SM-87-1

of the Research Department for Review and Comment

A Series of Occasional Papers in Draft Form Prepared by Members

STAFF MEMORANDA

WAGE GROWTH AND SECTORAL SHIFTS: NEW EVIDENCE ON THE STABILITY OF THE PHILLIPS CURVE

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ABSTRACT

Several researchers have noted a worsening trade-off between inflation and unemployment over the seventies. This paper argues that the apparent shift in the parameters of the Phillips curve is illusory. Standard Phillips curve analyses employ the unemployment rate or some function of the unemployment rate as a proxy for general economic conditions. The unemployment rate is only a good measure of the overall state of demand as long as the underlying structure of the economy is stable. This paper provides an alternative measure of general economic conditions by attempting to filter out the effects of sectoral shifts on the unemployment rate. Structural realignment in the seventies, measured as permanent changes in the distribution of employment across industries, appears to account for much of the higher unemployment rates observed.

Wage Growth and Sectoral Shifts: New Evidence on the Stability of the Phillips Curve

Ellen R. Rissman*

Advocates of the Phillips curve approach to modeling the inflation process typically relate wage or price inflation to some measure of current economic activity and an inflationary expectations variable. Economic activity is usually proxied by the unemployment rate or some function of the unemployment rate, and expected inflation is specified as a function of lagged values of actual inflation. This traditional approach to the inflation process appears to have dwindled in popularity in recent years as many researchers have suggested that, most noticeably over the seventies, the parameters of the Phillips curve have changed. Cagan (1975), Sachs (1980), and O'Brien (1985) argue that the trade-off between inflation and unemployment has worsened in that a given decline in the rate of wage inflation requires a larger increase in the rate of unemployment than previously needed, i.e. the short run Phillips curve has become flatter. In contrast, Wachter (1976) and Schultze (1981) suggest that the opposite is instead the case so that wage inflation has actually become more cyclically sensitive, implying a steeper short run Phillips curve. Although disagreement exists over the direction of change in the slope of the curve, most empirical work suggests that the parameters have indeed shifted.

Attempts to resurrect the Phillips curve have typically fallen along two lines of reasoning. First, unstable parameters may be the result of an initial misspecification of the inflationary expectations process. Thus, several researchers have experimented with alternative expectations formulations and more sophisticated models of wage and price dynamics. However, as Sachs (1980) notes, the evidence indicates that more complicated models of expected price inflation do not account for what appears to be a significant change in the short run trade-off between inflation and unemployment.

An alternative tack that has been taken is to assume that the inflationary expectations and economic activity variables in the regression models are correctly specified while other relevant explanatory variables have been

*The author would like to thank Bob DeFina, Craig Hakkio, Gary Koppenhaver, and Steve Strongin for helpful comments and suggestions, and Eric Klusman for excellent research assistance. The author alone is responsible for any remaining errors. The views expressed here are not necessarily those of the Federal Reserve Bank of Chicago or the Federal Reserve System.

omitted from the analysis. When these other factors are considered, the apparent instability of the parameters is explained. Several researchers, including Gordon (1977), Perry (1980), Hamilton (1983), and Loungani (1985), have suggested that demographic changes combined with the oil price shocks, the acceleration and termination of the Vietnam War, and the implementation of wage and price controls account for much of the observed change in the parameters.

A somewhat different explanation that is not as reliant on the timing of historical events is that the perceived instability is the result of a change in the underlying structure of the inflation process. For example, Sachs (1980), Barro (1977), and Taylor (1980) among others argue that an increase in the use of longer term labor contracts and a belief held by the public that monetary and fiscal policy will be used so as to promote high employment and stable prices have served to worsen the trade-off. Although there is some theoretical justification for this hypothesis, empirical evidence on this point is lacking.

Oddly, in light of recent research on sectoral shifts and their relation to the unemployment rate, the search for alternative specifications of the Phillips curve has led to relatively little consideration of the appropriate measure of economic activity to be employed in the analysis.[1] Typically, the unemployment rate or some function of the unemployment rate is used to capture the disequilibrium aspect although some researchers, including Sachs (1980) and Gordon (1977), have chosen instead to use alternative measures such as deviations of real Gross National Product from potential. These variables are all meant to measure the current level of economic activity.

