One for Some or One for All?
Taylor Rules and the Inter-Regional Unemployment Dispersion Puzzle

Olivier Coibion and Daniel Goldstein
University of Michigan*

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Abstract: We document a robust and surprising empirical phenomenon: both the U.S. Federal Reserve and the European Central Bank appear to set interest rates partly in response to regional disparities in unemployment rates. This result is exceedingly robust, even after controlling for a wide variety of factors, including the central bank's information set and a battery of explanatory variables. Furthermore, including measures of inter-regional unemployment dispersion in Taylor rule estimates also helps improve the identification of the central banks' responses to inflation and unemployment rates. We propose a variety of statistical and theoretical possibilities to account for this puzzling empirical result, but find that none is consistent with our findings.

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1 Introduction

This paper poses a simple question: do central banks systematically respond inter-regional variation in economic activity above and beyond that measured in aggregate variables? Surprisingly, we find very robust empirical evidence that regional variation in unemployment rates helps explain interest rate policies of both the U.S. Federal Reserve and the European Central Bank. In addition, we find that the inclusion of measures capturing the dispersion of unemployment rates across regions in a Taylor rule helps improve the identification of the response of the central bank to aggregate variables, particularly in the United States. We propose a list of possible explanations – both statistical and theoretical – that might explain this puzzling result, but find that none can adequately account for our findings.

The question of whether regional economic differences affect policy-makers' decisions is of growing importance as countries increasingly choose to surrender their independent national central banks in favor of multinational ones. Nations agree to give up independent monetary policies in exchange for increased exchange rate stability with major trading partners and as a mechanism to credibly lower inflationary expectations. Yet as central banks gain control of monetary policy over more heterogeneous economic entities, the pressure to respond to or accommodate off-cycle regions may rise. The possibility that central banks systematically respond to regional heterogeneity in economic welfare, rather than just to aggregate measures, has important implications.

For example, with Central and East European countries being considered for admission to the Euro-Zone, the degree of heterogeneity in regional welfare would likely increase for the European Central Bank. Additionally, the formation of a central bank and common currency for the entire continent of Africa is on the horizon. To the extent that responding to regional concerns is sub-optimal for aggregate welfare, as theory implies, then expansion of monetary zones to new areas could negatively affect the performance of these institutions.

A vast literature is devoted to the study of how central banks set interest rates. Following Taylor (1993), who found that the Federal Reserve's interest rate could be adequately represented via a simple rule in which interest rates change mechanically with various aggregate variables, empirical work has treated Taylor rules as a baseline for modeling a central bank's behavior.[1] This approach has been supported by the fact that optimal policy in simple New Keynesian models can often be

[1] See Clarida, Gali, and Gertler (2000) for such an example focusing on whether the behavior of US Federal Reserve has changed over time.

represented by such a rule.[2] In addition, the explicit goals laid out for central banks in their founding legal documents focus on a few aggregate variables, such as the stability of the price level and maximizing employment. The use of simple Taylor rules linking interest rates to aggregate inflation, output growth, and/or unemployment rates thus has strong empirical, theoretical and legal support.

We use this framework to address whether central banks respond to regional economic variables once aggregate variables are taken into account. In short, for our baseline model we use 2SLS to estimate Taylor rules for the U.S. Federal Reserve and European Central Bank and test whether various measures of inter-regional unemployment rate dispersion offers any additional explanatory power after taking account of expectations of future aggregate inflation and unemployment rates.

Our empirical results strongly reject the null that these central banks do not respond to the inter-regional dispersion of unemployment rates. The results hold for various measures of dispersion, such as gaps between high and low unemployment regions, or the weighted variance of unemployment rates across regions for each period. Interestingly, the estimated parameters in the Fed and the ECB models are remarkably similar to each other. This is true for both our dispersion measures and the other – conventional – parameters of the Taylor rule, though the latter is only true *after* including our measure of dispersion.

To understand the implications of this finding, we first verify that it is not a statistical anomaly. For this purpose, we consider whether our findings are robust to the inclusion of a variety of leading indicators and other economic variables that could be relevant to the central banks' expectations for aggregate inflation and unemployment, such as stock prices, oil prices, …, and find that our results are largely robust to the inclusion of such measures.

Another concern is that our unemployment dispersion measures could help forecast future values of inflation and unemployment. In this case, finding significant coefficients on the former could simply reflect a failure to adequately capture the central banks' expectations of future inflation, output growth, or unemployment. We address this important concern in two ways. First, we show that there is little evidence that our measures of regional heterogeneity are useful predictors of future values of the aggregate variables. Second, we reproduce our estimates using GreenBook forecasts from the Federal Reserve of future aggregate variables. Thus, we can control for the central bank's

[2] See Woodford (2003) for a thorough discussion of monetary policy in New Keynesian models.

information set. Even here, we continue to find an independent role for our inter-regional dispersion measures.

Having found that our results do not appear to be a statistical anomaly, we then consider some possible explanations for this result. The first theoretical possibility is simply that the central bank could care about differences in economic welfare across regions. We develop a simple model that allows for such a feature of central bank preferences and derive testable implications involving how the central bank should respond to the variance and skew of the distribution of unemployment rates across regions. Taking these predictions to the data, we find little evidence that central banks act in a way consistent with trying to minimize differences in economic welfare across regions.

A second, if related, possibility is that the central bank might care more about a certain type of region than others. The most likely possibility is that the central bank would place a disproportionate weight on states with high unemployment. To address this possibility, we break our regional dispersion measures into two components: the difference between a specific percentile of the unemployment distribution and the mean unemployment rate. In other words, this is asking whether the central bank responds in the same way to high unemployment states seeing higher unemployment and low unemployment states reaching even lower unemployment levels. For the ECB, we cannot reject the null that the central bank responds identically to low and high unemployment states. Interestingly, for the Fed, interest rates rise as low-unemployment states see their unemployment rates fall relative to the mean, but are unchanged as high-unemployment states move away from the mean.

Finally, we consider the possibility that institutional features of central bank voting patterns could account for this result. For example, we show that voting records of FOMC meetings since 1982 imply a disproportionately large representation of the New York and Boston districts, as well as disproportionately low representation of the Atlanta and San Francisco districts, as measured by votes of regional presidents and board members. In the ECB, the "big three" of Germany, France and Italy are substantially underrepresented in meetings of the Governing Council whereas small countries are heavily over represented. Small countries in the ECB have also had much lower unemployment rates than big countries (Finland is the lone exception). We construct measures of the weighted sum of differences between regional and aggregate unemployment where the weights reflect the voting power of the respective regions at each meeting of the FOMC and Governing Council of the ECB. We find that our original measures of the dispersion of regional unemployment rates are robust to including these series, and in fact find no evidence that these series affect interest rates in any way.

The paper is organized as follows. Section 2 presents our estimation approach, baseline results, and some robustness checks. Section 3 focuses on whether our results are due to a correlation between the central bank's information set and our measures of regional heterogeneity in unemployment rates. Section 4 proposes and tests potential explanations based on alternative objective functions for the central bank and institutional features of voting patterns. Section 5 concludes.

2 Baseline Results

The official goals of the US Federal Reserve, as stated in the Federal Reserve Act, are to

> "maintain long run growth of the monetary and credit aggregates commensurate with the economy's long run potential to increase production, so as to promote effectively the goals of maximum employment, stable prices, and moderate long-term interest rates."

while those of the European Central Bank, as laid out by the Treaty on European Union, are

> "The primary objective of the ESCB shall be to maintain price stability. … without prejudice to the objective of price stability, the ESCB shall support the general economic policies [a high level of employment and sustainable and non-inflationary growth[3]] in the Community with a view to contributing to the achievement of the objectives of the Community as laid down in Article 2." (Treaty article 105.1)

These two institutions, by mandate, therefore have very similar objectives: price stability and employment/output growth, though it has been emphasized that the ECB should, by mandate, place more emphasis on price stability. Nonetheless, to the extent central bankers of each institution are dedicated to achieving their stated goals, the objective function of the two central banks should be very similar.

2.1 Estimation strategy and data

Given that interest rates are the primary tool used by these central banks in achieving their goals, policy-makers' decisions are naturally modeled by an interest rate rule of the type proposed by Taylor (1993)

$$i_t = \phi_\pi E_t \pi_{t+j} + \phi_{ue} E_t ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \varepsilon_t \qquad (1)$$

[3] Added by authors and drawn from Article 2 of Treaty on European Union. The "ECSB" is the European System of Central Banks, composed of the European Central Bank as well as national central banks.

Such a rule implies that interest rates rise by ϕ_π (ϕ_{ue}) basis points on impact when expectations of inflation (unemployment) rises by one percentage point. Lagged interest rate terms are included to allow for interest-smoothing. The i.i.d. (by assumption) error term ε_t represents monetary policy shocks. This specification assumes that interest rates are set in response to current expectations of future values of the independent variables, capturing the well-known fact that monetary policy acts with a lag, forcing policy-makers to be forward-looking.

Because expectations are not always available, our primary estimation strategy will employ the additional assumption of rational expectations for policy-makers, such that $E_t x_{t+j}=x_{t+j}+v_{t+j}$ where v_{t+j} is unforecasteable using time t information. Substituting this into equation (1) yields

$$i_t = \phi_\pi \pi_{t+j} + \phi_{ue} ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \zeta_t \tag{2}$$

where ζ_t consists of the monetary policy shock and the sum of rational expectations errors. Thus, following from our assumptions, $E_{t-j}\zeta_t=0$ for all $j\geq 1$. Because future values of inflation and unemployment can be expected to be correlated with time-t monetary policy shocks, we propose to use 2SLS to estimate the parameters of equation (2).

We use monthly data from January 1982 to September 2005 for the U.S. and from January 1999 to September 2005 for the Euro-Zone. For the US, we use the effective federal funds rate as our primary measure of interest rates, the 12-month log percentage change in the CPI for inflation and the BLS series for aggregate unemployment rates. For the Euro-Zone, we use the interbank overnight rate for our interest rate series, and harmonized aggregate inflation and unemployment rates. We use a six-month forecast horizon, though the results are largely insensitive to this assumption.

Throughout the paper, we allude to the inter-regional dispersion of unemployment rates. In practice, we measure this dispersion in a variety of ways. The first is to take the difference between the unemployment rates of the 90th and 10th percentiles of the time-t distribution of regional unemployment rates (UEP9010). The second is a narrower band: the difference between the 75th and 25th percentiles (UEP7525). Our third measure is to compute the variance of the distribution of unemployment rates each month, weighted by the population share of each region (var(UE)). We define regions in the US as each of the fifty states plus the District of Columbia. For the ECB, each region is one of the eleven member states.

Figure 1 plots interest rates, inflation, and aggregate unemployment for the US over our time sample, while Figure 2 plots our three measures of regional heterogeneity in UE rates. The first thing worth noting is that these measures are broadly similar (all cross-correlations exceed 0.9). However, they do exhibit some different patterns. In particular, the variance of unemployment rates went up much more in the 1982 recession than is true of the other series. One similarity is that all three track the aggregate unemployment rate, by rising in recessions and falling as aggregate unemployment falls. This property fails in two instances. The first is the 2000 recession: as aggregate unemployment rose with the recession, none of the dispersion measures changed in this time period, indicating that the last recession was borne similarly by all states. The second occurs in 1986, when the price of oil fell dramatically. This led to sizeable increases in unemployment in oil-producing states, which shows up as an upsurge in dispersion of unemployment rates with no commensurate increase in aggregate unemployment.

Figures 3 and 4 plot aggregate variables and our measures of unemployment dispersion respectively for the Euro-Zone. Again, the three dispersion measures are broadly similar, with a gradual convergence of unemployment rates being the most prominent feature of each series. Unlike with the US, these series are uncorrelated with aggregate unemployment.

