

Monetary Policy and Regional Price and Wage Dispersion in the U.S.

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Abstract: It is generally accepted that certain fiscal policies have strong local impacts. However, it is considerably less obvious that national fiscal and monetary policies will have significantly different impacts on areas as large as states or census regions. This paper focuses on the impact of U.S. monetary policy on relative inflation and wage growth rates for four large regions within the U.S. The results suggest that the conditional variance of aggregate monetary policy shocks, estimated from an ARCH process, increases the dispersion of regional inflation and regional real wage growth rates for the four U.S. census regions.

I. INTRODUCTION

Macroeconomic theory and empirical analysis imply that national fiscal and monetary policies have important impacts on national policy outcomes like output growth, inflation, and wage growth. The macroeconomic literature has paid less attention to the possibility that aggregate fiscal and monetary policies may have important regional impacts. These regional impacts should concern policy makers most if they are asymmetric and thus differ significantly across regions. For instance, a fiscal or monetary policy that had the same effect on all areas would impose no unintended burdens on regions within the national economy, as all would share equally in the pain/gain imposed by the federal authority. However, an aggregate federal policy that created significantly different outcomes across regions would be problematic, with some regions experiencing more of the costs/benefits of the aggregate policy than other regions.

The possibilities for asymmetric impacts are easy to imagine in the case of fiscal policies, particularly those aimed directly at local economies. For instance, the decision to close a particular military base will obviously affect the local economy more than the national average. The possibility that aggregate monetary policy may have asymmetric impacts within a nation is somewhat less easy to imagine, but also has important implications for policy makers. It highlights an unintended consequence of national monetary policy, which is aimed at the macroeconomy but which may distort regional outcomes by affecting economic performance in one region more than in another.

Distortions of regional outcomes by aggregate monetary policy may impose a variety of real costs, including the dispersion of per capita income and jobs, as well as relative inflation and wage growth rates. These distortions are also likely to lead to costly and inefficient labor migration from temporarily depressed regions to temporarily growing regions.

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This paper focuses on regional differences in inflation and wage growth rates within the U.S., in part because these indicators have been ignored by previous work. However, regional differences in inflation and wage growth performance have important implications for the health of a regional economy because they influence overall economic efficiency, regional costs of living, effective tax burdens, and returns to work.

Asymmetric changes in inflation rates imply regional differences in well-known aggregate inflation costs. Recent work on the costs of inflation suggests that even moderate rates of inflation may have significant welfare costs. Feldstein (1983, 1997) emphasizes the interaction of inflation and the U.S. tax code as an important determinant of the economic cost of inflation. He argues that inflation exacerbates the impact of the distortionary tax treatment of capital income. Thus, inflation interacts with capital income taxes to distort returns to saving. This, in turn, impacts intertemporal consumption incentives and the incentives to invest in owner-occupied housing. Feldstein (1997) estimates that the annual gain resulting from reducing the U.S. inflation rate from 2 percent to zero would range from 0.76 to 1.04 percent of U.S. gross domestic product (GDP). To the extent that regional inflation rates are dispersed across regions, the costs of aggregate U.S. inflation will be unevenly distributed across regions.

The purpose of this paper is to explore the relationship between aggregate monetary policy and regional inflation and wage growth rates. The plan of this paper is as follows: Section II reviews the literature on the regional impacts of national monetary policy; Section III describes the inflation and wage growth data sets; Section IV describes the measures of aggregate variability; Section V presents results relating aggregate variability and the dispersion of U.S. regional inflation and wage growth rates; and the paper concludes with Section VI.

II. PREVIOUS INVESTIGATIONS OF THE REGIONAL IMPACT OF MONETARY POLICY

The importance of examining the asymmetric impacts of national monetary policy on regional economies in the U.S. has received some attention by researchers. Purely theoretical work by Dow (1987) and Moore and Hill (1982) has explored the regional money supply process and the influence of national monetary policy on regional monetary aggregates and interest rates.

Empirical work has examined the extent to which regional money markets exist within the U.S., with Miller (1978), Bias (1992), and Roberts and Fishkind (1979) concluding that there is significant regional segmentation of money markets. Their work concludes that both regional economic and national monetary forces influence regional money markets and regional economic activity.

Another line of investigation abstracted from regional money markets and instead relied on reduced-form regressions of national monetary variables on regional measures of economic performance, such as real and nominal personal income and jobs.

Beare (1976) is an early example of this type of study, which investigated national monetary influences on Canadian provinces. Garrison and Chang (1979), Mathur and Stein (1980), and Garrison and Kort (1983) follow the same basic approach using U.S. data and reach the conclusion that national monetary policies influence regional economies to varying degrees.¹ More recently, Carlino and DeFina (1995, 1998) and Kozlowski (1995) bring more sophisticated econometric procedures to bear and come to the same conclusion: national monetary policies have significant asymmetric impacts on regional economic performance in terms of personal income and job growth.

The studies by Carlino and DeFina (1995, 1998) use a structural vector autoregression framework and show that the Great Lakes region seems to respond the most to national monetary shocks while the Southwest and Rocky Mountain regions respond the least.² Their analysis goes on to assess channels of transmission from national monetary shocks to regional personal income. Echoing earlier work (especially Fishkind 1977) they postulate that industry mix may play an important role in transmitting national monetary shocks to regional economies. The idea is that regions differ in their industry mix, so that some regions specialize in interest-sensitive industries (construction and durable manufacturing, for example). Thus, a national monetary policy that influences interest rates will have a larger impact on interest-sensitive industries and thus also a larger impact on regions that specialize in these industries. Their results suggest that this is the main channel of transmission of national monetary shocks to regional economies.