Discussions of the Phillips curve are typically expressed in terms of the labor market and labor market tightness. Controlling for inflationary expectations, the negative coefficient on the unemployment rate in a regression model of wage inflation is interpreted as indicating that tight labor markets, as signaled by lower unemployment rates, put upward pressure on wages and therefore on prices through some sort of markup equation. According to the sectoral shifts hypothesis of unemployment [Lilien (1982a)], changes in the distribution of demand, given its level, cause the unemployment rate to increase as it is costly both for the employer and employees to adjust instantaneously. Therefore, the sectoral shifts hypothesis suggests that higher unemployment is not necessarily indicative of a weakening economy, while similarly lower unemployment need not signify stronger aggregate demand.

If the relation between the actual unemployment rate and the underlying state of the economy has changed over time, then it is not surprising that the parameters of the Phillips curve as typically specified have shifted. Such a shift would occur, for example, if the volume of structural unemployment, i.e. that unemployment which is generated by compositional shifts in the structure of demand, were to change. This suggests that one possible way of explaining the apparent instability of the Phillips curve in recent years is to adjust the unemployment rate by filtering out the effects of sectoral shifts so as to obtain a more meaningful measure of labor market activity.

The purpose of this paper is to analyze the effects of permanent sectoral shifts on the wage inflation process. The remainder consists of three sections. In Section I a measure of sectoral shifts proposed by Neumann and Topel (1984) is used to construct estimates of the natural rate of unemployment defined in Lilien (1982a) as the rate of unemployment that would occur in the absence of cyclical fluctuations. The measure of changes in the industrial composition of employment used in the analysis is constructed by filtering changes in employment shares for the ten major industry groupings into temporary and permanent components. It is argued that, although the timing of the various post-World War II business cycles coincides with marked permanent shifts in the distribution of employment across industries, there is sufficient variation in the employment response over the cycles of the various industries examined to suggest that these changes in employment distribution are not driven solely by the business cycle, in contrast to Abraham and Katz (1986).

The effect of sectoral shifts on wage inflation is examined in Section II. If the wage inflation process depends upon the general level of economic activity rather than the distribution of demand across industries, then the difference between the actual rate of unemployment and the constructed natural rate series provides a truer measure of the overall state of the economy. The difference between the two reflects only cyclical fluctuations by construction provided that the measure of sectoral shifts employed is truly independent of the cycle. This suggests that standard Phillips curve type regressions are misspecified in that they include the level of the unemployment rate as an explanatory variable rather than the difference between the actual unemployment rate and the natural. In a less restrictive form, the actual unemployment rate should enter negatively while the estimate of the natural rate enters positively in a regression model of wage inflation. Assuming that the Phillips curve in its correct specification depends upon the general level of economic activity, acceptance of the hypothesis of equal and opposite signs on the coefficients of the actual unemployment rate and the natural rate adds some additional, although indirect, support for the sectoral shifts theory as it suggests that the decomposition of unemployment into its underlying cyclical and structural components provides meaningful economic content. Section III contains conclusions and suggestions for further research.

Section I: Sectoral Shifts and the Natural Rate of Unemployment

The sectoral shifts hypothesis examined by Lilien (1982a, 1982b) argues that changes in the composition of labor demand, rather than simply the level, cause unemployment in the short run due to the existence of adjustment costs. In testing the hypothesis a measure of sectoral shifts is constructed and then related in a regression model to the unemployment rate to test whether sectoral shifts add any significant explanatory power. Such a procedure is an appropriate test if the measure of sectoral shifts adopted is independent of the level of general economic activity. However, it is well-established that business cycles have a differential impact across industries. [See Mitchell (1941).] For this reason, measures of sectoral shifts constructed from observations of employment shares over time are not in general independent of the business cycle.

This problem of interpreting distributional changes in employment has received recent attention by several researchers. Lilien (1982b) and Davis (1986a) provide evidence in support of the sectoral shifts hypothesis by attempting to filter out the effects of the business cycle on the composition of employment. In a similar manner Neumann and Topel (1984) and Loungani (1985) suggest that changes in the distribution of employment that are derivative of the business cycle are temporary whereas structural shifts are more permanent in nature. This distinction enables one to construct a measure of permanent compositional changes that are assumed to be independent of the cycle for use in testing the sectoral shifts hypothesis. In contrast, Abraham and Katz (1986), using vacancy data, find that distributional changes in employment are driven in large part by the business cycle rather than exogenous shifts in resource allocation.