To test whether these measures affect central bank decision-making, we augment equation (2) with a measure of regional dispersion of unemployment rates

$$i_t = \phi_\pi \pi_{t+j} + \phi_{ue} ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \beta D_t + \zeta_t \tag{3}$$

where the null is that $\beta=0$ and D_t is a measure of regional unemployment dispersion. As instruments in our 2SLS regression, we will consistently use six lags of the endogenous RHS variables (inflation, unemployment, and UE dispersion), a constant, and the same lags of the interest rate used in estimating (3).[4]

2.2 Estimation Results

Our basic results are presented in Table 1. Consider first the Baseline results for the US (panel A), which exclude UE dispersion measures. The coefficient on future inflation is positive and highly significant, but, surprisingly, we cannot identify a statistically significant response of interest rates to the aggregate unemployment rate. The coefficients on lagged interest rate imply an important amount of interest smoothing (the sum of the coefficients is 0.97), yielding a long-run response to inflation of

[4] Standard errors are Newey-West HAC with a truncation at 6 lags to account for the overlapping errors due to the use of 6-month ahead values on the RHS.

about 3, consistent with the post-1982 estimates of Clarida, Gali and Gertler (2000). For the ECB (panel B), the degree of interest smoothing is once again high. However, in a complete reversal of the U.S. results, we can now identify a strong response to unemployment, but no statistically significant response to aggregate inflation.

Including a measure of inter-regional dispersion does two things. First, for both the U.S. and the ECB, it yields estimates of the Taylor rule that are more consistent with central banks responding to both aggregate inflation and unemployment rates. For the US, this means the response to unemployment becomes negative and statistically significant, while for the ECB the response to aggregate inflation becomes positive and significant. Second, each of the measures of UE dispersion enters the regression with a positive and highly significant coefficient. A positive coefficient implies that *as the degree of heterogeneity in UE rates increases across regions, the central bank tends to raise interest rates*. Interestingly, the results are similar for the ECB and the Fed, with the Fed seemingly responding more strongly to UE dispersion than the ECB.[5]

Because statistical significance need not imply economic significance, we consider how important our measures of regional dispersion are in explaining interest rate volatility. To do so, we break our equations into exogenous and endogenous components of interest rates such that

$$Endog_t \equiv \phi_\pi E_t \pi_{t+j} + \phi_{ue} E_t ue_{t+j} + \beta D_t \tag{4}$$

The endogenous component of interest rates is thus defined here as that due to expectations of future inflation and unemployment, as well as the response of the interest rate to our measures of the dispersion of regional unemployment rates. Movements in this endogenous component of interest rates can thus be decomposed into changes in expectations of aggregate variables ($agg_t \equiv \phi_\pi E_t \pi_{t+j} + \phi_{ue} E_t ue_{t+j}$) and dispersion measures. The variance of the endogenous component of interest rates is thus

$$\operatorname{var}(Endog_t) = \operatorname{var}(agg_t) + \beta^2 \operatorname{var}(D_t) + 2\beta \operatorname{cov}(agg_t, D_t). \tag{5}$$

We present the results of such a decomposition in Table 2, normalizing each of the RHS elements of (5) by the variance of the endogenous movements of the interest rate.

For the Euro-Zone, the covariance of the aggregate and regional measures is close to zero for each measure of regional unemployment dispersion. The variance of the aggregate component

[5] We also found similar results when looking at the dispersion of inflation rates, as well as for dispersion of real exchange rates for the ECB. However, the inflation results were much more sensitive for the US, reflecting the fact that regional price level data is only at a monthly frequency for the four Census Bureau divisions. The RER series for the ECB were very highly correlated with the UE dispersion measures, so we focus exclusively on the unemployment measures here.

accounts for two-thirds of the endogenous response of interest rates, while the variance of the regional dispersion measure approximately accounts for the remaining third. For the U.S., changes in aggregate variables account for 57% to 87% of the endogenous variance in interest rates, the rest being due to the variance in regional unemployment dispersion and the covariance term. The large positive covariance between aggregate unemployment and regional heterogeneity in unemployment rates translates into a negative covariance between the aggregate and regional components of the endogenous movements in the interest rate. The implication is that, if the results of Table 1 are indeed capturing an important element of policy-making, regional differences in unemployment rates have economically important effects on interest rates and account for (very) approximately one-third of the endogenous movements in interest rates.

2.3 Robustness.

As a first step to investigating the robustness of these results, Figure 5 provides a scatterplot of the US (weighted) cross-sectional variance of state unemployment rates against the orthogonalized component of interest rates. A clear positive relationship exists regardless of the outliers that appear. (These outliers are almost exclusively from 1982.) Not surprisingly given this scatter plot, we have found the positive relationship between the degree of heterogeneity in states' UE rates and interest rates to be very robust to sub-sample analysis.[6] Unfortunately, given the short time sample available since the inception of the ECB, no time-sample verification can be provided for the ECB.

A second type of robustness check is to consider an alternative estimation approach. Following Clarida, Gali, and Gertler (2000), we estimated equations (2) and (3) by GMM with Newey-West weighting matrix. The results were qualitatively the same, but all the p-values were smaller. Thus, all our results hold even more convincingly when done by GMM. Because of the very short time sample available for the ECB (80 observations) and the well-known poor properties of GMM in such short time samples, we focus on the 2SLS estimates.[7]

A third issue to be concerned about is whether our choice of interest rates is the correct one. For example, the U.S. Federal Reserve chooses a target for the Federal Funds rate (FFR), from which

[6] See Appendix Table 1 for results for US in different time-samples.

[7] See Hansen, Heaton, and Yaron (1996) and Christiano and den Haan (1996) for discussions of how GMM estimators fare in short time-samples. We reproduce the results of Table 1 using GMM with Newey-West weighting matrix in Appendix Table 2.

the effective FFR may differ, sometimes for reasons unrelated to monetary policy.[8] It is therefore possible that our measures of heterogeneity in regional unemployment rates could be picking up shocks that affect both the endogenous movements of the effective FFR around the target rate and regional asymmetric responses to these shocks. To see whether this is important, we reproduced the results of Table 1 using the target FFR for the Fed and the refinancing rate for the ECB and found nearly identical results (see Appendix Table 3).[9] We also find that our results are insensitive to using the GDP deflator or the non-farm business deflator to calculate inflation, rather than the CPI.

3 Capturing the Central Bank's Information Set

While the above results thus do not appear to be a statistical anomaly, a fundamental problem with the estimation procedure presented above is that if the central bank's information set is not properly conditioned for in 2SLS, finding that measures of the regional dispersion of unemployment rates are statistically significant predictors of interest rates could just be reflecting some predictive power of these measures for future values of the aggregate variables. To see this, suppose that the central bank's forecasts of future inflation and unemployment rates contain much more information than is embodied in our instruments. In this case, if dispersion measures are useful predictors of future aggregate measures, then they may show up as significant predictors in a Taylor rule simply because they are capturing elements of the central bank's information set that we are not controlling for.[10]

We consider three ways of addressing this legitimate concern. The first is determining whether our dispersion measures do appear to contain useful information for predicting future aggregate variables. The second consists of augmenting our Taylor rule estimates with variables that are well-known leading indicators. The third is to control directly for the central banks' expectations by using the real time forecasts used by these agencies in making their interest rate decisions. We find our results to be robust to all three possibilities.

3.1 Do Inter-Regional Unemployment Dispersion Measures Forecast Aggregate Variables?

Figure 6 plots the dynamic cross-covariances of the weighted variance of regional unemployment rates each month against leads and lags of aggregate inflation and unemployment rates for the U.S. and the

[8] See Romer and Romer (2004).
[9] The only difference is that future inflation is no longer statistically different from zero in the Euro-Zone estimates, likely reflecting how little variation there is to explain given the discrete and infrequent changes in the Euro refinancing rate.
[10] See Orphanides (2001) for a discussion of the importance of controlling for the central bank's real time expectations.

ECB over the time periods used for the empirical analysis.[11] For the US, there is little evidence that regional dispersion in unemployment rates tends to lead aggregate measures, instead the opposite appears to be true. Both high inflation and high aggregate unemployment rates tend to be followed by increased dispersion of unemployment rates across states. With aggregate inflation, there is an upswing in the dynamic cross-covariance at higher leads of inflation, which indicates that when the dispersion of unemployment rates is high, inflation tends to be higher over the subsequent time period.

For the Euro-Zone, however, there is stronger evidence that the dispersion of unemployment rates contains information about the future time path of aggregate unemployment. In this case, the peak covariance is between current dispersion and six-month ahead unemployment. However, there is little evidence that our measures of dispersion contain much useful information about future aggregate inflation.

To see this more formally, Table 3 presents Granger-Causality tests of the weighted variance of regional unemployment rates with aggregate inflation and unemployment for the US and Euro-Zones over the same time periods as used in the empirical analysis.[12] For the US, we can weakly (at the 10% level) reject the null that inflation does not granger-cause the variance measure, but fail to reject (with a *p*-value of 13%) the null that the variance of unemployment rates granger-causes inflation. This is consistent the dynamic cross-covariance of the two series in Figure 6. While we cannot reject the null that the unemployment rate does not granger-cause the variance measure at conventional levels, the F-statistic of this test far exceeds that of the reverse, which is also consistent with the dynamic cross-covariances. Indeed, for the US, there is little evidence that our measures of regional heterogeneity in unemployment rates contain information useful for forecasting inflation and unemployment rates.

Turning to the evidence from the Euro-Zone, one can strongly reject the null that inflation does not granger-cause the variance of unemployment rates, but not the reverse. However, the reverse holds true for aggregate unemployment, confirming the intuitive results of Figure 6. Thus, it does appear that the dispersion of unemployment rates across members of the Euro-Zone is a useful predictor of future aggregate unemployment rates. As a result, it is possible that our baseline result for the ECB could be capturing the predictive power of our measures of regional dispersion of unemployment rates for future aggregate unemployment rates. For the US, this appears unlikely since

[11] Use of our other measures of regional dispersion of unemployment rates yields nearly identical results.
[12] Similar results hold using our alternative measures of the regional dispersion of unemployment rates.

these measures appear to be, if anything, lagging, or at least not leading, indicators of inflation and aggregate unemployment.

On balance, these results do not strongly signal that the unemployment and inflation expectations generated by our 2SLS regressions are failing to account for the information contained in the dispersion variables. This is because, especially for the United States, there does not appear to be much extra information in the dispersion variables to be had.

3.2 Including forward-looking variables

An alternative approach to determining whether our measures could be capturing some forward-looking behavior not adequately measured in our specifications is to augment these specifications with additional variables which are typically used as leading indicators. For example, consumer confidence measures are often viewed as having predictive capacity for business cycles. Stock prices are another frequently used measure with strong properties as leading indicator. Finally, because the prices of raw materials and intermediate inputs typically are slow to feed through to final goods prices, these receive much attention as valuable indicators of the future direction of the prices for final goods.

Our test including forward-looking variables consists of augmenting equation (3) such that

$$i_t = \phi_\pi \pi_{t+j} + \phi_{ue} ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \beta_1 D_t + \beta_2 LI_t + \zeta_t \qquad (6)$$

where D_t is one of our measures of regional dispersion of unemployment rates while LI_t is the leading indicator added to the regression. The results are presented in Table 4, using the weighted variance of cross-sectional unemployment rates each month as our measure of dispersion.[13] As forward-looking variables we use consumer confidence indicators, stock prices, WDI oil prices, and PPI inflation.[14] For the US, none of these variables eliminates the influence of the cross-sectional variance of unemployment rates. Only consumer confidence enters the regression with a coefficient that is (weakly) statistically different from zero. But in this specification, the coefficient on aggregate unemployment falls to zero, indicating that the relative contribution of these two series is likely not well identified. For each of the other variables, the results for aggregate inflation and unemployment are close to those found earlier. Again, there is little evidence that the measures of regional dispersion

[13] The results are very similar using the other measures of regional dispersion of unemployment rates.

