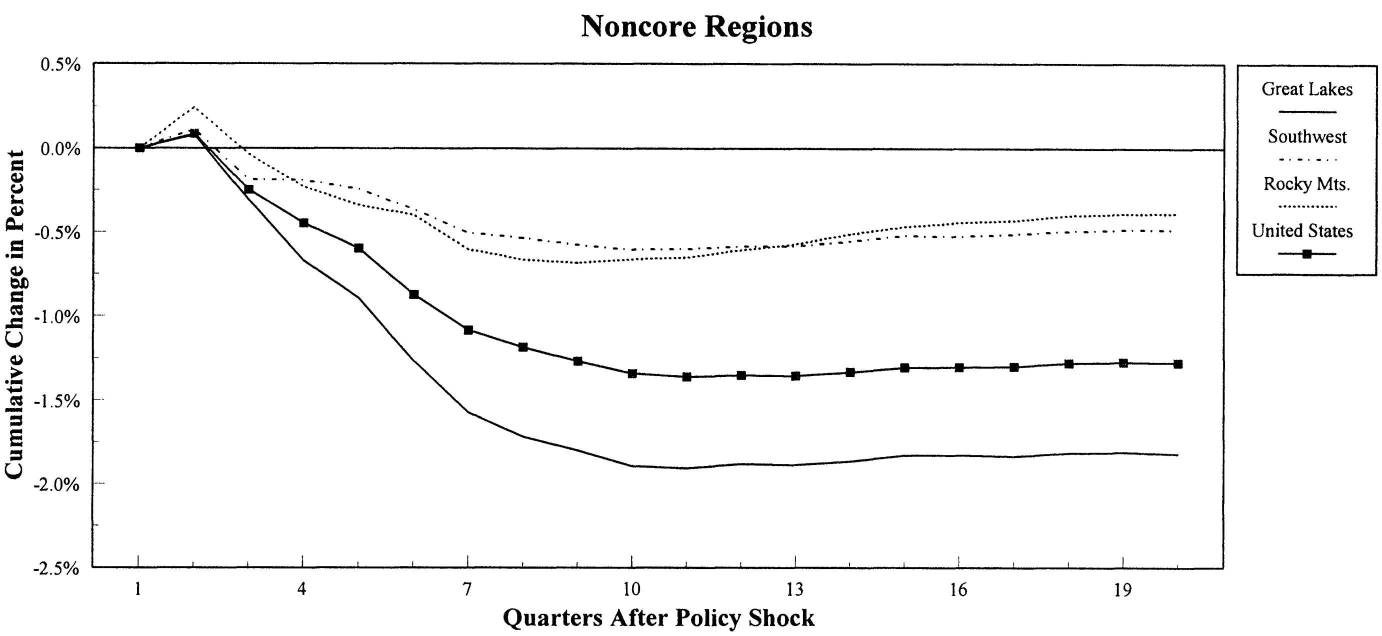

Notes: The estimated impulse response functions were computed from ten-variable VARs containing real personal income growth in each of the eight major BEA regions; growth in the producer price index for fuels and related products relative to the overall producer price index; and a monetary policy variable equal to the Boshen/Mills index. Shocks constitute one-standard-deviation unanticipated increases in the monetary policy variable.

variables were added to the baseline system—the core consumer price index (CPI) inflation rate (overall minus food and energy effects) and the first difference in the log of the index of leading indicators.[20] The former magnitude is of direct concern to the Fed and the latter proxies the array of real-side variables of interest to the Fed. Each variable was allowed to affect policy contemporaneously, but to be unaffected by other system variables contemporaneously. Figure 4 displays the cumulative impulse responses to a federal funds rate shock. As can be seen, the same core–noncore distinction arises as in the baseline, with very similar magnitudes.

[20] We thank the referee for this suggestion. The ADF unit-root tests indicated that first-differencing is required to make the log of the core CPI and the log of the index of leading indicators stationary.