Although resolution of this debate is not imminent, some insight may be gained by examining the behavior of specific industries over the business cycle. If changes in the distribution of employment across industries are fundamentally related to the business cycle rather than some sort of structural shift, then a given industry's employment response should be approximately the same over the various cycles. Table 1 provides some evidence on this point for the ten two-digit durable manufacturing industries over the eight post-World War II business cycles. Durable manufacturing industries are analyzed due to the common perception that these industries are more cyclically sensitive than others.

Define Δe as the change in employment in durable manufacturing from business cycle peak to subsequent trough. Similarly, define Δe_i as the change in employment between peak and trough for the i^{th} two-digit industry. Thus, $\Delta e_i/\Delta e$ measures the fraction of the change in durable manufacturing employment occurring over the business cycle attributable to the i^{th} industry. If each industry were a constant proportion of total nondurable manufacturing employment at the peak of each business cycle, then $\Delta e_i/\Delta e$ should be approximately the same magnitude for each of the eight cycles. However, the measure must be adjusted to control for compositional changes in durable manufacturing employment over time. For example, suppose that the employment share of industry i is decreasing over time. Because the i^{th} industry accounts for a much smaller share of total employment in later years, it should also account for a much smaller share of total changes in employment over the more recent business cycles.

The numbers reported in Table 1 have been adjusted for changes in employment shares by subtracting the fraction of durable manufacturing employment in industry i at the peak of the cycle from $\Delta e_i/\Delta e$. If the industry responds in the same manner over the various cycles, then the numbers reported should be roughly the same within an industry. From Table 1 it is apparent that not all business cycles are alike with respect to the composition of employment. For example, between the fourth quarter of 1948 and the fourth quarter of 1949 the primary metals industry accounted for 30 percent of the change in durable manufacturing employment even though its employment share at the peak of the cycle was only 15 percent. In contrast, primary metals contributed only 7 percent to the change in total employment from 1973 to 1975 although its employment share in the fourth quarter of 1973 was approximately 11 percent. By comparison, transportation equipment accounted for only 12 percent of the total change in durable manufacturing employment between 1948 and 1949 but 37 percent between 1969 and 1970 even though its employment share grew from only 16 to 18 percent between the two business cycle peaks.

If the business cycle hypothesis is correct, then the response of a given industry should be the same across cycles. Thus, the cross-cycle variance of the observations should be relatively small. Standard tests of variance based upon the X^2 distribution are inappropriate because of the small sample size and, more fundamentally, a problem of determining how small the variance should be under the null hypothesis. A somewhat weaker variant of the business cycle hypothesis is that the underlying distributions of the observations are the same across cycles. A likelihood ratio test of the equality of the variances of the observations across cycles can be easily constructed from the sample variances within a cycle. The computed test statistic, $-2\log\lambda$, equals 12.607 and is asymptotically distributed as X^2 with (8-1) degrees of freedom.[2] The probability value associated with the calculated statistic is 0.08. However, the small sample properties of the test statistic are unknown.

Although some industries may fare relatively better than others in terms of employment over the course of the business cycle, there is sufficient ation in the employment response of a given industry across cycles to suggest that changes in the sectoral composition of employment over time are driven by more than simply general cyclical fluctuations. While a cyclical interpretation is not precluded, any cyclical explanation must necessarily explain the differing character of these cycles.[3]

To attempt to filter the effects of cyclical disturbances from allocative disturbances on the distribution of employment across industries, Neumann and Topel (1984) have suggested that cyclical shocks are temporary whereas sectoral disturbances are more permanent in nature stemming from some fundamental long term change in the structure of the economy.[4] Accordingly, they present a method for measuring permanent and temporary changes in the distribution of employment across industries. Essentially the computation involves a decomposition of deviations in employment shares from trend, measured as a weighted average of past employment shares, into two components—one predictable from observations on future employment shares and one which is orthogonal to it. The series are calculated using employment data for the ten major industry categories which are reported in the BLS' *Employment and Earnings*.[5] The measures of permanent and temporary changes in the industrial composition of employment, Δ^P and Δ^T respectively, are shown in Figure 1. The series have been constructed in the manner described in Neumann and Topel (1984) and extended through the first quarter of 1981.[6]

From Figure 1, the measure of permanent changes in the distribution of employment across industries is much more variable than the measure of temporary change. The measure of permanent shifts in the industrial composition of employment exhibits quite noticeable peaks in 1958, 1961, 1970, and 1975, and possibly in 1980. If Δ^P truly measures sectoral or structural shifts, then it appears that these events are intimately related to the timing of the business cycle. While it is possible to interpret this evidence as implying that distributional changes are driven by the cycle, an equally valid interpretation is that structural shifts cause the cycle. Temporary changes in the employment distribution do not appear to be significantly correlated with the general state of the economy.