[14] For the US, these are specifically the University of Michigan Consumer Sentiment Index, the Dow Jones Industrial Average, the spot price of WTI oil, and the PPI-all commodities index. For the Euro-Zone, we use the Consumer Confidence Indicator of the European Commission Consumer Survey, the DAX German stock price index, and the Euro-Zone PPI all industries excluding construction index.

of unemployment rates are capturing forward-looking behavior or information that is inadequately modeled.

For the Euro-Zone, however, the results are somewhat different. For two of the four regressions, the coefficient on the measure of dispersion is insignificantly different from zero at standard levels. Both consumer confidence and PPI inflation receive non-zero coefficients while pushing the coefficients on dispersion measures to zero. However, these cases also lead to changes in the coefficients on aggregate inflation and unemployment. In the case of consumer confidence, as in the US, the coefficient on aggregate unemployment goes to zero, with little change in standard errors. This is surprising given the fact that the response to unemployment was so cleanly identified in the baseline specification for the ECB. When including the PPI, it is the response of aggregate inflation that becomes insignificantly different from zero. This occurs primarily because the standard errors increase rather than because of any change in the coefficient. This increased imprecision in identifying responses to aggregate inflation and unemployment indicates that, given the short time sample involved since the inception of the ECB, the data just does not contain enough variation to adequately identify the contributions of each of the variables included.[15] The validity of including measures of the cross-sectional variation in unemployment rates in the ECB's Taylor rule will likely require a longer time sample to be verified.

3.3 Using Green Book Forecasts

To control for the central bank's information set, the ideal setup make use of the specific real-time forecasts of future variables that the central bank relied on to make their decisions instead of using ex-post realized values of these variables. For the US Federal Reserve, Green Book forecasts are available for much of our time sample and provide expectations for the Fed, at the time of each meeting, of inflation, output growth, and unemployment in the future. Unfortunately, these forecasts are unavailable over the last five years, so our time sample is restricted to January 1982 until December 2000. In addition, forecasts are available for each meeting only, so our time frequency is that of the meetings of the Board of Governors every six weeks. This leaves us with 152 observations.

We estimate the following equation

[15] The statistical insignificance of the regional measures when adding consumer confidence and PPI inflation is quite sensitive. For example, assuming that the 6-month ahead measure of regional dispersion in unemployment rates enters equation (4) instead of the contemporaneous value reverses the result. Similarly, the specific instrument set used in these cases is also important, therefore reinforcing the idea that the time sample is just too short, and with too little variation, to allow us to adequately identify the contribution of each variable.

$$i_t = \phi_\pi E_t\left[\sum_{j=0}^{2}\pi_{Q_j}\right]/3 + \phi_{gy}E_t\left[\sum_{j=0}^{2}gy_{Q_j}\right]/3 + \phi_{ue}E_t\left[\sum_{j=0}^{2}ue_{Q_j}\right]/3 + \rho i_{t-1} + \beta_1 D_{t-1} + \varepsilon_t \qquad (7)$$

where i_t is the new Target FFR chosen at each meeting. The expectations terms are the average Green Book Forecasts of expected inflation, output growth, or unemployment over the current quarter and the subsequent two quarters. The lagged interest rate term is the Target FFR chosen at the previous meeting, and D_{t-1} is the measure of cross-sectional heterogeneity in unemployment rates of the month before that in which the meeting occurs.[16] We include expectations of output growth because these appear to play an important role in affecting interest rate decisions. We use the one-month lag in the dispersion measure to ensure that these series are orthogonal to the error term, which captures monetary policy shocks. Because all of the RHS variables are determined prior to the decision about the new Target FFR, they should all be orthogonal to the error term, thus equation (7) can be estimated by OLS.

The results are presented in Table 5. The first column presents baseline results excluding unemployment dispersion measures. The results are quite consistent with our priors about how the central bank sets interest rates. The coefficients on inflation and output growth are positive and statistically significant, while that on unemployment is negative and also statistically different from zero. Adding our measures of dispersion has little effect on the coefficients on expectations of future aggregate inflation and output growth. However, as in Table 1, the estimated response to expected unemployment becomes larger (in absolute value), more than doubling when the weighted variance of state unemployment rates is added to the regression. The coefficients on measures of unemployment dispersion are still positive, as they were in the estimates presented in Table 1. However, when using the difference between the 75th and 25th percentiles of the distribution of unemployment rates across states, the coefficient is not statistically different from zero. The gap between the 90th and 10^{h} percentiles has a positive coefficient that is statistically different from zero only at the 10% level. The variance measure, however, remains positive and statistically different from zero at the 1% level, indicating that even after controlling for the central bank's expectations (and thereby losing many

[16] We include only one lag of the target FFR because higher order lags are all insignificant and have no effect on other coefficients.

observations), the variance of state unemployment rates continues to be an important predictor of the Federal Reserve's interest rate policy.[17]

After making the above efforts to try and determine whether the initial results of Table 1 were due to a statistical anomaly or a failure to adequately capture the central bank's information set, we conclude that our results appear to be remarkably robust, particularly for the U.S. Federal Reserve. While the ECB results are less clear-cut, increasing the sample size may yield further empirical confirmation in the future. It is particularly striking, however, that augmenting the basic Taylor rule with the unemployment dispersion measures clarifies the parameter estimates on the aggregate variables in such a way that both banks' monetary policies now appear to be consistent with their legal obligations and our priors as to how policy is set. Moreover, the sign on the coefficient for the dispersion variable is the same for each of the two institutions. This qualitative consistency in the monetary policy of the two central banks is an appealing feature of the dispersion-augmented Taylor rules – a feature which was not evident when omitting the dispersion variable. We now turn to possible theoretical explanations of why inter-regional unemployment dispersion might appear to be important for policy.

4 Theoretical Explanations

4.1 Do Central Banks Dislike Regional Heterogeneity?

In our augmented Taylor rules, we consistently find a positive coefficient on our measures of inter-regional unemployment dispersion, implying that central banks tend to raise interest rates in the face of increased regional differences in unemployment rates. We have, in fact, no priors as to how or even why the central bank should respond to such variation. Here, we first consider the simple possibility that the central bank dislikes dispersion in regional welfare, in addition to wanting to maximize aggregate welfare. This is a natural first step, and nests typical models of monetary policy-making that ignore regional differences in welfare. The goal is to identify how the central bank should respond to regional dispersion in unemployment rates if it did care about minimizing such dispersion.

Suppose that each region i has a loss function L_i over aggregate inflation π and its local unemployment rate u_i such that

[17] Because the ECB does not release forecasts used for each meeting, we cannot replicate this analysis for the ECB. This is particularly unfortunate because the results of the previous section indicated that it was more likely that our dispersion measures were capturing forward-looking information of the central bank in the Euro-Zone than in the Fed.

$$L_i = \frac{1}{2}\pi^2 + \frac{\lambda}{2}\left(u_i - \bar{u}_i\right)^2 \tag{8}$$

where we implicitly assume that all regions have the same target rate of inflation equal to zero and place the same relative weight λ on inflation and the deviation of unemployment from its natural level. We do not impose, however, that all regions have the same natural rate of unemployment $\bar{u}_i$. Each region is assumed to face an expectational Phillips Curve relating the deviation of inflation from expectations to the deviation of unemployment from its natural level

$$u_i - \bar{u}_i = -\alpha\left(\pi - \pi^e\right) + \varepsilon_i \tag{9}$$

where π^e is expected inflation (common to all regions) and ε_i is a regional shock to the Phillips Curve.[18]

We define the loss function for the central bank in control of monetary policy for all these regions as

$$L_a = \sum_i \omega_i L_i + \frac{\kappa}{2}\sum_i \omega_i \left[L_i - \sum_i \omega_i L_i\right]^2 \tag{10}$$

The first term is the weighted sum of regional loss functions, where the weights are assumed to be the population or GDP/capita share of each region and sum to 1. We add a second term which captures the squared deviation of each region's loss function from the weighted average of regional loss functions. In other words, when $\kappa>0$, the central bank seeks to minimize the weighted sum of regional loss functions but does not want to do so by inflicting a disproportionate amount of pain to any single region. Of course, when $\kappa=0$, the central bank does not care at all about differences in regional loss functions at all. This oft-cited scenario leads to the following optimal policy, when the central bank chooses inflation conditional on expected inflation

$$\pi^{opt} \equiv \frac{\alpha\lambda}{1+\alpha^2\lambda}\left[\varepsilon_a + \alpha\pi^e\right] \tag{11}$$

where $\varepsilon_a \equiv \sum \omega_i \varepsilon \lambda_i$. The optimal policy, when $\kappa=0$, is independent of each region's loss function and depends only on expected inflation and the aggregate shock, where the aggregate shock is just the weighted sum of regional shocks. In addition, the optimal policy is independent of all higher order moments of the distribution of regional shocks, and therefore of the distribution of regional unemployment rates.

[18] This setup closely follows that of Barro and Gordon (1983).

However, if we impose that the central bank cares about how each region fares relative to the aggregate, the optimal policy depends on the second and third moments of the distribution of regional unemployment rates. Specifically, we can show

Proposition 1: *The optimal choice of inflation for the central bank (π^*) that minimizes (8) subject to (6) and (7) is of the form*

$$\pi^* = \pi^{opt} f\left(\sigma_u^2\right) + skew\left(u_i - \bar{u}_i\right) g\left(\sigma_u^2\right)$$

where $\sigma_u^2 \equiv \sum_i \omega_i \left[\left(u_i - \bar{u}_i\right) - \left(u_a - \bar{u}_a\right)\right]^2$ *is the weighted variance of cross-sectional unemployment rates,* $u_a \equiv \sum_i \omega_i u_i$, $\bar{u}_a \equiv \sum_i \omega_i \bar{u}_i$, *and f and g are continuous functions of this variance.*

Proof: See Appendix.

This proposition shows that the optimal policy depends only on the first three moments of the distribution of shocks across regions. Because the second and third moments of the distribution of shocks are the same as the second and third moments of the distribution of unemployment rates (around their natural levels), the optimal policy augments that of (11) with functions only of the variance and skew of regional unemployment rates. In addition, we can show

Corollary 1: *a)* $f\left(\sigma_u^2\right) \geq 1$, *with equality when* κ*=0 or* $\sigma_u^2 = 0$.

b) $\dfrac{df\left(\sigma_u^2\right)}{d\sigma_u^2} > 0$.

c) $g\left(\sigma_u^2\right) \leq 0$, *with equality when* κ*=0 or* $\sigma_u^2 = 0$.

d) $\dfrac{dg\left(\sigma_u^2\right)}{d\sigma_u^2} < 0$.

Proof: See Appendix.

The first result in the corollary establishes that if the central bank tries to minimize the differences in aggregate welfare among states or regions, it must respond *more strongly* to the determinants of optimal policy (aggregate shock and expectations of inflation) when the variance of unemployment rates is nonzero. Suppose all states experience a common shock that tends to increase aggregate unemployment with no change in the variance of regional unemployment rates. In the case with $\kappa=0$, the central bank chooses to raise inflation to offset some of the increase in aggregate unemployment, because the loss function is quadratic in both inflation and unemployment. If the cross-sectional variance of unemployment rates is positive, then some states have higher unemployment rates than the average state. These states suffer disproportionately from the increase in unemployment from the shock, again because of the quadratic nature of the loss function. Thus, if $\kappa>0$ so that the central bank has the additional goal of avoiding imposing disproportionate welfare losses on an single region, the central bank must raise inflation more than it would otherwise to accommodate the disproportionate loss suffered by high unemployment states. The second result of the corollary establishes that as the variance of unemployment rates rises, this phenomenon becomes increasingly important as larger fractions of states have disproportionately large welfare losses.

The third result indicates that the coefficient on the skew of the distribution of unemployment rates must be positive (when $\kappa=0$ and the variance of unemployment rates is nonzero). The skew captures the asymmetry of the distribution. When it is positive, there is a fat tail of high unemployment rates, whereas when it is negative there is a fat tail of low unemployment states. States in the fat tails tend to suffer disproportionate welfare losses and therefore have to be accommodated by the central bank when $\kappa>0$. The fourth element indicates that, holding the skew constant, an increase in the variance diminishes the response of optimal policy to the skew of the unemployment distribution.