Another possible channel for the transmission of monetary shocks to regional economies is the credit channel. The idea here is that larger firms have more access to national financial markets than do smaller firms, which must rely on local financing.³ Thus, the health of the local financial sector becomes a crucial factor in determining regional growth, particularly the growth of smaller businesses. Samolyk (1994) finds evidence supporting this version of the regional credit view, while Carlino and DeFina (1995, 1998) find weak evidence of a regional credit channel using the eight Bureau of Economic Analysis regions and little evidence using state data.⁴

The empirical work to this point suggests that national monetary policy disperses income and job growth across U.S. regions. This study introduces a new consideration into the literature by considering the extent to which national monetary policy produces asymmetric regional price and wage responses across regions within the U.S.

In examining the impact of national monetary policy on regional prices and wages, we follow the literature relating intersectoral relative price variance (ISPV) to national monetary policy. It is a well-established phenomenon in the

¹Mathur and Stein (1980) take issue with the reduced-form approach in general, arguing that the estimated coefficients are not stable over time.

²Their regions are defined using the Bureau of Economic Analysis definition.

³See Bernanke (1993) for a more detailed treatment.

⁴Carlino and DeFina also find little evidence supporting a version of the regional credit channel based on regional differences in bank size.

monetary literature that an increase in the variability of monetary policy contributes significantly to an increase in the variability of sectoral prices around the national average.⁵ Thus, national monetary policy may interfere with the efficiency of the economy by creating disturbances in relative prices across industries. Likewise, national monetary policy may also create disturbances in relative prices across regions.

This opens up an additional regional channel of transmission of national monetary policy to regional economies. As we will argue below, regional economies differ in their industrial composition, and thus monetary policy that disperses relative prices across industries may also disperse relative prices across regions.

III. REGIONAL RELATIVE PRICES AND WAGES FOR THE FOUR CENSUS REGIONS

The empirical analysis to follow uses seasonally adjusted price and wage data for the four census regions of the U.S. Prices are measured by quarterly consumer price indexes for urban wage earners and clerical workers (CPIW) for all items. Wages are measured by average hourly earnings of production workers in manufacturing.

Consumer price index data is available for the U.S., the four census regions, and for selected metropolitan statistical areas (MSAs). Data is collected by the Bureau of Labor Statistics (BLS) through a representative sample of products and establishments in the relevant MSA.

Data on prices are collected for a representative market basket of goods and services from 85 urban areas and 21,000 retail and service outlets within the U.S. Individual prices are combined into an all-item average for each region using weights derived from the share of spending on each good in total spending (these shares are estimated from data gathered in the Consumer Expenditure Survey). BLS compiles indexes for the U.S. (monthly), the four census regions (monthly), and 29 local MSAs (monthly, bimonthly, or semiannually, depending on sample size). Note that the U.S. average and census region data reflect prices gathered from 85 urban areas, even though only 29 local area indexes are separately published.

In the empirical analysis to follow, we will focus on the data for the four census regions. We do so because these regions give us the longest regional time series and also because they encompass both published data for the large MSAs and unpublished data for the smaller MSAs. This makes the census region data more reliable than the local area indexes, which rely on much smaller samples. The tradeoff imposed by this choice is that we cannot make specific statements about the impact of national monetary policy on specific MSAs (or states). However, the focus of this paper is on the power of national monetary policy to disperse regional inflation rates from the average.

⁵Glejer (1965), Vining and Elwertowski (1976), and Fischer (1981, 1982) found that increases in monetary variability increased the variability of prices across industries. Using a Lucas (1973) type framework, Hercowitz (1981, 1982), among others, for the U.S., Neumann and von Hagen (1990) for Germany, and Blejer and Leiderman (1980) for Mexico generally found similar results.

Average hourly earnings of production workers in manufacturing is published monthly for all states, also by BLS. This data is collected by BLS as part of the Current Employment Survey (which is also sometimes referred to as BLS 790). Average hourly earnings are calculated by dividing gross payrolls by total hours. Gross payrolls include overtime pay but exclude the value of employee benefits and are calculated before deductions for payroll taxes. In order to maintain comparability with the CPIW data, the average hourly earnings in manufacturing data is aggregated to the census region level and we focus on relative wage growth rates.

Regional Economic Characteristics

Table 1 summarizes the relative size and industrial characteristics of the four census regions. Industry shares of gross state product (GSP) are calculated by averaging industry GSP during the 1977-94 period and dividing by the average total GSP in the region during the same period. As the table shows, industry shares of total GSP varied significantly across regions during the period.⁶ For instance, census region mining shares vary from 0.3 percent in the Northeast to 5.4 percent in the South. We can summarize the regional dispersion of industry shares by calculating a coefficient of variation for each industry. The coefficient of variation normalizes the standard deviation of region industry shares around the U.S. average by dividing it by the U.S. industry share and multiplying by 100. Again using the mining industry as an example, the standard deviation of mining GSP shares was 2.26 and the U.S. industry share was 2.7, which results in a coefficient of variation for the mining industry of 83.7.

The largest coefficients of variation arise in agriculture, forestry, and fishing; mining; and manufacturing (particularly durable goods manufacturing). Manufacturing as a share of total GSP varies from a high of 26 percent in the Midwest to a low of 14.9 percent in the West. The dispersion is even greater in the durable manufacturing sector, with GSP shares ranging from a high of 17.1 percent in the Midwest to a low of 8.4 percent in the South.