The correct interpretation of permanent sectoral shifts and their relation to the cycle is left unresolved. However, the variability of the employment response in durable manufacturing over the cycle suggests that the relation between the two is not simply one of cause and effect. Furthermore, although the timing of contractions and sectoral shifts is coincidental, strong contractions do not necessarily occur when permanent employment adjustments are most pronounced. The NBER ranks the downturns of 1969-1970 and 1973-1975 respectively as mild and strong contractions. However, sectoral shifts were more pronounced in the earlier period than in the latter. In the analysis that follows, an agnostic approach is taken to the debate over the exogeneity of permanent sectoral shifts. An alternative question is asked: Does the decomposition of the unemployment rate into a 'cyclical' and a 'natural' or 'structural' component provide useful economic information?

Sectoral Shifts and the Unemployment Rate

According to the sectoral shifts theory of unemployment, permanent changes in the distribution of employment across industries should increase the unemployment rate in the short run as both employers and employees adapt to the changing conditions. [See Lilien (1982b) and Davis (1986b) for a discussion of the adjustment to allocative disturbances.] Therefore, in a regression model of the unemployment rate the measure Δ^P should enter with a positive coefficient.

Various other factors also affect the unemployment rate.[7] The demographic composition of the labor force has been cited as one cause of the rise in the unemployment rate over the seventies as a larger proportion of women, nonwhites, and younger workers entered the labor market. Because empirically the effect of nonwhites and females on the unemployment rate is much smaller in magnitude than the effect of the age composition, the dependent variable employed in the analysis that follows (WUR_t) is constructed assuming a fixed age composition of the labor force given by its distribution in the first quarter of 1960.

Aside from sectoral shifts and demographics, the business cycle also has an effect upon the unemployment rate. Following Lilien (1982a), this cyclical effect is assumed to be captured by 'unanticipated money growth' (DM_t). As defined in Barro (1978), unanticipated money growth is constructed as the least squares residuals from a regression of M1 on a vector of explanatory variables. Deviations of real Gross National Product from trend ($DGNP_t$) are also assumed to capture the business cycle effect on unemployment where the explanatory variable is constructed as the estimated residuals from an ARIMA model of real GNP.

Finally, unemployment insurance and other social welfare programs are thought to have an effect upon the unemployment rate. [See Mortensen (1977) for an analysis of the effect of unemployment insurance on job search.] The effect of such programs on the unemployment rate is assumed to be measured by social insurance expenditures expressed as a percentage of Gross National Product (SI_t). The data are reported annually in the Social Security Administration's *Social Security Bulletin: Annual Statistical Supplement*. Linear interpolation was used to obtain quarterly figures. Because social insurance expenditures typically rise during economic downturns, the ratio of such expenditures to GNP is cyclical. Thus, SI may simply proxy for additional cyclical information independent of any hypothesized effect it may have upon job search or job matching.

The regression results are reported below in equation [1] where the associated t-statistics are given in parentheses below the corresponding coefficient estimates. The regression was computed over the period from 1954 through the third quarter of 1981 using ordinary least squares.

$$
\begin{aligned}
WUR_t = \underset{(0.178)}{0.018} &- \underset{(3.678)}{0.009DGNP_t} - \underset{(3.198)}{0.008DGNP_{t-1}} - \underset{(0.623)}{0.002DGNP_{t-2}} \\
&- \underset{(2.324)}{7.740DM_t} + \underset{(0.883)}{3.083DM_{t-1}} - \underset{(0.149)}{0.563DM_{t-2}} + \underset{(6.410)}{94.448\Delta_t^P} \qquad [1] \\
&- \underset{(5.871)}{86.213\Delta_{t-1}^P} + \underset{(3.211)}{3.332SI_t} + \underset{(15.156)}{1.185WUR_{t-1}} - \underset{(3.182)}{0.247WUR_{t-2}}
\end{aligned}
$$

$R^2 = 0.978$.