These results are similar to the theoretical arguments laid out by Dixit (2000) and Fuchs and Lippi (2006). They each consider the problem of an aggregate central bank trying to maximize aggregate welfare subject to the constraint that the regional members find it optimal to stay in the monetary union. They find that the central bank should respond disproportionately to regions for whom the participation constraint is binding. Thus, aggregate policy is affected by regional concerns above and beyond those embodied in aggregate variables. The approach considered here naturally yields a similar conclusion but can be applied to both the ECB and the Fed, whereas the notion of states exiting the monetary union is inapplicable to the US.

We focus on testing two of these implications: the amplification effect of the variance on the response to aggregate variables and the skew of the distribution. To do so, we treat the following as the desired rule when $\kappa=0$

$$i_t^{des} = \phi_\pi \pi_{t+6} + \phi_{ue} ue_{t+6} + v_{t+6} \tag{12}$$

and use the following estimating equation

$$i_t = c + i_t^{des} + \sum_{j=1}^{J} \rho_j i_{t-j} + \beta_1 \sigma_{u,t}^2 i_t^{des} + \beta_2 \sigma_{u,t}^2 + \beta_3 skew\left(u_{i,t} - \bar{u}_{i,t}\right) + \varepsilon_t \tag{13}$$

which allows us to estimate c, ϕ_π, ϕ_{ue}, ρ_j, β_1, β_2, and β_3 jointly. The null is that $\beta_1=\beta_2=\beta_3=0$, whereas the theory predicts that, with $\kappa>0$, we should have $\beta_1>0$ and $\beta_3<0$. Because the theory also implies that the relevant distribution when $\kappa>0$ is that of unemployment gaps for each region, we consider the variance and skew across regions of unemployment rates in levels (as before, which is correct if $\bar{u}_i = \bar{u}$ for all i) and where each region's time-t unemployment rate is relative to the mean unemployment rate of that region over the time sample.[19]

The results are presented in Table 6. For the US, regardless of whether we use demeaned regional unemployment levels, we find no evidence that $\beta_1>0$. The same is true for the ECB. This indicates that the variance of regional unemployment rates does not affect interest rates by amplifying the response to aggregate variables as the dispersion of UE rates rises, as would be predicted by the theory. Instead, the positive level effect of UE dispersion on interest rates continues to hold in almost all cases. A novel result of this table is that the skew of the distribution of UE rates also appears to be an important predictor of interest rates. When we include both the skew and variance measures, we can identify different effects for the ECB and the Fed. In the US, interest rates appear to fall with the skew of regional rates, particularly when each state's unemployment rate is measured relative to that state's average UE rate over the whole sample. This is the sign predicted by the theory, since an increase in the skew means a longer tail of high unemployment states. Because these high unemployment states are suffering disproportionately large losses, the central bank should accommodate them by lowering interest rates. However, we find exactly the opposite result for the ECB, where interest rates rise with the skew of the distribution.

The evidence is thus mostly inconsistent with the theoretical possibility that the central bank has the additional goal of minimizing regional disparities in welfare. There is no evidence that the central bank's response to aggregate variables is amplified when the cross-sectional variance of regional unemployment rates rises. Instead, the level effect of regional dispersion measures is

[19] We present estimates done by GMM rather 2SLS because the interaction terms are highly correlated with the level effects. GMM performs better at identifying the independent contribution of each variable in the presence of such high correlation. We use as instruments six lags of inflation and unemployment, the same lags of the interest rate as included in the RHS of equation (13), plus 3 lags of any dispersion measures and interaction terms included in the regression.

basically unchanged from the baseline: higher dispersion implies higher interest rates, which actually indicate a policy less accommodation for the high unemployment states. In addition, for the ECB, the sign on the central bank's response to the skew of the distribution is contrary to that predicted by the theory. The central bank should lower interest rates when there is a longer tail of high unemployment states, whereas our estimates indicate that the ECB has tended to do the reverse. The one result consistent with the theory is that US interest rate tends to fall with the skew of unemployment, once the variance of unemployment rates is taken into account.

4.2 Do Central Banks Care Equally about All Regions?

Because the evidence supporting the notion that the central bank has an additional goal of minimizing welfare differences across regions receives little support, we consider in this section the possibility that the central bank responds asymmetrically to different regions. Specifically, we focus on whether the ECB and Fed respond to unemployment rates of high and low unemployment states in the same way. Our measures of the dispersion of regional unemployment rates can rise for two reasons. First, high unemployment states can see their unemployment rates rise. Second, low unemployment states may see their unemployment rates fall. Both raise the dispersion, but could lead to different responses by the central bank.

To test this hypothesis, we separate our percentile gap measures into two components

$$\begin{aligned} UEP9010_t &\equiv UEP90_t - UEP10_t \\ &= UEP90_t - UE_t - \left(UEP10_t - UE_t\right) \\ &= UEP90UE_t - UEP10UE_t \end{aligned}$$

where we've defined UEP90UE≡UEP90-UE and equivalently for the 10th percentile of the regional UE distribution. The same procedure can be applied to the 75th-25th percentile gap of UE distribution. Thus, we consider two sources of dispersion: the deviation of high unemployment states from the mean and the deviation of low unemployment states from the mean. We can test the null that the central bank responds symmetrically to both using

$$i_t = \phi_\pi \pi_{t+j} + \phi_{ue} ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \beta_1 UEP90UE_t - \beta_2 UEP10UE_t + \zeta_t \tag{14}$$

The null of no difference between the two groups is simply $\beta_1 = \beta_2$.

Results are presented in Table 7 of estimates done by GMM.[20] For the US, the point estimates of β_1 are consistently not statistically different from zero, whereas those of β_2 are positive and statistically significant. A positive β_2 implies that when low-unemployment states experience even lower unemployment rates relative to the mean, the central bank lowers interest rates controlling for the effect on aggregate unemployment. In addition, we can reject the null that the central bank responds identically to increases in the dispersion of unemployment rates coming from changes in either high or low unemployment states. For the ECB, we can never reject the null that the central bank responds identically to both groups. When using the 75th and 25th percentiles of the regional unemployment distribution, both β_1 and β_2 are positive and significantly different from zero.

The results for the ECB are consistent with the central bank trying to minimize the dispersion of unemployment rates. Conditional on the desired interest rate response to aggregate variables, the central bank lowers interest rates when high-unemployment countries are far from the average unemployment rate and raises interest rates when low-unemployment countries move further away from the average unemployment rate. The US estimates, on the other hand, imply that the central bank only follows such a policy for low-unemployment states. This is somewhat surprising, since one would expect that if central bankers did respond asymmetrically to distribution effects in either high or low unemployment states, it would be to respond disproportionately to high unemployment regions. Our findings, in fact, imply the opposite.

4.3 Do Institutional Features Matter?

The result that the Fed and the ECB appear to respond differently to regional unemployment gaps raises the possibility that the institutional features of each organization could be driving the result. In the ECB, interest rate decisions are made by the Governing Council which consists of the twelve governors of national central banks and six members of the Executive Board, of which four are typically from the "big" countries: Germany, France, Italy, and Spain. A majority vote decides interest rate policy, but no minutes or records of voting patterns are released. In contrast, the Federal Reserve sets interest rates by a majority vote among seven Board members and five (rotating) regional Bank Presidents (Dominguez (2006)).

[20] Again, GMM is used to better identify the independent contribution of highly correlated variables. 2SLS yields the same qualitative results with the only exception that standard errors for the ECB dispersion measures are larger, so that one cannot reject the null that $\beta_1=0$ *or* $\beta_2=0$.

Aksoy et al (2002) argue that if some voting members focus on regional concerns, aggregate interest rate decisions can be sub-optimal because of majority voting. Heinemann and Huefner (2004) test this notion and find weak evidence of such an effect in the ECB. Meade and Sheets (2005) provide evidence that members of the Federal Open Market Committee of the Fed systematically respond to the unemployment rate of their home regions in voting for interest rate changes. If certain regions are over-represented within the decision-making process and members vote in a manner reflecting the state of their regional economy as well as aggregate considerations, then this could lead to interest rate decisions being related to the dispersion of unemployment rates across regions as found in section 2.

We first consider whether the composition of voting members of the FOMC and Governing Council of the ECB are representative of the population shares accounted for by each region. For the US, we use data from Meade and Sheets (2005), which provides the voting decisions of each member of the FOMC for each meeting since 1982 out to 2000. Each voting member is assigned a region of origin, including Board members. Table 8 presents the fraction of votes accounted for by members of each Federal Reserve District. For the ECB, we present a similar breakdown by country from January 1999 to September 2005.

For the US, the Northeast, i.e. New England and New York, appear to be the most heavily over-represented regions in FOMC meetings. Both districts 1 and 2 of the Federal Reserve have been the source of a disproportionate amount of voting Board members relative to their share of the population. In addition, because the New York Fed always has a vote at FOMC meetings, it also accounts for a disproportionate share of voting done by regional presidents. On the other hand, the Southeast (Atlanta-based) and the West (San Francisco-based) are the most under-represented in voting decisions relative to their share of the population. The Southeast is particularly unaccounted for in terms of voting Board members. All other districts have accounted for a share of votes approximately equal to their share of the population. Note that there appears to be no close relationship between voting representation in FOMC meetings and the average difference between regional unemployment and the aggregate unemployment rate.[21]

[21] Because some districts account for parts of states, we had to arbitrarily place some states entirely within some districts and therefore exclude them from districts of which they are partially part of. This was done only for the purposes of calculating unemployment rates per district. See Appendix 2 for a complete description of which district includes which state.

For the ECB, France, Germany, and Italy are heavily underrepresented in the voting decisions of the Governing Council of the ECB, despite each of them consistently occupying a seat on the Executive Board, in addition to their representation via their national central banks. Most dramatically, while Germany accounts for over 25 percent of the Euro-Zone's population, it only accounts for ten percent of votes cast. Instead, the smaller countries, which each receive at least one vote are over-represented relative to their share of the population. Luxembourg, for example, accounts for little over one-tenth of a percent of the population but has cast about six percent of all votes in meetings of the Governing Council. Interestingly, the over-represented "small" countries of the ECB have had much lower unemployment rates than those of France, Germany, and Italy (Finland is the only exception).

To determine whether our finding that dispersion measures of regional unemployment rates affect interest rates is due to the over- and under-representation of regions in central banks' decision-making procedures, we construct a measure of weighted regional unemployment gaps, where the weights are given by the voting representation of each region. For the US, we take the fraction of votes associated with each region on any given meeting of the FOMC and multiply these fractions by the difference between the (one-month lagged) unemployment rates of each region from the (one-month lagged) aggregate unemployment rate. This yields a series with frequency given by Fed meetings. For the ECB, we apply the same procedure using contemporaneous values of unemployment rates. Because the ECB Governing Council meets monthly, this is a monthly series.

We then use the following equation to test whether including these measures eliminates the predictive power of the dispersion measures used in the previous sections

$$i_t = \phi_\pi E_t \pi_{t+j} + \phi_{gy} E_t gy_{t+j} + \phi_{ue} E_t ue_{t+j} + \sum_{i=1}^{T} \rho_i i_{t-i} + \beta_1 D_t + \beta_2 D_t^{voting} + \varepsilon_t \qquad (15)$$

where D_t one of our measures of the cross-sectional dispersion of regional unemployment rates and D_t^{voting} is our new measure of dispersion using voting shares of each region. For the US, we estimate equation (15) using GreenBook forecasts of inflation, output growth, and aggregate unemployment by OLS from 1982:01 to 2000:12 at the frequency of Fed meetings. Both the dispersion measures are lagged one period to ensure orthogonality of RHS variables to error term. For the ECB, we use ex-post values of inflation and aggregate unemployment and estimate (15) by 2SLS, as in section 2, at a monthly frequency.[22]

The results are presented in Table 9. For all three measures of the regional dispersion of unemployment rates used, we find the same result for the ECB and the Fed. Specifically, the coefficient β_2 on the voting-share-weighted dispersion measure is never statistically different from zero. The coefficients on β_1 are always positive and have the same significance levels as found before. The other parameters of the Taylor rule are also unchanged. Thus, it appears that our measures of dispersion are not capturing institutional biases for and against various regions. While regional concerns may have an effect on the individual decisions of voting members, as argued by Meade and Sheets (2005), these appear to have no aggregate effect on interest rates and cannot explain why interest rates appear to systematically respond to the regional dispersion of unemployment rates each period.