FIGURE 4.—CUMULATIVE IMPULSE RESPONSE OF REAL PERSONAL INCOME TO FED FUNDS SHOCK, CORE INFLATION AND LEADING INDICATORS ADDED



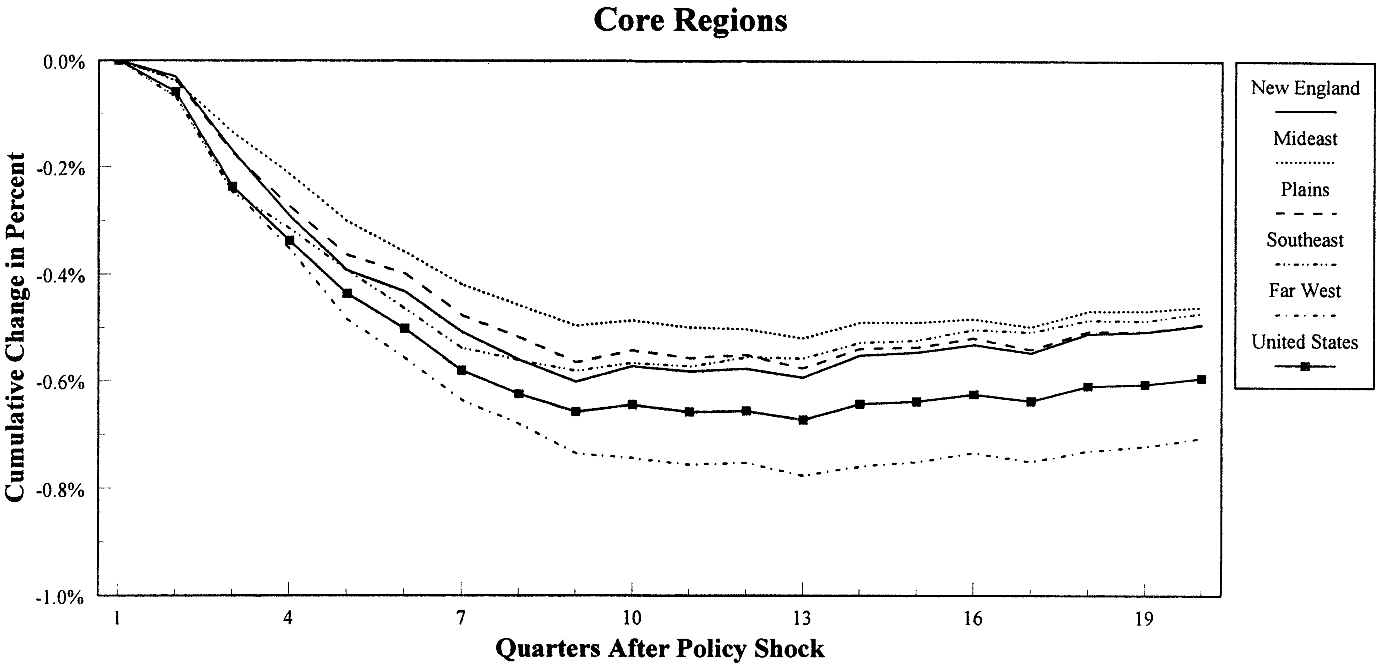

Note: The VAR is the same as the one generating the responses shown in Figure 1 except that core inflation and leading indicators are added to the model.