Preliminary analysis unreported here revealed that inclusion of lagged dependent variables increased the performance of the equation. The OLS results are asymptotically equivalent to maximum likelihood estimates provided that the errors are homoskedastic. The Ljung-Box statistic testing for autocorrelation of the estimated residuals for a lag length of six quarters has a computed value of 3.30 and is distributed X^2 . Small magnitudes of the test statistic suggest that the residuals are serially uncorrelated.

From the regression results, the initial impact of a given permanent change in the distribution of employment across industries is to increase the unemployment rate.[8] However, the effect damps quickly so that within four quarters the unemployment rate is negligibly different from its predisturbance level. Furthermore, positive deviations in real GNP from trend have a negative effect upon the unemployment rate, affecting future unemployment both directly via lags of the assumed exogenous variables and indirectly via the lagged endogenous variables. In accord with Rush (1985), positive values of monetary disturbances are associated with significant contemporaneous declines in the unemployment rate. The standard errors on additional lags are large relative to the coefficient estimates so that lagged direct effects of monetary disturbances, as opposed to indirect effects working through lags of the dependent variable, are negligible.

Social insurance expenditures enter positively into the regression model. One possible interpretation is that more generous insurance benefits prolong or possibly promote unemployment. As noted above, however, the measure employed, social insurance expenditures as a percentage of GNP, is in theory procyclical. Therefore, rather than truly capturing the disincentive effect of welfare insurance on job search, the variable may simply be an additional indicator of the general level of economic activity. Although this is possible, the actual series exhibits no clearly cyclical pattern. In constructing the natural rate of unemployment it is assumed that SI represents other institutional factors affecting the unemployment rate independent of cyclical unemployment.

Construction of the Natural Rate of Unemployment

Following Lilien (1982a), the natural rate of unemployment is defined as the rate of unemployment that would occur in the absence of cyclical fluctuations. Thus, the natural rate can be thought of as measuring the amount of unemployment attributable to frictional and structural factors alone. Given this definition, it is a relatively straightforward exercise to calculate the natural rate from the regression results previously obtained in equation [1]. The natural rate in this context is simply calculated as the rate of unemployment that occurs when the cyclical variables, namely DGNP and DM, are set identically equal to zero over the entire time period analyzed. Therefore, the natural rate changes over time due to sectoral shifts in the composition of employment and, to a lesser extent, changes in social insurance expenditures relative to GNP. To implement these computations, it is necessary to specify initial values of the natural rate. The effect of these initial values on the calculations decreases rapidly. As a result, within two years the natural rate is virtually independent of these assumed initial values.

Figure 2 presents the actual age-weighted unemployment rate and the estimate of the natural rate of unemployment calculated from the parameter values of equation [1]. Initial values of the natural rate are taken to be the actual values of the age-weighted unemployment rate for the first and second quarters of 1954. The figure shows the estimates over the period from 1958 through the third quarter of 1981 so as to minimize the influence of the assumption of the specific initial values. It should be noted that because the dependent variable in the regression analysis, WUR, is a demographically fixed-weight unemployment rate, the estimate of the natural rate constructed is also independent of the demographic effects of age.

As seen from the graph, the natural rate of unemployment has at times been below the actual unemployment rate and at other times has been above it. Until late 1966 the natural rate was consistently below the actual by as much as two percentage points. From late 1966 through 1973 the reverse occurred although the natural rate never exceeded the actual by more than one percentage point. The rise in the natural rate over this time is due predominantly to the relatively large amount of sectoral shifts that occurred as measured by Δ^P and, to a lesser extent, the increase in social insurance expenditures as a percentage of GNP. From 1974 through 1977 the actual rate again exceeded the natural rate while for the brief period from 1978 to 1981 the opposite was true.

Generally, it is presumed that low unemployment rates signal a healthy economy while high unemployment rates indicate low aggregate demand. Such an interpretation is valid only if other factors aside from cyclical ones are stable components of the unemployment rate since, in this case, movements in the unemployment rate correspond to changes only in the cyclical component of unemployment. The estimates of the natural rate constructed indicate that the importance of such other factors has varied considerably over the time period examined. The natural rate has been as high as 7.01 percent in the third quarter of 1981 and as low as 3.48 percent in the first quarter of 1966.