5 Conclusion

This paper presents robust, if puzzling, evidence that interest rate decisions of policy-makers are systematically affected by the distribution of unemployment rates across regions. Specifically, the greater is the spread of unemployment rates across regions, the higher interest rates tend to be. We find that this phenomenon appears to hold both for the US Federal Reserve and the European Central Bank. In addition, an appealing feature of augmenting estimates of the Taylor rule with these dispersion measures is that they help more clearly identify the response of the central bank to inflation and unemployment, and make the Fed and ECB monetary policies appear to be more qualitatively consistent with one another.

We show that this result is economically significant and does not appear to be a statistical fluke. The very symmetry of the results across the Fed and the ECB as well as the fact that it holds for various measures of dispersion of regional unemployment rates is also consistent with this result not being an anomaly. Even after controlling for the Fed's expectations via Green Book forecasts, we continue to find a strong predictive role for the lagged cross-state variance of unemployment rates.

This result is surprising because according to theory and the legal foundations of these institutions, regional concerns should not affect interest rates decisions other than by their effect on

[22] We have to use different approaches because Fed meetings are not held monthly and we do not have ECB forecasts of future aggregate variables. As in section 2, we drop output growth (industrial production) for ECB estimates because these have no explanatory power.

macroeconomic aggregate variables. We find little evidence that our results are driven by a concern of the central banks to minimize welfare losses across regions or by the institutional representation of regions in voting decisions. The exception to this finding is that in the US, we find that the central bank appears to respond to deviations from the mean by low-unemployment states, but not high unemployment states.

This finding is puzzling and potentially important for central banks of heterogeneous regions. For example, with the Euro-Zone likely to expand in future years, and an African Central Bank in its formative stages, the cross-sectional variance of unemployment rates is likely to be higher for these institutions. Deviations from an "optimal" monetary policy – or more precisely, a policy that is aimed exclusively at maximizing welfare based on aggregate measures of inflation and unemployment – may increase. Given that responding to such regional variation is not optimal in an aggregate sense, the gains from a single monetary policy-making institution could be lower than expected by both current and future member states.

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Appendix 1

Proof of Proposition 1 and Corollary 1:

Note first that minimization of (8) over inflation yields the optimality condition

$$\sum_i \omega_i \frac{dL_i}{d\pi} + \kappa \sum_i \omega_i \left[L_i - \sum_j \omega_j L_j \right] \left[\frac{dL_i}{d\pi} - \sum_j \omega_j \frac{dL_j}{d\pi} \right] = 0 \tag{A1}$$

Minimization of (6) combined with (7) yields

$$\frac{dL_i}{d\pi} - \sum_j \omega_j \frac{dL_j}{d\pi} = -\alpha\lambda(\varepsilon_i - \varepsilon_a) \tag{A2}$$

and we also have

$$L_i - \sum_j \omega_j L_j = \frac{\lambda}{2}\left(\varepsilon_i^2 - \sum_j \omega_j \varepsilon_j^2 \right) - \alpha\lambda(\pi - \pi^e)(\varepsilon_i - \varepsilon_a) \tag{A3}$$

Substituting (A2), (A3), and equation (9) into (1) and rearranging yields, in terms of the optimal inflation level π^*

$$(1+\alpha^2\lambda)\pi^* - \alpha\lambda\varepsilon_a - \alpha^2\lambda\pi^e = \kappa\alpha\lambda^2 \left[\sum_i \omega_i \left[\frac{(\varepsilon_i - \varepsilon_a)}{2} \left(\varepsilon_i^2 - \sum_j \omega_j \varepsilon_j^2 \right) - \alpha(\pi^* - \pi^e)(\varepsilon_i - \varepsilon_a)^2 \right] \right]$$

which simplifies to

$$(1+\alpha^2\lambda)\pi^* - \alpha\lambda\varepsilon_a - \alpha^2\lambda\pi^e = \kappa\alpha\lambda^2 \left[\frac{1}{2}\left(\sum_i \omega_i \varepsilon_i^3 - \varepsilon_a \sum_i \omega_i \varepsilon_i^2 \right) - \alpha(\pi^* - \pi^e) \sum_i \omega_i (\varepsilon_i - \varepsilon_a)^2 \right] \tag{A5}$$

Now note that the weighted variance of the observed regional shocks is given by

$$\mathrm{var}(\varepsilon_i) = \sum_i \omega_i (\varepsilon_i - \varepsilon_a)^2 = \sum_i \omega_i \varepsilon_i^2 - \varepsilon_a^2 \tag{A6}$$

and the weighted skew of these observed shocks is

$$\begin{aligned} skew(\varepsilon_i) &= \sum_i \omega_i (\varepsilon_i - \varepsilon_a)^3 = \sum_i \omega_i \varepsilon_i^3 - 3\varepsilon_a \sum_i \omega_i \varepsilon_i^2 + 2\varepsilon_a^3 \\ &= \sum_i \omega_i \varepsilon_i^3 - \varepsilon_a^3 - 3\varepsilon_a \,\mathrm{var}(\varepsilon_i) \end{aligned} \tag{A7}$$

where the last equality makes use of (A6).

Substituting both (A6) and (A7) into equation (A5) yields

$$(1+\alpha^2\lambda + \alpha^2\lambda\kappa\,\mathrm{var}(\varepsilon_i))\pi^* = \alpha\lambda(\varepsilon_a + \alpha\pi^e) + \frac{\kappa\alpha\lambda^2}{2}\left[skew(\varepsilon_i) + 2\varepsilon_a \,\mathrm{var}(\varepsilon_i) \right] + \alpha^2\lambda^2\kappa\,\mathrm{var}(\varepsilon_i)\pi^e \tag{A8}$$

Defining $\Psi = 1 + \kappa\lambda\,\mathrm{var}(\varepsilon_i)$, we can rewrite (A8) as

$$(1+\alpha^2\lambda\Psi)\pi^* = \Psi\alpha\lambda(\varepsilon_a + \alpha\pi^e) + \frac{\alpha\lambda^2\kappa}{2} skew(\varepsilon_i) \tag{A9}$$

Note that $\sigma_u^2 \equiv \sum_i \omega_i \left[(u_i - \bar{u}_i) - (u_a - \bar{u}_a) \right]^2 = \sum_i \omega_i [\varepsilon_i - \varepsilon_a]^2 \equiv \mathrm{var}(\varepsilon_i)$

and $skew(u_i - \bar{u}_i) \equiv \sum_i \omega_i \left[(u_i - \bar{u}_i) - (u_a - \bar{u}_a) \right]^3 = \sum_i \omega_i [\varepsilon_i - \varepsilon_a]^3 \equiv skew(\varepsilon_i)$

then defining $f\left(\sigma_u^2\right) \equiv \dfrac{\Psi + \alpha^2 \lambda \Psi}{1 + \alpha^2 \lambda \Psi}$ and $g\left(\sigma_u^2\right) \equiv \dfrac{\alpha \lambda^2 \kappa}{2\left(1 + \alpha^2 \lambda \Psi\right)}$ yields

$$\pi^* = \pi^{opt} f\left(\sigma_u^2\right) + skew\left(u_i - \bar{u}_i\right) g\left(\sigma_u^2\right)$$

Note that $\Psi = 1 + \kappa\lambda \operatorname{var}\left(\varepsilon_i\right) \geq 1 \Rightarrow f\left(\sigma_u^2\right) \geq 1$ and $g\left(\sigma_u^2\right) \geq 0$.

Note also that

$$\begin{aligned}
\frac{df}{d\sigma_u^2} &= \left[\frac{1+\alpha^2\lambda}{1+\alpha^2\lambda\Psi} - \frac{\Psi\alpha^2\lambda\left(1+\alpha^2\lambda\right)}{\left(1+\alpha^2\lambda\Psi\right)^2}\right]\frac{d\Psi}{d\sigma_u^2} \\
&= \left[\frac{1+\alpha^2\lambda}{1+\alpha^2\lambda\Psi}\right]\frac{d\Psi}{d\sigma_u^2} > 0
\end{aligned}$$

while $\dfrac{dg}{d\sigma_u^2} < 0$ since $\dfrac{d\Psi}{d\sigma_u^2} > 0$.

Appendix 2: Data Details for Voting Measures

A- States associated with each Fed District

District 1: Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont.
District 2: New York
District 3: Pennsylvania, Delaware, New Jersey.
District 4: Ohio.
District 5: Washington DC, Maryland, Virginia, North and South Carolina, West Virginia.
District 6: Alabama, Florida, Georgia, Louisiana, Mississippi, and Tennessee.
District 7: Iowa, Illinois, Indiana, Michigan, and Wisconsin.
District 8: Arkansas, Missouri, and Kentucky.
District 9: Minnesota, Montana, North and South Dakota.
District 10: Colorado, Kansas, Nebraska, Oklahoma, Wyoming, and New Mexico.
District 11: Texas.
District 12: California, Arizona, Utah, Nevada, Oregon, Washington, Idaho, Alaska, and Hawaii.
Because districts typically include parts of states, this division only approximately captures the division of states across districts. It was necessary to divide districts into states because employment data by month is only available at the state level.

B- ECB members of the Governing Council.

Every nation has one representative through the head of its central bank. Greece joined in January 2001. In addition, the following were members of the Executive Board and had voting rights in interest rate decisions:

President: Duisenberg (ND) from Jan. 1999 to Oct. 2003. Replaced by Trichet (FR) in Nov. 2003, to present.

Vice-President: Noyer (FR) from June 1998 to May 2002. Replaced by Papademos (GR) in June 2002, to present.

Members:

Solans (ES) from June 1998 to May 2004. Replaced by Gonzalez-Paramo (ES) June 2004 to present.
Hamalainen (FI) from June 1998 to May 2003. Replaced by Tumpel-Gugerell (AU) June 2003 to present.
Issing (DE) from June 1998 to May 2006. Replaced by Stark (DE) June 2006 to present.
Padoa-Schioppa from June 1998 to May 2005. Replaced by Bini Smagghi (IT) June 2005 to present.

Table 1: Does Regional Variation in UE Affect Interest Rates?