Service-producing industries generally record lower levels of dispersion, with the exception of finance, insurance, and real estate. This sector exhibits notable concentrations in the Northeast and West, with less activity in the South and Midwest regions.

Average population shares during the 1977-94 period range from 34.0 percent in the South census region to 20.2 percent in the West census region. The Northeast registered an average population share of 21.0 percent while the Midwest accounted for 24.8 percent of U.S. population. Overall, the four census regions are similar in population size, but differ in important ways in terms of industrial specialization.

⁶GSP data is currently available for the 1977-94 period. For the sake of comparability, we will calculate population shares for the same period.

TABLE 1
Regional Population Shares and Industry Shares of Regional Gross State Product (GSP)
(Percent)

	Northeast	Midwest	South	West ^a	U.S.	Coefficient of Variation ^b
Region Shares of U.S. Population Average: 1977-94						
Population	21.0	24.8	34.0	20.2	—	—
Industry Shares of Region GSP Average: 1977-94						
Ag., For., Fish.	0.8	3.1	2.1	2.5	2.1	34.4
Mining	0.3	0.9	5.4	3.0	2.7	83.7
Construction	3.9	4.0	4.7	4.9	4.4	9.5
Manufacturing	19.9	26.0	18.5	14.9	19.8	20.5
Durables	11.5	17.1	8.4	9.9	11.5	32.3
Nondurables	8.5	8.9	10.1	5.0	8.3	20.2
TCPU ^c	8.7	9.1	9.4	8.4	8.9	4.4
Wholesale Trade	7.3	7.0	6.6	6.5	6.8	4.7
Retail Trade	8.4	9.2	9.4	9.7	9.2	4.9
FIRE ^d	21.1	14.7	14.4	18.4	16.9	17.0
Services	18.6	14.9	14.8	17.7	16.3	10.4
Government	11.0	11.1	14.8	13.9	12.9	13.2

^aCensus region definitions. **Northeast:** Connecticut, Maine, Massachusetts, New Hampshire, New Jersey, New York, Pennsylvania, Rhode Island, Vermont. **Midwest:** Illinois, Indiana, Iowa, Kansas, Michigan, Ohio, Wisconsin, Minnesota, Missouri, Nebraska, North Dakota, South Dakota. **South:** Alabama, Arkansas, Delaware, District of Columbia, Florida, Georgia, Kentucky, Louisiana, Maryland, Mississippi, North Carolina, Oklahoma, Tennessee, South Carolina, Texas, Virginia, West Virginia. **West:** Alaska, Arizona, California, Colorado, Hawaii, Idaho, Montana, Nevada, New Mexico, Oregon, Utah, Washington, Wyoming.

^bCoefficient of variation is the standard deviation around the U.S. average divided by the U.S. share, multiplied by 100.

^cTCPU is Transportation, Communications, and Public Utilities.

^dFIRE is Finance, Insurance, and Real Estate.

Sources: GSP and Population data: Regional Economic Information Service CD, 1997, Bureau of Economic Analysis.

Regional Relative Inflation and Wage Growth

Since we are interested in regional differences from the national average, we focus on regional inflation and wage growth rates minus the comparable national average. Table 2 shows the standard deviations and means, as well as the 10th and 90th percentiles, of the resulting relative inflation and relative nominal and real wage growth rates. All rates are expressed as compound annual rates of growth. The standard deviations of our relative inflation rates tell us the extent to which regional inflation rates have deviated from the national average during the period. For the four census regions, standard deviations range from 0.8 (South) to 1.5 (West) percentage points. To put this in perspective, note that the average U.S. inflation rate over the period was 5.5 percent. This implies that the standard deviation of regional inflation rates ranges between 15 and 27 percent of the average U.S. inflation rate. The 10th percentile ranges from -0.8 to -1.6 percent and the 90th percentile varies from 0.9 to 2.1 percent. On average over the nearly 30-year period, regional inflation rates range from 0.1 percent above the national rate to 0.1 percent below the national rate. Regional inflation rates are close to the U.S. infla-

tion rate on average over long periods of time, but deviate substantially from the national average during shorter periods.

TABLE 2
Descriptive Statistics
Relative Inflation and Wage Growth Series
Quarterly Data at Compound Annual Rates (Percent)
(Quarterly, 1967Q2 - 1995Q4)

	Northeast	Midwest	South	West
Relative Inflation				
Std. Deviation	1.3	1.1	0.8	1.5
Mean	0.1	-0.1	0.1	0.0
10th Percentile	-1.4	-1.5	-0.8	-1.6
90th Percentile	1.6	1.1	0.9	2.1
Relative Nominal Wage Growth				
Std. Deviation	1.2	1.5	1.5	1.8
Mean	0.1	0.0	0.2	-0.4
10th Percentile	-1.5	-1.8	-1.4	-2.7
90th Percentile	1.7	1.9	2.2	1.8
Relative Real Wage Growth				
Std. Deviation	1.6	2.0	1.7	2.3
Mean	0.1	0.2	0.2	-0.4
10th Percentile	-1.8	-2.1	-2.1	-3.4
90th Percentile	1.9	2.7	2.6	2.5

Source: Calculations by author.