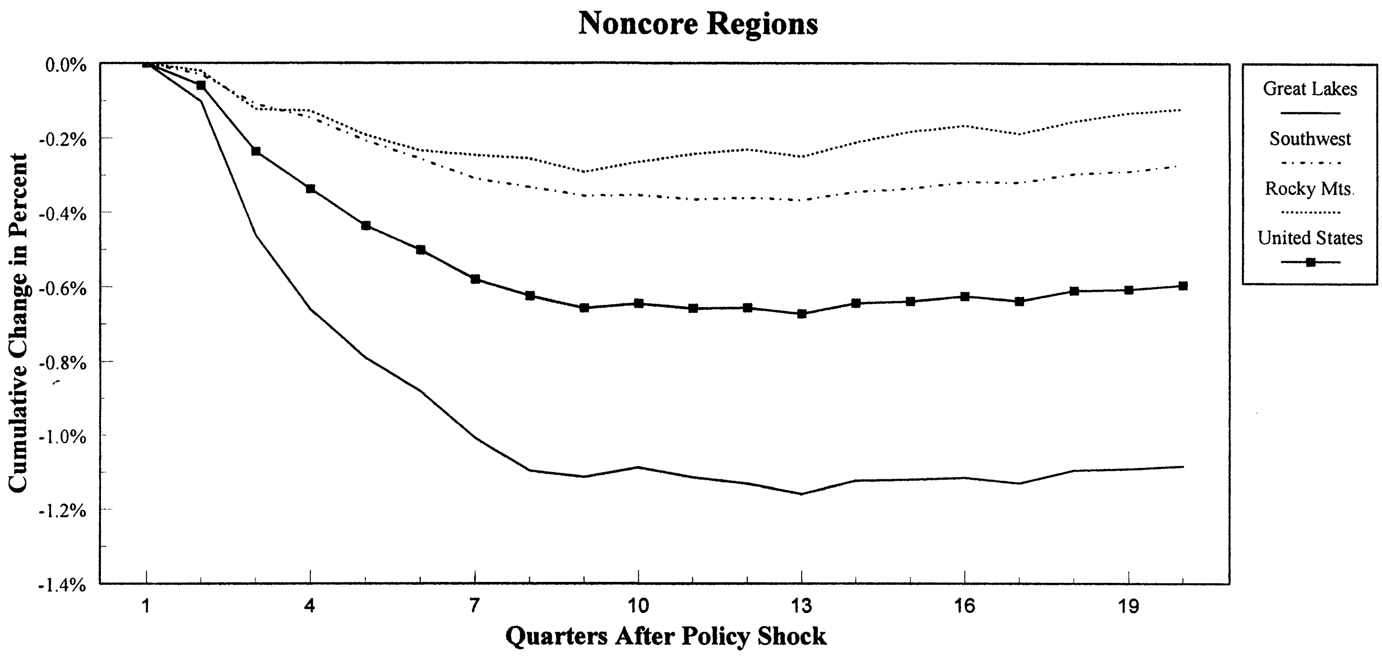
Taken together, these findings suggest that our core–noncore groupings are generally robust to the choice of a monetary policy indicator. Given the similarity of results, the rest of the analysis will rely on the earlier ten-variable baseline model, using the federal funds rate as the policy variable.

*Measures of Economic Activity:* To gauge the sensitivity of the results to our chosen measure of economic activity, we consider the effect of positive shocks to the federal funds rate change on employment growth instead of real personal income growth. Although personal income growth is a broader measure that captures growth in employment and productivity, employment data are in real terms and thus avoid issues associated with appropriately deflating income.

Figure 5 reports the cumulative impulse response functions of regional employment resulting from a one-standard-

FIGURE 5.—CUMULATIVE IMPULSE RESPONSE OF EMPLOYMENT TO FED FUNDS SHOCK, CORE AND NONCORE REGIONS




Notes: The estimated impulse response functions were computed from ten-variable VARs containing employment growth in each of the eight major BEA regions; growth in the producer price index for fuels and related products relative to the overall producer price index; and a monetary policy variable equal to the first difference of the federal funds rate. Shocks constitute one-standard-deviation unanticipated increases in the monetary policy variable.

deviation increase in the federal funds rate. The findings are generally consistent with those based on real personal income. The main difference is that the Far West region is somewhat more responsive to federal funds rate shocks than the other core regions. Nonetheless, the response of the Far West regions is closer to that of the other regions in the core grouping than it is to the Great Lakes region, which is in the noncore grouping.

*Methods for Making the Data Stationary:* A potentially important aspect of the empirical analysis is the manner in which the data are made stationary. Our analysis has relied on standard unit-root tests, which indicate that the levels of virtually all variables are nonstationary, while the first differences are stationary. Thus first-differenced data are used. There are several concerns with this approach. To begin with, unit-root tests have low power. In addition, as is now well known, first-differencing data to achieve stationarity potentially ignores valuable information about a system's long-run dynamics if the level of some variables are cointegrated. However, practical obstacles make the identification of cointegrating relationships difficult. First, sample sizes are typically too small to uncover long-run relationships. Our sample spans a 34-year period, which is likely insufficient to describe long-run relationships in the data. Moreover, as discussed in Bewley et al. (1994), small-sample problems of bias and kurtosis make the appropriate technique for estimating cointegrated systems unclear.[21] Furthermore, as noted in Johansen and Juselius (1990), interpreting estimated cointegrating vectors can be difficult when two or more cointegrating relationships are present, as could be the case in the present study.