Panel A: United States				
	Baseline	***(1)***	***(2)***	***(3)***
c	-0.08	-0.04	-0.09	0.21
	(0.17)	(0.15)	(0.16)	(0.19)
ϕ_{π}	0.10**	0.09***	0.11***	0.11***
	(0.05)	(0.03)	(0.04)	(0.04)
ϕ_{ue}	-0.01	-0.07**	-0.04	-0.09***
	(0.02)	(0.03)	(0.03)	(0.03)
ρ_1	1.31***	1.26***	1.26***	1.24***
	(0.08)	(0.08)	(0.08)	(0.08)
ρ_2	-0.34***	-0.30***	-0.30***	-0.29***
	(0.08)	(0.07)	(0.08)	(0.07)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β		0.10***	0.12**	0.11***
		(0.03)	(0.05)	(0.03)
Sample	January 1982 - September 2005			

Panel B: Euro-Zone				
	Baseline	***(1)***	***(2)***	***(3)***
c	2.42***	0.67	1.29**	1.42*
	(0.95)	(0.89)	(0.61)	(0.75)
ϕ_{π}	0.01	0.24***	0.16***	0.19*
	(0.06)	(0.07)	(0.06)	(0.10)
ϕ_{ue}	-0.26***	-0.15*	-0.20***	-0.20***
	(0.10)	(0.08)	(0.07)	(0.07)
ρ_1	0.92***	0.88***	0.89***	0.88***
	(0.04)	(0.03)	(0.03)	(0.03)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β_1		0.07***	0.08***	0.05**
		(0.02)	(0.03)	(0.02)
Sample	January 1999 - September 2005			

Note: All estimates done by 2SLS with Newey-West standard errors in parentheses. Instruments include 6 lags of each endogenous variable (inflation, unemployment, and additional variables when included). Dependent variables are interest rates, while ϕ_{π}, ϕ_{ue}, and ρ_i are coefficients on 6-month ahead inflation, 6-month ahead unemployment, and i lags of the interest rate respectively. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Table 2: Decomposing Variance of the Endogenous Component of Interest Rates

Panel A: United States

	Fraction of Var(Endog) due to		
	var(agg)	***var(D)***	***cov(agg,D)***
Dispersion Measure			
UEP90-UEP10	74%	93%	-67%
UEP75-UEP25	57%	51%	-7%
var(UE)	87%	114%	-101%

Panel B: Euro-Zone

	Fraction of Var(Endog) due to		
	var(agg)	***var(D)***	***cov(agg,D)***
Dispersion Measure			
UEP90-UEP10	64%	39%	-3%
UEP75-UEP25	66%	27%	6%
var(UE)	66%	34%	0%

Note: The table presents decompositions of the variance of the endogenous component of interest rates as defined in equation (5), but normalized by the variance of the endogenous interest rate component. Var(*agg*), var(*D*), and cov(*agg,D*) are the variance of the endogenous component of interest rates due to aggregate inflation and unemployment, the variance of the measure of regional unemployment dispersion, and the covariance of the two respectively. Each expression is multiplied by relevant constants from equation (5). The estimated coefficients used come from the results of Table 1.

Table 3: Granger Causality Tests

	F-Statistic	*p-value*
United States		
π does not Granger-Cause var(UE)	1.68*	0.07
var(UE) does not Granger-Cause π	1.47	0.13
UE does not Granger-Cause var(UE)	1.48	0.13
var(UE) does not Granger-Cause UE	0.92	0.53
Euro-Zone		
π does not Granger-Cause var(UE)	2.75***	0.005
var(UE) does not Granger-Cause π	0.88	0.57
UE does not Granger-Cause var(UE)	0.81	0.64
var(UE) does not Granger-Cause UE	2.07**	0.03

Note: Granger Causality tests done with 12 lags. The var(UE) series is the weighted variance of regional unemployment rates each month. Data is from 1982:01 to 2005:09 for US and 1999:01 to 2005:09 for Euro-Zone. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Table 4: Including Leading Indicators

Panel A: United States				
Added Variable:	***Consumer Confidence***	***Stock Prices***	***Oil Prices***	***PPI Inflation***
c	-3.88*	-0.17	0.26	0.16
	(2.12)	(1.29)	(0.24)	(0.19)
ϕ_{π}	0.14***	0.12**	0.11***	0.16***
	(0.03)	(0.05)	(0.04)	(0.05)
ϕ_{ue}	0.00	-0.08**	-0.08***	-0.09**
	(0.05)	(0.04)	(0.03)	(0.04)
ρ_1	1.23***	1.24***	1.25***	1.23***
	(0.08)	(0.07)	(0.07)	(0.08)
ρ_2	-0.28***	-0.29***	-0.30***	-0.29***
	(0.08)	(0.07)	(0.07)	(0.07)
β_1	0.07**	0.13***	0.11***	0.11***
	(0.03)	(0.04)	(0.03)	(0.04)
β_2	0.81*	0.04	-0.02	-0.01
	(0.43)	(0.11)	(0.05)	(0.01)
Sample	January 1982 - September 2005			

Panel B: Euro-Zone				
Added Variable:	***Consumer Confidence***	***Stock Prices***	***Oil Prices***	***PPI Inflation***
c	0.12	1.02	0.18	0.73
	(0.52)	(2.21)	(1.05)	(0.52)
ϕ_{π}	0.17***	0.21**	0.14*	0.15
	(0.06)	(0.08)	(0.07)	(0.10)
ϕ_{ue}	0.01	-0.18	-0.12	-0.12**
	(0.05)	(0.12)	(0.08)	(0.05)
ρ_1	0.89***	0.88***	0.91***	0.91***
	(0.02)	(0.03)	(0.03)	(0.03)
β_1	-0.01	0.05**	0.07***	0.03
	(0.02)	(0.03)	(0.03)	(0.02)
β_2	0.02***	0.02	0.15	0.02**
	(0.01)	(0.16)	(0.11)	(0.01)
Sample	January 1999 - September 2005			

Note: All estimates done by 2SLS with Newey-West HAC standard errors in parentheses. Instruments include 6 lags of each endogenous variable (inflation, unemployment, the weighted variance of regional unemployment rates and the additional variable included). Dependent variables are interest rates, while ϕ_{π}, ϕ_{ue}, and ρ_i are coefficients on 6-month ahead inflation, 6-month ahead unemployment, and i lags of the interest rate respectively. β_1 and β_2 are the coefficients on the dispersion measure and the additional variable respectively. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Table 5: Using Green-Book Forecasts

	Panel A: United States			
	Baseline	***(1)***	***(2)***	***(3)***
c	0.09	0.19	0.19	0.85**
	(0.25)	(0.25)	(0.29)	(0.40)
ϕ_{π}	0.31***	0.32***	0.31***	0.34***
	(0.10)	(0.10)	(0.10)	(0.10)
ϕ_{gy}	0.18***	0.14***	0.15***	0.12**
	(0.05)	(0.05)	(0.05)	(0.05)
ϕ_{ue}	-0.11***	-0.20***	-0.15**	-0.27***
	(0.04)	(0.07)	(0.06)	(0.08)
ρ	0.88***	0.85***	0.86***	0.84***
	(0.04)	(0.05)	(0.05)	(0.05)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β		0.18*	0.13	0.23***
		(0.09)	(0.09)	(0.08)
Sample	Fed Meetings from Jan. 1982 - Dec. 2000			

Note: All estimates done by OLS with Newey-West HAC standard errors in parentheses. Dependent variable is the target FFR chosen at each meeting, while ϕ_{π}, ϕ_{gy}, *and* ϕ_{ue} are coefficients on Green-Book forecasts of average inflation, output growth, and unemployment over current quarter through next two quarters. ρ is the coefficient on the target FFR from the previous meeting. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Table 6: Testing whether Central Banks Minimize Dispersion of Regional Losses

	Panel A: United States						Panel B: Euro-Zone					
	Regional UE in Levels			*Demeaned Regional UE*			*Regional UE in Levels*			*Demeaned Regional UE*		
	(1)	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*
c	0.13*	0.10	0.30***	0.08	0.09	0.05	0.37	1.63**	0.25	3.58***	2.48***	1.89***
	(0.07)	(0.11)	(0.10)	(0.06)	(0.12)	(0.08)	(0.62)	(0.64)	(0.62)	(0.75)	(0.56)	(0.73)
ϕ_{π}	0.06***	0.08***	0.10***	0.06***	0.07***	0.10***	0.01	0.18***	0.00	-0.05	-0.05	-0.02
	(0.02)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.06)	(0.01)	(0.04)	(0.05)	(0.02)
ϕ_{ue}	-0.04**	-0.04**	-0.11***	-0.04**	-0.03**	-0.07***	-0.04	-0.22***	-0.02	-0.37***	-0.25***	-0.18**
	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)	(0.02)	(0.08)	(0.06)	(0.08)	(0.07)	(0.06)	(0.08)
ρ_1	1.28***	1.27***	1.25***	1.28***	1.30***	1.19***	0.92***	0.89***	0.93***	0.88***	0.90***	0.90***
	(0.06)	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)
ρ_2	-0.32***	-0.30***	-0.29***	-0.31***	-0.32***	-0.24***						
	(0.05)	(0.06)	(0.05)	(0.05)	(0.06)	(0.06)						
β_1	0.25		-0.04	0.63		0.10	0.87		2.02	-0.18		0.82
	(0.19)		(0.05)	(0.48)		(0.14)	(2.52)		(11.25)	(0.12)		(0.65)
β_2	0.10***		0.16***	0.17**		0.36***	0.28		0.34*	-0.62		1.32**
	(0.04)		(0.03)	(0.08)		(0.06)	(0.18)		(0.17)	(0.44)		(0.58)
β_3		0.04***	-0.04*		0.05**	-0.12***		-0.03***	0.03***		0.19***	0.27***
		(0.01)	(0.02)		(0.02)	(0.03)		(0.01)	(0.01)		(0.04)	(0.07)
Sample	Jan 1982 - Sep 2005			Jan 1982 - Sep 2005			Jan 1999 - Sep 2005			Jan 1999 - Sep 2005		

Note: The table presents estimates of equation (13) in the text. Estimates are done by GMM with New-West weighting matrix, with a truncation of 6 lags. The dependent variable is the interest rate. ϕ_{π}, ϕ_{ue}, and ρ_j (for j=1 or 2) are the responses of the central bank to expected inflation, expected unemployment and lag j of the interest rate respectively. β_1, β_2, and β_3 are the responses to the interaction term, the level of the variance of regional unemployment rates each month, and the skew of unemployment rates each month. We allow for the variance and skew measures to be taken across regional unemployment rates as reported (in levels) and across regional demeaned unemployment rates. Statistical significance at the 1%, 5%, and 10% levels are indicated by ***, **, and * respectively. Standard errors of estimates are in parentheses below coefficients.

Table 7: Decomposing the Dispersion of Regional UE Rates

	United States		Euro-Zone	
UE Percentiles:	***90th and 10th***	***75th and 25th***	***90th and 10th***	***75th and 25th***
c	0.08	-0.06	0.80**	1.17***
	(0.06)	(0.07)	(0.37)	(0.28)
ϕ_π	0.08***	0.11***	0.24***	0.18***
	(0.02)	(0.02)	(0.03)	(0.03)
ϕ_{ue}	-0.10***	-0.11***	-0.17***	-0.20***
	(0.02)	(0.02)	(0.03)	(0.03)
ρ_1	1.27***	1.23***	0.88***	0.89***
	(0.05)	(0.05)	(0.01)	(0.01)
ρ_2	-0.30***	-0.25***		
	(0.05)	(0.05)		
β_1	0.02	-0.10	0.03	0.07**
	(0.05)	(0.07)	(0.03)	(0.03)
β_2	0.18***	0.37***	0.09***	0.09***
	(0.05)	(0.08)	(0.02)	(0.01)
Wald ($\beta_1 = \beta_2$)	3.32*	11.15***	1.93	0.17
Sample	Jan 1982 - Sep 2005		Jan 1999 - Sep 2005	

Note: This table presents estimates of equation (14) in text. β_1 is the coefficient on the difference between the 90th or 75th percentile of the regional unemployment distribution and the mean unemployment rate. β_2 is the same using the 10th or 25th percentiles. Estimates done by GMM with Newey-West weighting matrix (6 lags). Standard errors in parentheses below coefficients. Wald is the Wald test statistic of the restriction that $\beta_1 = \beta_2$. Statistical significance at the 1%, 5%, and 10% levels are indicated by ***, **, and * respectively.