The standard deviation for relative nominal wage growth ranges from 1.2 to 1.8 percent, with the 10th (90th) percentile varying from -1.5 to -2.7 percent (1.7 to 2.2 percent). Standard deviations of relative nominal wages range from 22 to 33 percent of U.S. average nominal wage growth.

The standard deviation of relative real wages varies between 1.6 and 2.3 percent, with the 10th (90th) percentile varying from -1.8 to -3.4 percent (1.9 to 2.7 percent). These standard deviations are quite large relative to average U.S. real wage growth, which is close to zero during this period. Thus, nominal and real wage growth patterns are similar to those found for inflation rates.

An additional consideration is the persistence of shocks to regional relative prices and wages. If shocks to relative price levels are very persistent, then the arbitrage processes across regions work only slowly to eliminate price level differentials.

Results of the Augmented Dickey-Fuller tests on regional prices and wages suggest that the series exhibit substantial persistence in their levels but not in their growth rates.⁷

Strong persistence in relative price and wage levels implies that shocks to relative inflation and wage growth rates produce changes in relative price and wage levels that remain for a long time. This is consistent with results obtained by Blanchard and Katz (1992), who examine the adjustment of state economies in the U.S. to relative economic shocks. They find that relative state employment and wage levels exhibit strong persistence, while employment and wage growth rates are less dependent on past shocks. Carlino and DeFina (1995, 1998) also find evi-

dence of strong persistence in personal income levels at the state and region level. Evidence regarding the persistence of shocks to important regional indicators is important because it highlights the point that national policies that impact regional growth rates will have long-term impacts on the level of economic activity across regions.

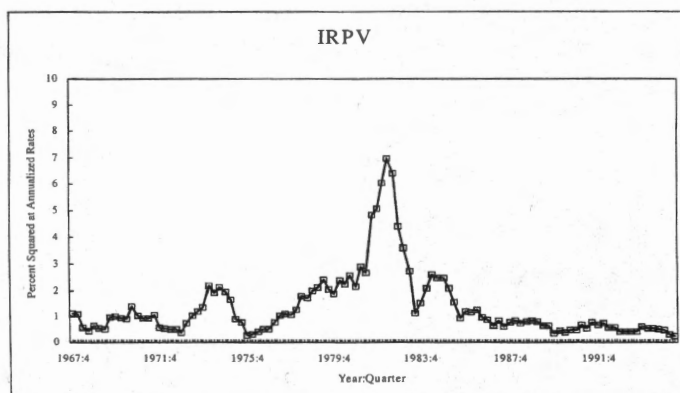
Using our relative regional inflation rates, we can parsimoniously depict average deviations from national performance by constructing a regional analog to the intersectoral relative price variance measure used in the monetary literature. Thus, we construct an index of regional relative price variance measured as the weighted average of the squared relative inflation rates, as shown in Equation (1).⁸

$$(1) \quad \text{IRPV}_t = \sum w_{i,t} (\pi_{i,t} - \pi_{us,t})^2.$$

Here, $\pi_{i,t}$ is the inflation rate in region i at time t , $\pi_{us,t}$ is the U.S. city average inflation rate, and $w_{i,t}$ is the population weight in region i .⁹ Similarly, the measure of relative nominal wage variance (REV) is calculated by summing the population weighted, squared, relative nominal wage growth rates. Finally, relative real wage variance (RREV) is calculated as the population weighted average of squared, relative real wage growth rates.

The indexes of IRPV, REV, and RREV are plotted in Figures 1-3. As is evident from the figures, there is substantial variation over time. Increases seem to be generally associated with national recessions, as the indexes rise during the 1969-72, 1973-75, and 1979-83 periods. For the 1967-95 period, the mean of IRPV is 1.3 percent, which implies that the average regional deviation from the U.S. inflation rate is 1.14 percentage points. This percentage point deviation amounts to 21 percent of the average U.S. inflation rate.

Figure 1



⁸Results of Augmented Dickey-Fuller tests available from the author upon request.

⁹BLS uses population weights to aggregate regional price indexes. See BLS Handbook of Methods, April 1997, <www.bls.gov/homtoc.htm>. For the sake of comparability, we will use population weights to calculate nominal and real average hourly earnings in manufacturing dispersion measures.

^{*}Population weights are derived from census population estimates for the years 1960, 1970, 1980, and 1990. Data for years between census counts are derived by interpolation.

Figure 2

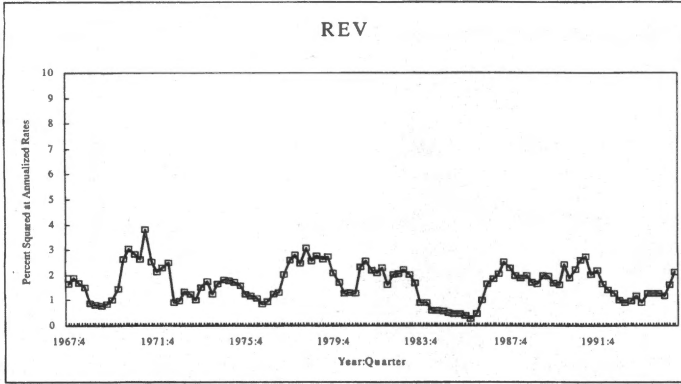
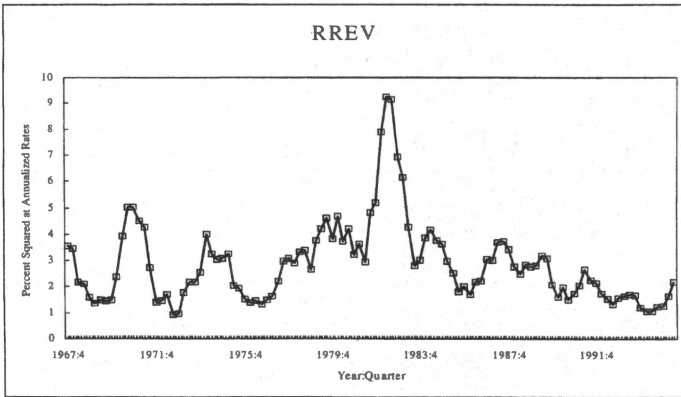