To explore the sensitivity of the results to the method used for making the variables stationary, we follow the practical solution offered by Hamilton (1994, pp. 652–653). That is, we reestimate our system in levels without restrictions and compare the resulting dynamics to those from the differenced specification. If the results are similar, it is reasonable to conclude that our results are not influenced by the assumptions made about unit roots.[22]

Figure 6 displays the cumulative impulse responses to a one-standard-deviation positive shock to the level of the federal funds rate when levels of all variables are used to estimate the system. The findings for level data are generally quite similar to those based on first-differenced data. During the first nine to eleven quarters following a shock to the federal funds rate, the level of real personal income in the Great Lakes region responds much more than in regions composing the core grouping. Similarly, the Southwest region and the Rocky Mountains region respond much less to increases in the federal funds rate than do regions composing the core grouping. We thus conclude that the results do not depend on the method for making the data stationary.

## VI. Differential Responses in State Economic Activity

Given the robustness of the regional results, we next extended our analysis to examine how monetary policy actions differentially affect each state's economic activity. To do so, we estimated a separate, eleven-variable structural VAR for each state. Variables included are lagged values of the state's personal income growth, lagged values of personal income growth for the state's region less the state's income, lagged values of each of the other seven regions' personal income growth, the change in the log of the relative price of energy, and the change in the federal funds rate.

In an effort to conserve space, the state-level cumulative impulse responses are not presented (available on request). Three aspects of the state responses are noteworthy. First, the estimate state responses all take the same general profile found using the regional data. That is, income declines, with the maximum cumulative effect occurring eight to ten quarters following a positive fed funds rate shock. Second, not only do the estimated state responses exhibit noticeable within-region variation, there is considerable cross-region variation as well. Finally, despite these variations, the central tendency of responses for states within a given region correlates well with the earlier estimated overall response for the region. Thus, the core–noncore regional dichotomy remains.

### A. *What Caused the Differential State Responses to Monetary Policy Actions?*

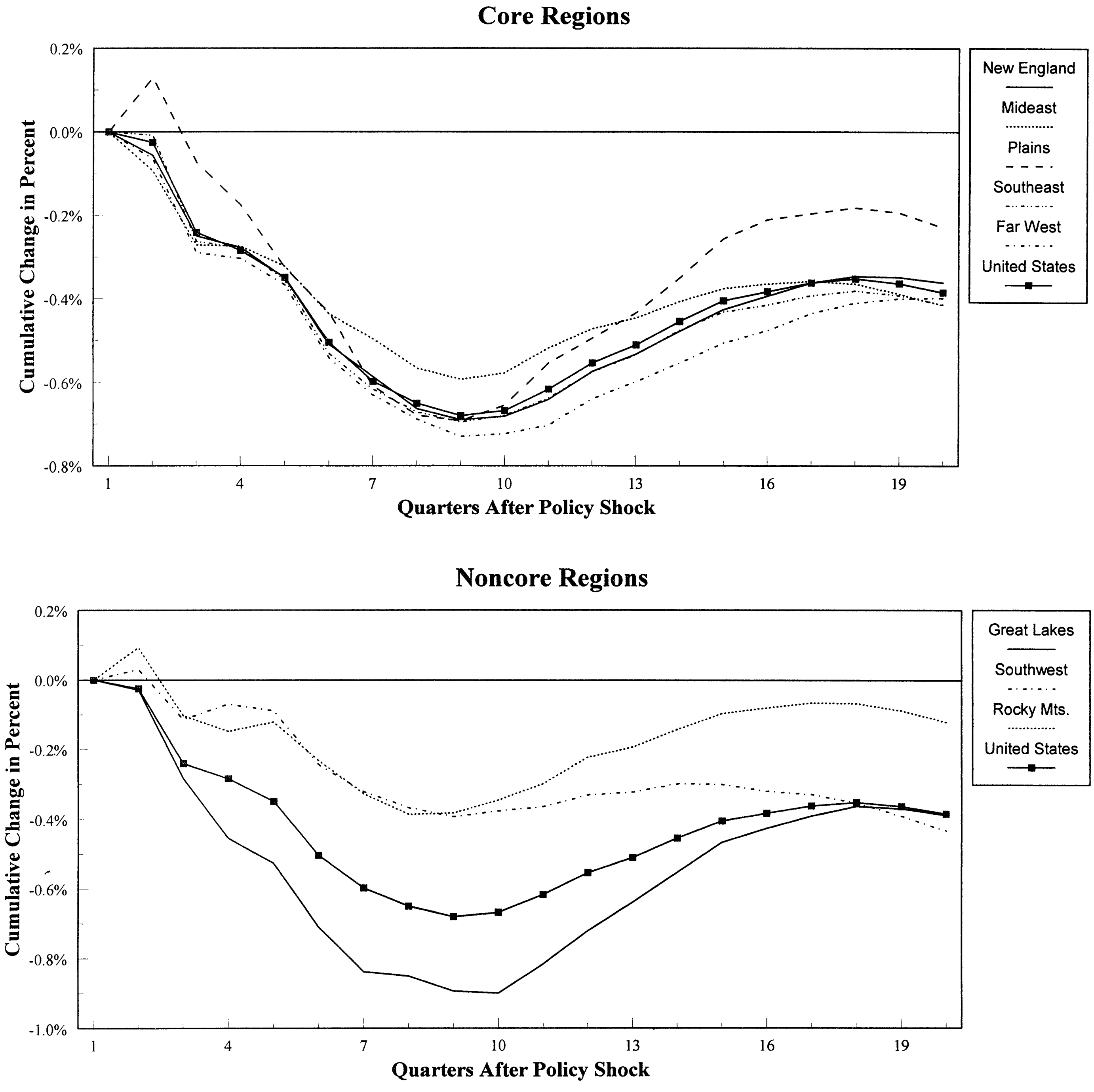
We identified three ways in which monetary policy could have differential effects in different parts of the country: area differences in the mix of interest-sensitive industries, in the mix of large and small firms, and in banks' abilities to adjust their balance sheets. How important are these factors in accounting for the different state responses to monetary policy innovations? To answer the question, we exploit the differential real income responses estimated for each state, and ask whether cross-state differences in the size of the responses are systematically related to variables capturing the hypothesized three explanatory factors. One advantage of using state-level data is that it provides 48 separate responses to monetary policy shocks instead of only eight responses when using the findings at the regional level.