Table 8: Regional Representation in Interest-Rate Decision-Making

Panel A: US Federal Reserve Bank Districts

District:	***1***	***2***	***3***	***4***	***5***	***6***	***7***	***8***	***9***	***10***	***11***	***12***
Share of Board Member Votes	0.12	0.15	0.09	0.00	0.14	0.00	0.13	0.07	0.00	0.10	0.11	0.08
Share of Regional Pres. Votes	0.06	0.20	0.07	0.10	0.07	0.07	0.10	0.06	0.06	0.07	0.07	0.07
Share of Total Votes	0.10	0.17	0.08	0.04	0.11	0.03	0.12	0.07	0.03	0.08	0.09	0.08
Share of Population	0.05	0.07	0.08	0.04	0.09	0.14	0.13	0.04	0.03	0.05	0.07	0.19
Mean UE gap	-1.1	0.1	-0.1	0.4	-0.7	0.2	0.3	0.1	-1.4	-0.9	0.3	0.5

Panel B: Members of European Central Bank

Country:	***AU***	***BE***	***FI***	***FR***	***DE***	***GR***	***IR***	***IT***	***LX***	***ND***	***PR***	***ES***
Share of Total Votes	0.08	0.06	0.09	0.10	0.11	0.10	0.06	0.11	0.06	0.10	0.06	0.11
Share of Population	0.03	0.03	0.02	0.19	0.27	0.04	0.01	0.19	0.00	0.05	0.03	0.13
Mean UE gap	-4.3	-0.8	0.7	0.8	-0.5	1.8	-4.0	0.5	-5.2	-5.1	-3.2	2.4

Note: US Federal Reserve Bank Districts are based in Boston (1), New York (2), Philadelphia (3), Cleveland (4), Richmond (5), Atlanta (6), Chicago (7), St. Louis (8), Minneapolis (9), Kansas City (10), Dallas (11), and San Francisco (12). Members of the ECB are Austria (AU), Belgium (BE), Finland (FI), France (FR), Germany (DE), Greece (GR, since Jan. 1 2001), Ireland (IR), Italy (IT), Luxembourg (LX), Netherlands (ND), Portugal (PR), and Spain (ES). Data for FOMC votes is from Meade and Sheets (2005) and contains votes by Board members and Regional Presidents. For the ECB, votes are of members of the Governing Council, which include Executive Board members and Presidents of each national central bank. Mean UE gaps are average difference between regional unemployment rates and the aggregate unemployment rates, from 1982:01 to 2005:09 for US and from 1999:01 to 2005:09 for Euro-Zone.

Table 9: Does Regional Representation in Interest-Rate Decisions Matter?

	Panel A: United States			Panel B: Euro-Zone		
	(1)	*(2)*	*(3)*	*(1)*	*(2)*	*(3)*
c	0.25	0.28	0.90**	1.12	1.30**	1.25**
	(0.32)	(0.34)	(0.45)	(0.73)	(0.58)	(0.61)
ϕ_{π}	0.32***	0.32***	0.34***	0.14***	0.12***	0.16***
	(0.10)	(0.10)	(0.11)	(0.05)	(0.04)	(0.05)
ϕ_{gy}	0.13**	0.14**	0.12**			
	(0.06)	(0.05)	(0.06)			
ϕ_{ue}	-0.21**	-0.16**	-0.28***	-0.17***	-0.19***	-0.17***
	(0.08)	(0.07)	(0.09)	(0.07)	(0.06)	(0.06)
ρ	0.85***	0.86***	0.84***	0.89***	0.89***	0.89***
	(0.05)	(0.05)	(0.05)	(0.03)	(0.03)	(0.03)
β_1	0.17**	0.10	0.22***	0.06**	0.08**	0.05***
	(0.07)	(0.09)	(0.06)	(0.03)	(0.03)	(0.02)
β_2	-0.14	-0.23	-0.19	0.04	0.08	0.07
	(0.37)	(0.43)	(0.41)	(0.13)	(0.13)	(0.12)
Sample	Jan 1982: Sep 2005 (FOMC meetings)			Jan 1999: Sep 2005 (monthly)		

Note: Estimates for the US are done by OLS using GreenBook Forecasts of future inflation (ϕ_{π}), output growth (ϕ_{gy}), and unemployment (ϕ_{ue}) with interest rates measured by the target FFR on data with frequency of FOMC meetings. Estimates for ECB are done by 2SLS with 6-month ahead values of inflation and unemployment with interest rates measured by interbank overnight rate. β_1 is the coefficient on each measure of the cross-sectional regional dispersion of unemployment rates. β_2 is the coefficient on the weighted sum of differences between regional and aggregate unemployment rates, where the weights are the voting share of each region in the interest-rate decision process that period. All standard errors are Newey-West HAC. Statistical significance at the 1%, 5%, and 10% levels are indicated by ***, **, and * respectively.

Figure 1: US Aggregate Variables

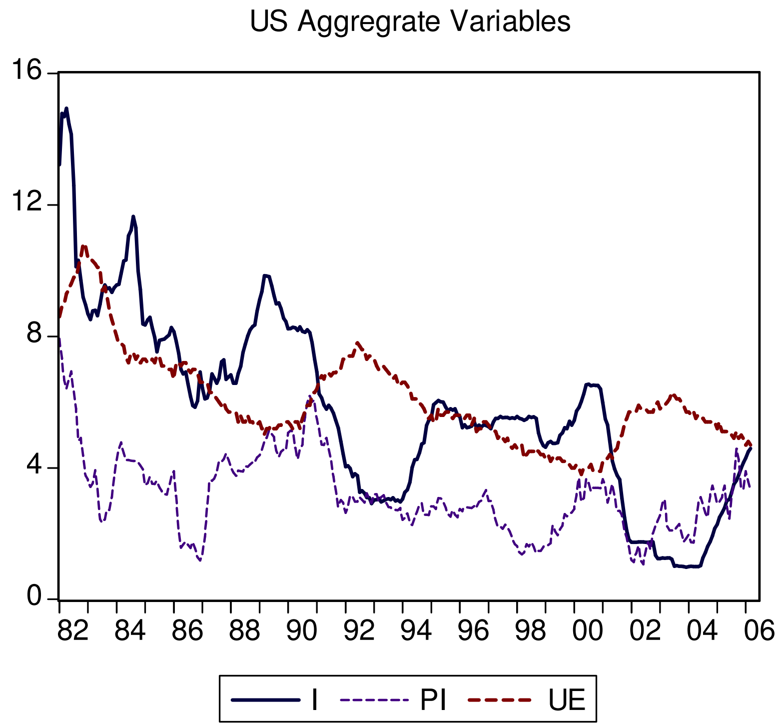


Figure 2: Measures of the Regional Dispersion of Unemployment Rates

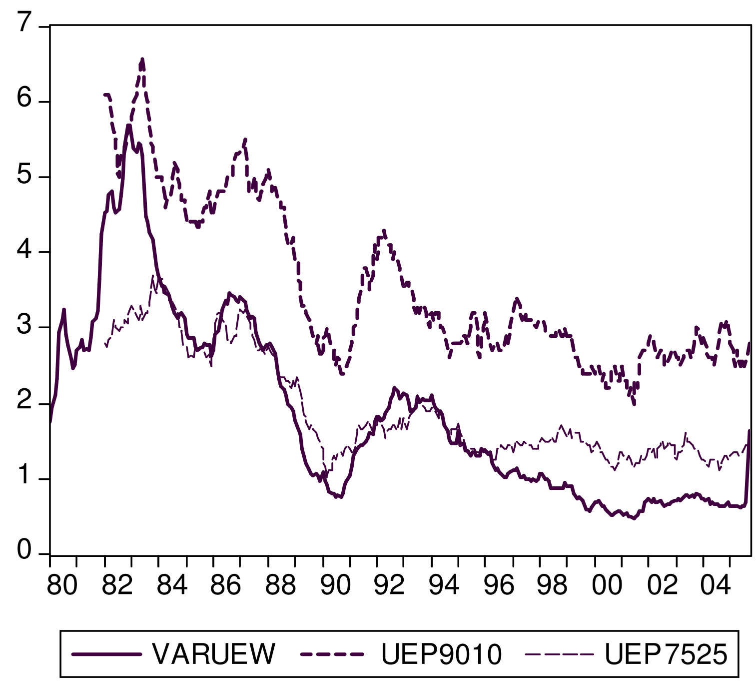


Figure 3: European Aggregate Variables

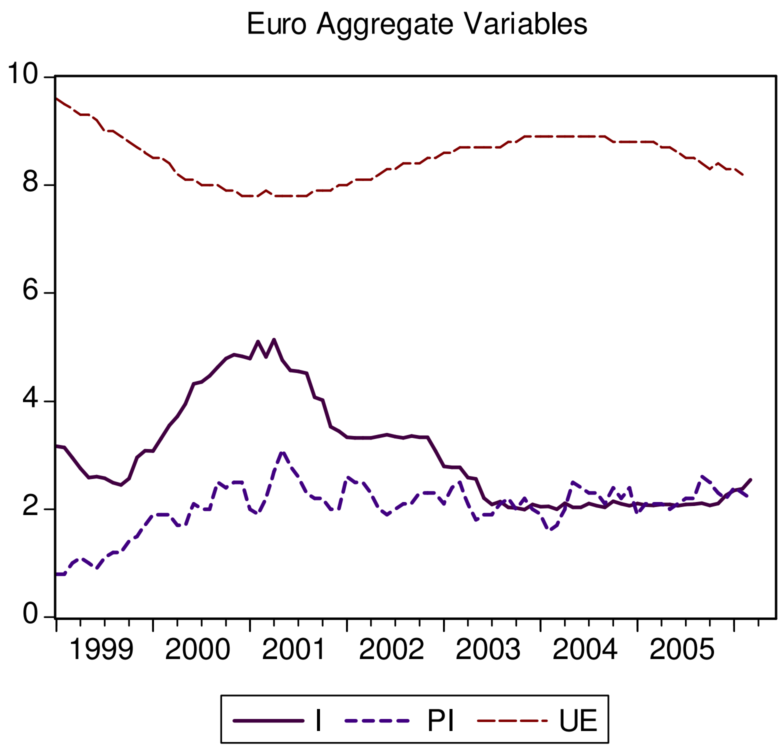


Figure 4: Measures of Regional Dispersion of Unemployment Rates in Euro-Area

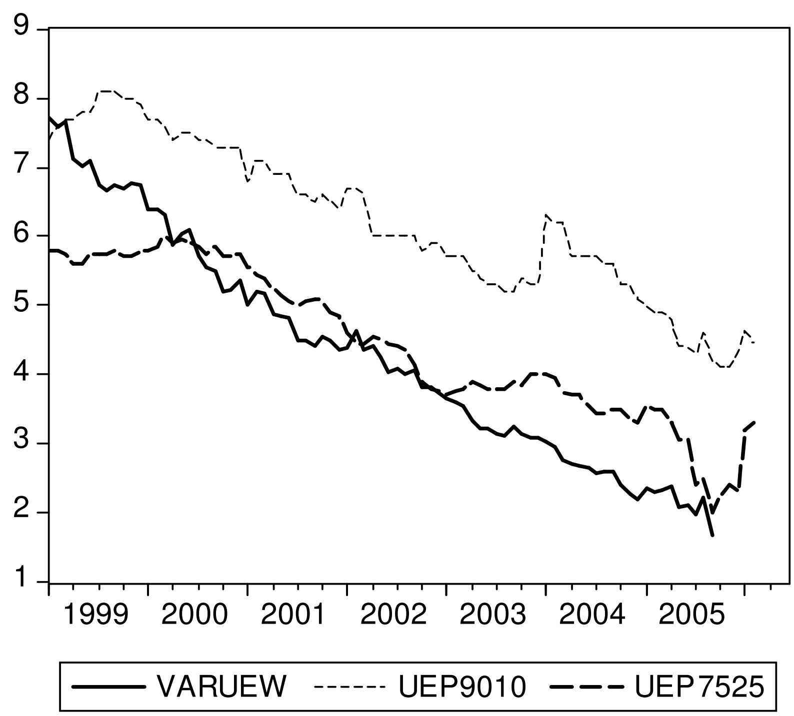


Figure 5: Scatter Plot of Variance of US state UE rates against Orthogonalized Component of Interest Rates