Figure 3



REV averages 1.7 percent during the 1967-95 period and the average deviation of regional wage growth rates from the U.S. rate amounts to 24 percent of average U.S. wage growth. RREV records a mean of 2.9 percent. The average deviation of regional real wage growth rates is quite large relative to the U.S. average growth rate because the U.S. growth rate is close to zero. Overall, the regional variation of relative inflation and wage growth rates is substantial and varies over time in an interesting way.

IV. MEASURING AGGREGATE VARIABILITY

If monetary policy shocks have asymmetric effects on regional inflation rates, then the variance of regional inflation rates around the U.S. average will depend on the variance of aggregate monetary shocks. To see that the relationship needs to be modeled in terms of variances, suppose that regional inflation rates are in fact affected differently by monetary policy shocks. Then we can write:

$$(2) \quad \pi_{i,t} = \gamma_i \text{MPShock}_t,$$

where i indexes regions, $\pi_{i,t}$ is the regional inflation rate, γ_i is the (asymmetric) regional response to the aggregate monetary policy shock, MPShock. Aggregating up to the national level gives:

$$(3) \quad \pi_{us,t} = \bar{\gamma} \text{MPShock}_{it}$$

Using Equations (2) and (3) and the definition of IRPV gives:

$$(4) \quad \text{IRPV}_t = \sum w_i (\gamma_i - \bar{\gamma})^2 \text{MPShock}_{it}^2$$

where w_i is a regional weight. Assuming MPShock_{it} to be mean zero leads to our interpretation of MPShock_{it}^2 as the variance of monetary policy shocks. This is the channel we will explore in the remainder of the paper.

Identifying ARCH Variance Processes

To carry out our test, we need a measure of the variance of monetary policy shocks. We use Engle's (1982) Autoregressive Conditional Heteroscedasticity (ARCH) technique for this purpose. The ARCH model allows the conditional variance of monetary policy shocks to change over time and provides a way to efficiently estimate a parametric model of a time-varying variance process. Intuitively, the ARCH model implies that large forecast errors tend to cluster together, as do small forecast errors. It is common to interpret a clustering of large forecast errors as a period of increased uncertainty.

To identify empirically a series of monetary policy shocks, we follow Bernanke and Blinder (1992), Gordon and Leeper (1994), and Carlino and DeFina (1995, 1998), among others, and focus our attention on the market for reserves. In particular, Bernanke and Blinder (1992) argue that shocks to the federal funds rate sensitively measure policy shocks in the market for reserves.¹⁰

To identify monetary policy shocks we begin with the monthly federal funds rate, and regress the funds rate on twelve own lags and twelve lags of each of the following series: the growth rate of total reserves, the growth rate of the Federal Reserve Board's industrial production index, and the growth rate of the consumer price index for all urban wage earners and clerical workers. F-tests reject the hypothesis that any of the independent variables should be excluded from the regression at the 1% significance level. In addition, an F-test for a structural break in October 1979 rejects the null hypothesis of no break at the 1% level. With this in mind, we rerun the regression including a binary variable and interaction terms for all right-hand side variables in order to reflect the structural break. The residuals from this regression are our measure of monetary policy shocks. Table 3 contains summary statistics for the final specification.

The next step is to test for an ARCH process in the residuals. The general form of the variance process of the ARCH model is:

$$(5) \quad h_t = \delta_0 + \sum \delta_j \epsilon_{t-j}^2$$

where h_t is the conditional variance of the process, ϵ_{t-j}^2 is the squared residual from the generating regression, and $\delta_0 > 0$ and $\delta_j \geq 0$ for all j . Both the intercept

¹⁰Bernanke and Blinder (1992) find that the federal funds rate works well both before and after October 1979, which is commonly believed to be the date on which the Federal Reserve switched from targeting the federal funds rate to targeting nonborrowed reserves. Our results hold for another measure of monetary policy suggested by Bernanke and Blinder (1992), the interest rate spread.

TABLE 3
 Summary Statistics
 Regressions to Estimate Monetary Policy Shocks,
 Relative Fuel Price Shocks, and Fiscal Policy Shocks

	Summary Regression Statistics	Estimated ARCH Process+
Federal Funds Rate (Monthly, 1965:1-1995:12)		
$\bar{R}^2$	0.97	$h_{ff,t} = 0.13 + 0.51 \sum_{t=1}^8 w_t \epsilon_{ff,t-\tau}^2$
F-Stat (P Value)	0.000	(0.06) (0.10)
Q-Stat (P Value)	0.55	$w_t = \frac{9-\tau}{36}$
LM Test ARCH Process (8 Lags)	61.4**	
Relative Fuel Price (Monthly, 1967:1-1995:12)		
$\bar{R}^2$	0.21	$h_{fu,t} = 732.5 + 0.102 \epsilon_{fu,t-1}^2$
F-Stat (P Value)	0.000	(242.9) (0.05)
Q-Stat (P Value)	0.79	
LM Test ARCH Process (1 Lag)	3.7*	
Real Federal Government Defense Purchases (Quarterly, 1967:1-1995:12)		
$\bar{R}^2$	0.23	$h_{gfd,t} = 43.74 + 0.6 \epsilon_{gfd,t-1}^2$
F-Stat (P Value)	0.000	(7.78) (0.09)
Q-Stat (P Value)	0.99	
LM Test ARCH Process (1 Lag)	3.1*	

+ standard errors in parentheses.