The absolute values of the estimated state cumulative responses at an eight-quarter horizon are used as dependent variables in a cross-state regression equation. An eight-quarter horizon was chosen for the cumulative response because this is generally when Fed policy has its maximum cumulative impact on the level of real state personal income. The model's independent variables are designed to proxy for the hypothesized explanatory factors given to explain why states' responses to monetary policy innovations differ. A state's interest-rate sensitivity is measured by the percent of a state's GSP accounted for by manufacturing. To capture the broad credit channel, we employ the percent of a state's firms that are small, defined as the percent of a state's firms with fewer than 250 employees. To account for the bank lending channel, we use two alternative variables: the percent of a state's total loans made by a state's banks that are at or below the 90th percentile in assets nationally in 1982; and the percent of a state's total loans made by a state's banks that are at or below the 90th percentile in assets nationally in 1982 and not part of a bank holding company.

[21] Bewley et al. (1994) use Monte Carlo experiments to evaluate the small-sample properties of the Box–Tiao (1977) and Johansen (1988) estimators. See also the related paper by Gonzalo (1994).

[22] Another well-known approach to removing nonstationarity in data is the application of the Hodrick–Prescott filter, a procedure that has become influential in the real business cycle literature. While straightforward to employ, Cogley and Nason (1995) present evidence that the filter can generate cyclical dynamics even when none are present in the origin data. Consequently, we chose not to use this procedure.

FIGURE 6.—CUMULATIVE IMPULSE RESPONSE OF REAL PERSONAL INCOME LEVEL TO FED FUNDS SHOCK, CORE AND NONCORE REGIONS




Notes: The estimated impulse response functions were computed from ten-variable VARs containing the level of real personal income in each of the eight major BEA regions; level in the producer price index for fuels and related products relative to the overall producer price index; and a monetary policy variable equal to the federal funds rate. Shocks constitute one-standard-deviation unanticipated increases in the monetary policy variable.

Because the estimated impulse responses represent average behavior during the sample period, averaging the data for the explanatory variables is appropriate. Data availability limited averaging to the period from the mid-1970s to the early 1990s. Averaging also minimizes the chance that the results depend on the data for a particular year and helps control for business-cycle dynamics.