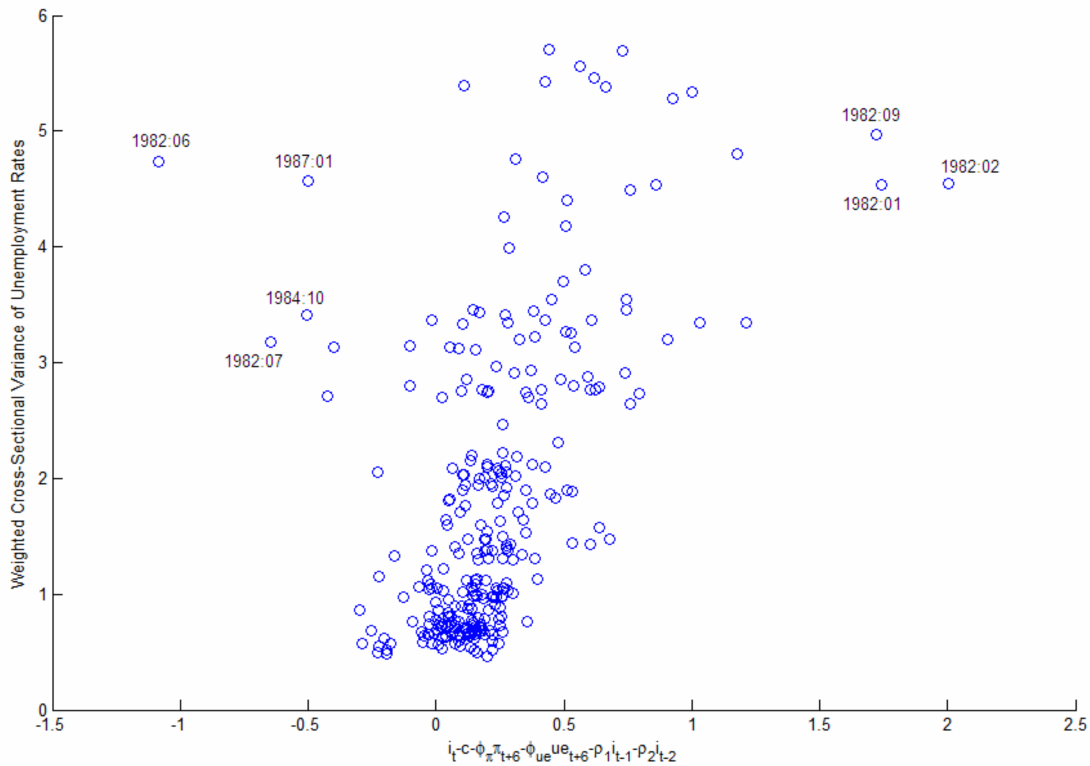
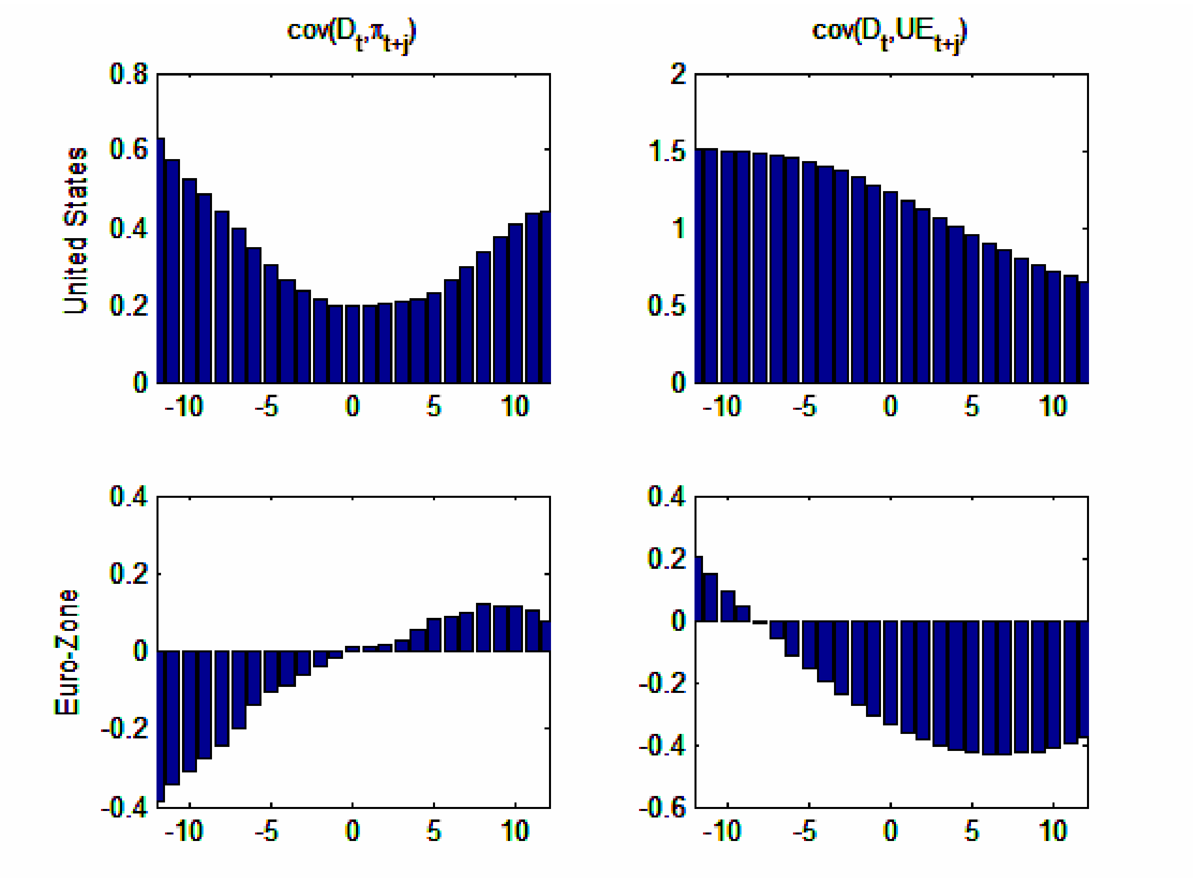


Figure 6: Dynamic Cross-Correlations of Dispersion of Regional Unemployment Rates with Aggregate Variables



Note: This figure plots the dynamic cross-correlations of the weighted variance of unemployment rates across regions with leads (positive values on x-axis) and lags of aggregate inflation and unemployment rates for the US from 1982:01-2005:09 and Euro-Zone 1999:01-2005:09. All data is monthly.

Appendix Table 1: Robustness to Time Sample

	Panel A: United States			
	Baseline	***(1)***	***(2)***	***(3)***
c	0.21	-0.49	0.33***	0.62***
	(0.19)	(0.58)	(0.10)	(0.19)
ϕ_π	0.11***	0.24**	0.06***	0.09***
	(0.04)	(0.11)	(0.02)	(0.03)
ϕ_{ue}	-0.09***	-0.03	-0.09***	-0.16***
	(0.03)	(0.05)	(0.02)	(0.03)
ρ_1	1.24***	1.21***	1.35***	1.15***
	(0.08)	(0.08)	(0.09)	(0.10)
ρ_2	-0.29***	-0.29***	-0.38***	-0.20**
	(0.07)	(0.08)	(0.08)	(0.09)
β	0.11***	0.14***	0.12***	0.22***
	(0.03)	(0.04)	(0.03)	(0.05)
Sample	82:1-05:9	82:1-91:12	87:1-05:9	92:1-05:9

Note: All estimates done by 2SLS with Newey-West HAC standard errors in parentheses. Instruments include 6 lags of each endogenous variable (inflation, unemployment, and additional variables when included). Dependent variables are interest rates, while ϕ_π, ϕ_{ue}, and ρ_i are coefficients on 6-month ahead inflation, 6-month ahead unemployment, and i lags of the interest rate respectively. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Appendix Table 2: GMM Estimates

	Panel A: United States			
	Baseline	***(1)***	***(2)***	***(3)***
c	0.00	0.03	-0.02	0.33***
	(0.09)	(0.07)	(0.08)	(0.09)
ϕ_{π}	0.08***	0.08***	0.10***	0.08***
	(0.03)	(0.02)	(0.02)	(0.02)
ϕ_{ue}	-0.02*	-0.07***	-0.05***	-0.10***
	(0.01)	(0.01)	(0.01)	(0.02)
ρ_1	1.32***	1.26***	1.27***	1.25***
	(0.06)	(0.05)	(0.05)	(0.05)
ρ_2	-0.34***	-0.30***	-0.31***	-0.30***
	(0.06)	(0.05)	(0.05)	(0.04)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β		0.10***	0.13***	0.12***
		(0.02)	(0.02)	(0.02)
Sample	January 1982 - September 2005			

	Panel B: Euro-Zone			
	Baseline	***(1)***	***(2)***	***(3)***
c	2.09***	0.26	1.23***	0.98**
	(0.76)	(0.42)	(0.37)	(0.44)
ϕ_{π}	0.02	0.27***	0.19***	0.23***
	(0.04)	(0.03)	(0.04)	(0.04)
ϕ_{ue}	-0.23***	-0.12***	-0.20***	-0.16***
	(0.08)	(0.04)	(0.04)	(0.04)
ρ_1	0.93***	0.88***	0.89***	0.88***
	(0.03)	(0.01)	(0.01)	(0.01)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β_1		0.08***	0.08***	0.06***
		(0.01)	(0.01)	(0.01)
Sample	January 1999 - September 2005			

Note: All estimates done by GMM with Newey-West weighting matrix. Instruments include 6 lags of each endogenous variable (inflation, unemployment, and additional variables when included). Dependent variables are interest rates, while ϕ_{π}, ϕ_{ue}, and ρ_i are coefficients on 6-month ahead inflation, 6-month ahead unemployment, and i lags of the interest rate respectively. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.

Appendix Table 3: Using Target Interest Rates

Panel A: United States				
	Baseline	***(1)***	***(2)***	***(3)***
c	-0.08	-0.02	-0.08	0.23
	(0.17)	(0.15)	(0.16)	(0.19)
ϕ_{π}	0.10**	0.09***	0.12***	0.12***
	(0.05)	(0.03)	(0.04)	(0.04)
ϕ_{ue}	-0.01	-0.07**	-0.05	-0.09**
	(0.02)	(0.03)	(0.03)	(0.04)
ρ_1	1.31***	1.20***	1.20***	1.17**
	(0.08)	(0.09)	(0.10)	(0.10)
ρ_2	-0.34***	-0.25***	-0.25**	-0.23**
	(0.08)	(0.09)	(0.10)	(0.10)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β		0.11***	0.12**	0.12***
		(0.03)	(0.05)	(0.04)
Sample	January 1982 - September 2005			

Panel B: Euro-Zone				
	Baseline	***(1)***	***(2)***	***(3)***
c	2.42***	1.78	1.97**	1.92**
	(0.95)	(1.14)	(0.82)	(0.88)
ϕ_{π}	0.01	0.12	0.07	0.11
	(0.06)	(0.09)	(0.08)	(0.08)
ϕ_{ue}	-0.26***	-0.24**	-0.26***	-0.23***
	(0.10)	(0.11)	(0.09)	(0.09)
ρ_1	0.92***	0.87***	0.87***	0.87***
	(0.04)	(0.04)	(0.04)	(0.03)
Variable Added:		*UEP90-EUP10*	*UEP75-EUP25*	*var(UE)*
β_1		0.06**	0.08**	0.05**
		(0.03)	(0.03)	(0.02)
Sample	January 1999 - September 2005			

Note: All estimates done by 2SLS with Newey-West HAC standard errors in parentheses. Instruments include 6 lags of each endogenous variable (inflation, unemployment, and additional variables when included). Dependent variables are central banks' target interest rates, while ϕ_{π}, ϕ_{ue}, and ρ_i are coefficients on 6-month ahead inflation, 6-month ahead unemployment, and i lags of the interest rate respectively. Statistical significance at the 1%, 5%, and 10% levels are indicated by a ***, **, and * respectively.