* indicates significance at the 10% level.

** indicates significance at the 1% level.

and the coefficient are restricted to be nonnegative to ensure that the estimated variance process does not become negative. In addition, the sum of the lag coefficients must be less than one to ensure dynamic stability of the variance process. Estimates of the δ 's come from the OLS regression of the squared residuals on a constant and a set of own lags, as in Equation (5). Using the regression results, the conditional variance, h_t , can be consistently estimated.

Applying this procedure to the residuals from the federal funds regression, we regress the squared residuals on twelve own lags. We drop insignificant lags as long as this does not increase the AIC criterion. An LM test for the ARCH process rejects the null of conditional homoscedasticity at the 1% level for a specification retaining eight lags, which indicates that the squared errors follow an ARCH process.

To estimate the ARCH process, we follow Engle (1982) and Engle, Lilien, and Robins (1987) and employ a two-parameter variance process, which eases the

burden of meeting the theoretical parameter restrictions. The two-parameter estimation imposes a set of linearly declining weights on past residuals. Our measure of the variability of monetary shocks, $h_{ff,t}$ is generated by the equation shown in Table 3.

In the IRPV, REV, and RREV regressions to follow, we will also include two other possible sources of dispersion of regional outcomes: oil price shocks and shocks to federal government defense purchases. We estimate the variability of oil price shocks using a measure of the relative price of motor fuel, calculated by dividing the U.S. CPIW for motor fuel by the U.S. CPIW for all items. We apply the same general procedure as above to identify the variability of relative fuel price shocks, with the results also presented in Table 3. The results reject the null of conditional homoscedasticity at the 6% level and we estimate the ARCH process as above. We will refer to our measure of oil price variability as $h_{fu,t}$.

For real federal government defense purchases (in constant 1987 dollars), we reject the null of conditional homoscedasticity at the 8% level and estimate the ARCH process as before. The results are presented in Table 3 and we will refer to our measure of the variability of real federal government defense purchases as $h_{gfd,t}$.

V. AGGREGATE VARIABILITY AND REGIONAL DISPERSION

The hypothesis to be tested is that an increase in the variance of monetary shocks increases the regional relative inflation and wage growth variance. Regression of IRPV on a set of own lags and on current and lagged values of $h_{ff,t}$, $h_{fu,t}$ and $h_{gfd,t}$ provides a test of this hypothesis. Consider the following autoregressive distributed lag model:

$$\begin{aligned} \text{In IRPV}_t = & a_0 + a_1 \text{In IRPV}_{t-1} + \dots \\ (6) \quad & + b_0 \ln h_{ff,t} + b_1 \ln h_{ff,t-1} + \dots \\ & + c_0 \ln h_{fu,t} + c_1 \ln h_{fu,t-1} + \dots \\ & + d_0 \ln h_{gfd,t} + d_1 \ln h_{gfd,t-1} + \dots + e_t, \end{aligned}$$

where $\ln$ indicates the natural logarithm and e_t is the regression residual. We estimate Equation (6) in logs to ensure that the estimated regional inflation and wage growth variance processes do not become negative. The autoregressive distributed lag framework is attractive in this context because it takes time for the impacts of monetary policy to filter through the economy. For instance, if monetary policy causes interest rates to change, this will impact private investment spending decisions, which typically take a significant amount of time to implement.

Significant coefficients on h_{ff} in this regression indicate that IRPV can be explained partly by the variance of monetary shocks. Eight own lags and current and eight lags of $h_{ff,t}$, $h_{fu,t}$ and $h_{gfd,t}$ were considered because empirical work in the area suggests that this is the lag length at which the regional impacts of monetary policy are at their greatest (two examples here are Carlino and DeFina 1995, 1998 and Garrison and Chang 1979). We eliminate lags that are insignificant at the 10%

level as long as this does not increase the AIC criterion or introduce significant serial correlation. As shown in Equation (7), with standard errors in parentheses, the final regression contains two own lags of IRPV and current and six lagged values of $h_{ff,t}$ and current and three lags of $h_{fu,t}$. The lag selection process eliminates $h_{gfd,t}$ from the regression altogether.