Four estimated regressions are presented in table 6. Equations (1) and (2) contain the three explanatory variables described. The banking variable in equation (1) is measured using all small banks, while the analogous variable in equation (2) excludes banks that are members of a holding company. Equations (3) and (4) are similar to equations (1) and (2), respectively, except that dummy variables identify-

TABLE 6.—ESTIMATED EQUATIONS EXPLAINING CROSS-STATE VARIATION IN POLICY RESPONSES[a]

| Variable | Equation 1 | Equation 2 | Equation 3 | Equation 4 |
|---|---|---|---|---|
| Intercept | −0.3567 | −0.3794 | −0.0732 | −0.0759 |
| | (0.7018) | (0.6784) | (0.6276) | (0.6076) |
| Percent Manufacturing[b] | 0.0199 | 0.0206 | 0.01124 | 0.0124 |
| | (0.0064)*** | (0.0064)*** | (0.0066)** | (0.0065)** |
| Percent Small Firms[b] | 0.0122 | 0.0125 | 0.0124 | 0.0126 |
| | (0.0089)* | (0.0085)* | (0.0080)* | (0.0077)* |
| Percent Small Bank Loans (all banks)[b] | −0.0024 | | −0.0046 | |
| | (0.0020) | | (0.0022)** | |
| Percent Small Bank Loans (no holding co.)[b] | | −0.0038 | | −0.0074 |
| | | (0.0027) | | (0.0029)*** |
| New England[c] | | | −0.1033 | −0.1270 |
| | | | (0.1116) | (0.1103) |
| Mideast[c] | | | −0.3382 | −0.3635 |
| | | | (0.1276)** | (0.1257)*** |
| Great Lakes[c] | | | 0.2244 | 0.1908 |
| | | | (0.1224)* | (0.1201) |
| Plains[c] | | | 0.0568 | 0.0554 |
| | | | (0.1142) | (0.1100) |
| Southwest[c] | | | −0.3009 | −0.2966 |
| | | | (0.1365)** | (0.1328)** |
| Rocky Mountains[c] | | | −0.1691 | −0.2255 |
| | | | (0.1348) | (0.1348) |
| Far West[c] | | | 0.0683 | 0.0559 |
| | | | (0.1457) | (0.1403) |
| Adjusted $R^2$ | 0.1586 | 0.1729 | 0.3821 | 0.4115 |

Notes: [a] Standard errors in parentheses. *, **, and *** indicates that a null hypothesis of zero is rejected at the 10%, 5%, and 1% levels, respectively.
[b] Null hypothesis is tested against an alternative hypothesis of a theoretically prescribed positive coefficient (one-tailed test).
[c] Null hypothesis is tested against an alternative hypothesis of a non-zero coefficient since an expected sign is unspecified by theory (two-tailed test).

ing the region in which a state is located have been included to control for region-specific factors (Southeast region is excluded).

Each regression is significant at the 1% level, explaining between 15% and 41% of the cross-state variation in cumulative responses. Because the impulse responses are measured with sampling error, particularly high $R^2$s are not expected in regressions using the estimated cumulative impulses as dependent variables. The significance of each variable can be gauged from two different perspectives. Because the theories that motivate the independent variables unambiguously predict positive coefficients, a one-tailed test can be used to test hypotheses. Alternatively, one can simply ask whether these variables provide significant explanatory power in the equation, regardless of their predicted sign. In the latter case, a two-tailed test is used. The results reveal that the percent of a state's GSP accounted for by manufacturing is positively and significantly related to the size of a state's response to federal policy shocks, whether a one-tailed or a two-tailed test is used. The result is robust to the choice of the loan variable and to the inclusion of regional dummies. Studies have shown that spending on manufactured goods, especially durable goods, tends to be interest sensitive.[23] Spending on services, in contrast, tends to vary little with interest rates. Our finding suggests that differences in interest rate sensitivities are one reason for different state responses.[24]

A one-tailed test reveals that states containing a large concentration of small firms tend to be more responsive to monetary policy shifts than states containing small concentrations of small firms.[25] While the firm-size variable is significant at only the 10% level, the finding is robust to the choice of loan variable and to the inclusion of regional dummies. The estimated coefficient is, however, insignificantly different from zero, when a two-tailed test is used. Thus our findings lend, at best, weak support to the credit view of monetary policy.