(7)

$$\begin{aligned} \ln \text{IRPV}_t = & 12.30 + 0.04 \ln \text{IRPV}_{t-1} + 0.32 \ln \text{IRPV}_{t-2} + 0.99 \ln h_{ff,t} \\ & (6.11) \quad (0.09) \qquad (0.09) \qquad (0.39) \\ & - 0.42 \ln h_{ff,t-1} - 0.70 \ln h_{ff,t-2} + 1.31 \ln h_{ff,t-3} \\ & (0.61) \qquad (0.64) \qquad (0.64) \\ & - 1.04 \ln h_{ff,t-4} + 0.07 \ln h_{ff,t-5} + 0.85 \ln h_{ff,t-6} \\ & (0.66) \qquad (0.64) \qquad (0.41) \\ & + 0.43 \ln h_{fu,t} - 0.27 \ln h_{fu,t-1} - 0.11 \ln h_{fu,t-2} \\ & (0.48) \qquad (0.47) \qquad (0.47) \\ & - 1.67 \ln h_{fu,t-3} \\ & (0.47) \end{aligned}$$

The variability of monetary shocks significantly impacts IRPV. As shown in Table 4, the $h_{ff,t}$ terms are jointly significant at the 1% level. In addition, the sum of the lag coefficients is positive and significantly different from zero at the 1% level. To put these results in perspective, note that the long-run elasticity of IRPV with respect to monetary variance, computed from Equation (7), is 1.66. This implies that a 10 percent increase in the long-run variance of monetary policy results in a 16.6 percent increase in the long-run variance of relative inflation rates (and thus an 8 percent increase in the standard dispersion of regional inflation rates around the U.S. average). Thus, variability of U.S. monetary policy has substantial impacts of the variance of regional relative inflation rates.

TABLE 4
Summary Statistics
IRPV, REV, and RREV Regressions
(Quarterly, 1967Q1-1995Q4)

Regression	F-Tests on $h_{ff,t}$, $h_{fu,t}$ and $h_{gfd,t}$ (P Values)						Serial Correlation Q Statistic (P Values)	
	$h_{ff,t}$ ^a		$h_{fu,t}$ ^b		$h_{gfd,t}$ ^c		$\bar{R}^2$	36 Lags ^d
IRPV	Joint	Sum	Joint	Sum	Joint	Sum	0.37	41.6 (0.24)
REV	0.00	0.00	0.01	0.08	**	**	0.17	37.6 (0.40)
RREV	**	**	0.22	0.07	0.00	0.38	0.11	36.0 (0.47)
	0.07	0.32	**	**	**	**		

*. * Eliminated from final specification by lag selection process.
^a $h_{ff,t}$ is ARCH estimated monetary variability, $h_{ff,t}$ is the average of monthly values.
^b $h_{fu,t}$ is ARCH estimated oil price variability, $h_{fu,t}$ is the average of monthly values.
^c $h_{gfd,t}$ is ARCH estimated federal defense purchases variability.
^dNull Hypothesis for the Q Statistic is No Serial Correlation.

While current and lagged $h_{gfd,t}$ are not individually or jointly significant in the IRPV regression, current and three lags of $h_{fu,t}$ are significant. In addition, their sum is negative and significantly different from zero at the 8% level, which

implies that an increase in the variability of oil price shocks reduces regional relative inflation variance.

One interpretation of this unexpected result is that oil price shocks are similarly represented in average inflation rates across regions, and that these common, direct price impacts overwhelm existing inflation rate differences. Computing an index of the variability of motor fuel price inflation across regions contributes to this view. The dispersion of motor fuel inflation rises only mildly during 1979 and falls during the early 1980s. Overall, this index is mildly negatively correlated with IRPV.

With regard to variability in defense purchases, previous reduced-form investigations (Garrison and Chang 1979; Mathur and Stein 1980; Garrison and Kort 1983) have found that government spending tends to have more similar impacts across states and regions than does monetary policy. Indeed, these studies find that the variance of fiscal policy response is one-third to one-quarter that of the variance of the monetary policy response. The similarity of regional responses to aggregate fiscal policy, combined with the level of aggregation employed, probably contributes to the lack of significance of defense purchase variability.

The regional price impacts of monetary policy may not be restricted to product prices, but may also affect the dispersion of wage growth rates. To explore this possibility, we regress our measures of monetary, oil price, and federal defense purchases variability on REV and RREV. Summary regression results are again contained in Table 4. For REV, there is no strong correlation with the variance of monetary policy. None of the coefficients on $h_{ff,t}$ are significantly different from zero at the 10% level and an F-test on the joint significance of current and eight lags of $h_{ff,t}$ fails to reject the null at the 10% level. Equation (8) shows the final specification (with standard errors in parentheses):

$$\begin{aligned}
 (8) \quad Vt = & -1.34 + 0.20 \ln \text{REV}_{t-1} - 0.05 \ln \text{REV}_{t-2} + 0.08 \ln \text{REV}_{t-3} + 0.19 \ln R \\
 & (8.86) \quad (0.10) \quad (0.10) \quad (0.10) \quad (0.10) \\
 & + 0.13 \ln h_{fu,t} + 0.49 \ln h_{fu,t-1} + 0.92 \ln h_{fu,t-2} \\
 & (0.50) \quad (0.50) \quad (0.51) \\
 & + 0.50 \ln h_{gfd,t} - 1.80 \ln h_{gfd,t-1} + 1.12 \ln h_{gfd,t-2} \\
 & (0.50) \quad (0.51) \quad (0.52) \\
 & + 0.11 \ln h_{gfd,t-3} - 0.63 \ln h_{gfd,t-4} + 1.54 \ln h_{gfd,t-5} \\
 & (0.53) \quad (0.55) \quad (0.53)
 \end{aligned}$$

Current and two lags of $h_{fu,t}$ survive the lag selection process. These are jointly significantly different from zero at the 22% level. In addition, current and five lags of $h_{gfd,t}$ also survive and are jointly significantly different from zero at the 1% level. The sum of the lags of $h_{gfd,t}$ is positive, but different from zero at the 38% level. This suggests that variability in federal defense purchases does have an impact on the regional dispersion of wage growth rates, but that the effect largely dissipates within five quarters. Overall, REV does not seem to be strongly

related to monetary policy, federal defense spending, or oil price shocks. This may be due to the higher rate of unionization in manufacturing and the prevalence in manufacturing industries of unsynchronized, multiyear contracts that set nominal wages. Further, the economics literature suggests that what workers, firms, and policy makers really care about is the real wage, and it is to this issue that we now turn.