Finally, we find that a region becomes less sensitive to a monetary policy shock as the percent of small banks in the region increases. The estimated coefficients are negative in all four equations and negative and significant in the equations adding the regional dummy variables. The finding of a negative sign on the small bank variable is inconsistent with the view espoused by Kashyap and Stein (1995).[26] One possibility for the inconsistency is that a bank's asset size may be a poor indicator of its ability to adjust its balance sheet to monetary policy actions. For example, Peek and Rosengren (1995) suggest that bank capital is a better indicator—better capitalized banks, have more and cheaper alternative sources of funds available. In addition, Kashyap and Stein (1995) point out that regional differences in the types of loans being made might also matter, a factor not controlled for in our study.

## VII. Conclusions

This paper uses time-series techniques to examine whether monetary policy had symmetric effects across regions in the United States during the 1958–1992 period. Impulse response functions from an estimated structural VAR reveal a

[23] See Bennett (1990) for a survey of relevant studies.

[24] At the suggestion of a referee, we reestimated the baseline system using regional manufacturing employment instead of real personal income or total employment. The logic is that, if differential policy responses are driven mainly or only by manufacturing, the policy responses of manufacturing employment across regions should be very similar. However, the resulting impulse responses continued to exhibit differences, an outcome consistent with our results that something in addition to the percent of manufacturing induces differential policy responses.

[25] At first glance, the estimated positive association between concentration of small firms and the size of the policy response appears to be inconsistent with the negative correlation implied by table 2. The regression results are consistent, however, because the state concentration of manufacturing firms is negatively correlated with the state concentration of small firms (simple correlation = −0.62). Thus the negative simple correlation mainly reflects the strong association with manufacturing. In the regression, the effect of manufacturing is held constant and the positive correlation between the concentration of small firms and size of policy response is uncovered.

[26] If small banks largely make loans to small firms, this relationship would be captured by the small firm variable. There is a moderate correlation between the small firm variable and the small bank variable used in the regression (a simple correlation of 0.5). This correlation could explain the lack of a positive response of the bank size variable to changes in monetary policy, but it does account for the estimated negative effect.

core of regions—New England, Mideast, Plains, Southeast, and the Far West—that respond to monetary policy changes in ways that closely approximate the U.S. average response. The core regions accounted for a little more than two-thirds of the aggregate 1980 GSP in the United States and for 70% of the total U.S. population. The study also identified three noncore regions, one (Great Lakes) that is noticeably more sensitive to monetary policy changes, and two (Southwest and Rocky Mountains) that are much less sensitive. State-level VARs were also estimated. State-level impulse responses to monetary policy shocks displayed noticeable variation both within and across major BEA regions. Despite these variations, the central tendency of state responses for each region correlates well with the estimated overall regional responses. Thus, the core/noncore regional dichotomy remains.

Concerning the reasons for the differential state responses, we find that the manufacturing-intensive states are more responsive to changes in monetary policy shocks than the more industrially diverse states (evidence for a regional interest-rate channel). Evidence for a broad credit channel is, however, weak. Although we find that states containing a relatively larger concentration of small firms tend to be more responsive to monetary policy shifts than states composed of smaller concentrations of small firms, the correlation is only marginally significant. In addition, our findings are inconsistent with the credit channel described by Kashyap and Stein (1995).

The existence of disparate responses underscores the difficulty of conducting a national monetary policy for an area as large and diverse as the United States and raises issues of cross-regional equity. In a related vein, the findings reported in this paper raise a cautionary note for regional economic compacts, such as the European Monetary Union, that would rely on a single central bank. The economies of these countries are more heterogeneous than are the regional economies that comprise the United States (Bayoumi and Eichengreen (1993) and Carlino and DeFina (1998)). In such cases, likely regional differences in policy responses will arise across sovereign nations. These differences will make it difficult to form and maintain a monetary union.

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## APPENDIX A

**Definitions of Regions**

*New England*
Connecticut
Maine
Massachusetts
New Hampshire
Rhode Island
Vermont

*Mideast*
Delaware
Maryland
New Jersey
New York
Pennsylvania

*Great Lakes*
Illinois
Indiana
Michigan
Ohio
Wisconsin

*Plains*
Iowa
Kansas
Minnesota
Missouri
Nebraska
North Dakota
South Dakota

*Southeast*
Alabama
Arkansas
Florida
Georgia
Kentucky
Louisiana
Mississippi
North Carolina
South Carolina
Tennessee
Virginia
West Virginia

*Southwest*
Arizona
New Mexico
Oklahoma
Texas

*Rocky Mountains*
Colorado
Idaho
Montana
Utah
Wyoming

*Far West*
California
Nevada
Oregon
Washington