For RREV, one own lag and current and five lags of $h_{ff,t}$ are retained, as shown in Equation (9):

$$\begin{aligned}
 \ln RREV_t = & 0.82 + 0.29 \ln RREV_{t-1} - 0.03 \ln h_{ff,t} + 0.52 \ln h_{ff,t-1} \\
 & (0.35) \quad (0.09) \quad (0.37) \quad (0.59) \\
 (9) \quad & - 0.74 \ln h_{ff,t-2} + 1.49 \ln h_{ff,t-3} - 1.62 \ln h_{ff,t-4} \\
 & (0.61) \quad (0.59) \quad (0.56) \\
 & + 0.57 \ln h_{ff,t-5} \\
 & (0.32)
 \end{aligned}$$

The coefficients on current and lagged $h_{ff,t}$ are jointly significant at the 7% level and the sum is positive although not significantly different from zero at the 10% level. All lags of $h_{fu,t}$ and $h_{gfd,t}$ are eliminated in the lag selection process. The results for real wage growth echo earlier investigations that found significant asymmetric impacts of monetary policy on real income.

Evidence from previous investigations suggests that regions specializing in interest-sensitive industries will be affected more by national monetary policy. If we focus on industry shares in durable manufacturing, the results are generally consistent with this hypothesis. During the 1980-83 period, which is a period of increased monetary variability, Midwest relative inflation rates record a higher standard deviation than the Northeast region, which in turn exceeds the standard deviation of the South region's relative inflation. The West census region breaks the pattern by recording the largest variability during this period, as well as on average during the 1967-95 period (see Table 2).

The results are similar for real wage growth during the 1980-83 period. Again, real wage growth variability in the Midwest exceeds that of the Northeast, which in turn exceeds that of the South. As with relative inflation, wage growth variability is greatest in the West region.

VI. CONCLUSION

This paper investigated the impact of monetary policy on regional inflation and wage growth rates in the U.S. The empirical results suggest that increased monetary variability increases the variance of regional inflation and real wage growth rates around the U.S. average. These results highlight an unintended regional consequence of variable national monetary policy, which is aimed at stabilizing the national economy. However, because regions within the overall macroeconomy differ in their industrial structure (and therefore also differ in their sensitivity to national monetary policy), national monetary policy creates asym-

metric regional responses.

The results for real wages confirm previous work that has found that national monetary policy has important asymmetric regional impacts on income growth. The results of this paper also suggest that national monetary policy has important impacts on the dispersion of regional inflation differentials.

During the 1967-95 period, IRPV averaged 1.3 percent, which implies an average dispersion of regional inflation rates of 1.14 percentage points around the national average. At first glance, the regional dispersion of regional inflation rates may seem small. However, as Feldstein (1997) highlights, the costs of moderate inflation may be quite high. With uneven inflation rates across regions, the welfare costs of inflation are likely to be unevenly distributed as well. Further, taken together with the results of Section V, this suggests that an increase in monetary variability causes an additional economic burden by contributing to the dispersion of inflation rates across regions.

What might the size of this distortion be? A simple back-of-the-envelope calculation relying on Feldstein (1997) will give us a general idea. Feldstein (1997) estimates that a U.S. inflation rate of 2 percent implies an aggregate annual welfare cost of between 0.8 and 1 percent of U.S. GDP. Using the lower bound estimate of 0.8 percent and assuming that if IRPV equals zero the aggregate welfare costs are evenly distributed across regions, we have each of the four census regions bearing a cost equal to 0.2 percent of U.S. GDP.

What if inflation rates differ across regions? We have estimated that on average during the 1967-95 period the standard dispersion of regional inflation rates (computed using the square root of IRPV) is 21 percent of the average U.S. inflation rate. Taking this as given, a U.S. inflation rate of 2 percent implies that the standard dispersion is 0.42 percentage points. To simplify, assume that all regions are the same size and that two regions register inflation rates of 2.42 percent and two regions register inflation rates of 1.58 percent. The U.S. average inflation rate in this case is 2 percent.

Using Feldstein (1997), this means that the two high-inflation regions bear welfare costs equal to 0.242 percent of U.S. GDP and the low-inflation regions bear welfare costs equal to 0.158 percent of U.S. GDP. The high-inflation regions bear inflation costs 21 percent ($0.242/0.2$) above average and the low-inflation regions bear inflation costs 21 percent below the national average (assuming that welfare costs from inflation are proportional to the rate of inflation, at least in the neighborhood of 2 percent).

What is the impact of monetary variability? We have estimated that a 10 percent increase in monetary variability increases the standard dispersion of regional inflation rates by 8 percent. Thus, a 10 percent increase in monetary variability increases the standard dispersion from 0.42 percent to 0.45 percent. In turn, this increases the *annual* welfare cost born by high-inflation regions by 0.003 percent (from 0.242 percent to 0.245 percent) of U.S. GDP. To put this increase in welfare cost into perspective, 0.003 percent of U.S. GDP in 1994 was \$208.4

million, which ranged from 0.014 percent of Northeast census region GSP to 0.009 percent of South census region GSP.

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