The changing level of unemployment, attributable in large part to sectoral shifts, suggests that the unemployment rate provides a noisy signal of general economic conditions at best with even less information conveyed in times of rapid and pronounced allocative disturbances. This idea has important implications for monetary and fiscal policy. First, a countercyclical policy may prove to be elusive as it requires knowledge of the current level of economic activity, which may be difficult to ascertain from currently used economic indicators such as the unemployment rate. Secondly, if monetary and fiscal policy are neutral in their effects upon the distribution of employment, then countercyclical policy is effective only in reducing cyclical and not structural unemployment. Finally, if expansionary policy is undertaken in times of low cyclical unemployment and high structural unemployment, then it is likely to lead to inflation rather than reduction in the unemployment rate.

Section II: Wage Inflation and the Natural Rate of Unemployment

Traditional approaches to modeling the inflation process describe the rate of change of wages or alternative price variable in terms of its equilibrium and disequilibrium components. Thus, it is hypothesized that the rate of growth of wages, w_t, is a function of expected price inflation and the difference between labor demand and labor supply. If labor demand exceeds labor supply, then wages grow more rapidly whereas if the opposite occurs, then wages fall or grow more slowly.

Assuming that wage inflation is a linear function of these variables, then wage growth is expressed as:

$$\dot{w}_t = \delta_0 + \delta_1(L_t^d - L_t^s) + \delta_2\dot{p}_t^e \qquad [2]$$

where $\dot{w}_t$ is the logarithmic rate of change of wages at time t, and L_t^d and L_t^s are respectively labor demand and labor supply at time t, $\dot{p}_t$ is the inflation rate at time t, and the superscript 'e' indicates that the variable is in expectations form. The parameter δ_1 is hypothesized to be positive.

In order to estimate the above Phillips curve relation, many researchers have assumed the labor market conditions variable, $L_t^d - L_t^s$, to be proxied by the actual unemployment rate. Theorists suggest that as labor market conditions become tighter, the wage response is larger. Therefore, the unemployment rate is usually entered in inverse form to attempt to capture this nonlinearity.

Implementation of equation [2] also requires that the expectations process be specified. Various forms have been investigated in the literature. Typically it is assumed that expected inflation depends upon lagged realizations of actual inflation so that $\dot{p}_t^e = b(L)\dot{p}_t$ where $b(L)$ is a polynomial in the lag operator L. In the analysis that follows it is assumed that expected inflation is described by a second order polynomial distributed lag model so that the coefficient on the i^{th} lag, b_i, is expressed as $(a_0 + a_1 i + a_2 i^2)$. The inclusion of beginning and endpoint constraints changes this three parameter model to a one parameter model. Finally, for purposes of estimation it is assumed that an additive error term is included at the end of equation [2].

Any measure of expected price inflation that may be used in the regression analysis is measured with error. It is well known that such a classical errors in variables problem biases the OLS parameter estimates. Specifically, the coefficient on the expected inflation variable will be biased towards zero with the magnitude depending upon the variance of the measurement error relative to the true series. Furthermore, the other parameters of the model are also biased with the direction depending upon the variance-covariance matrix of the observations. [See Levi (1973) for details.] Without prior knowledge of the variance of the errors, it is difficult to correct for the problem. However, the estimates of the coefficient on expected inflation are surprisingly robust to alternative specifications. Therefore, it is hoped that the biases introduced are small.

Table 2 provides OLS parameter estimates and associated t-statistics for the wage inflation model described above using quarterly observations from the second quarter of 1960 through the third quarter of 1981. The dependent variable is the difference in logarithms of average hourly earnings for production workers in manufacturing and nonsupervisory workers in nonmanufacturing. The labor market activity variable is the civilian unemployment rate, UR, entered linearly. Price inflation is calculated as the difference in logarithms of the Consumer Price Index for urban workers. The coefficient on $\dot{p}^e$ reported is the estimate of $\delta_2 a_2$ in an eight quarter polynomial distributed lag on past inflation with beginning and endpoint constraints.[9] Thus, a negative sign on the coefficient estimate indicates that the b_i are all positive and concave in i assuming that $\delta_2 > 0$.

The coefficient estimates of the standard specification are found in column (1). The sign of the coefficient on the unemployment rate is negative and clearly significant at traditional confidence levels implying that increases in labor market tightness are associated with more rapidly rising wages. In addition, higher expected price inflation is associated with higher wage growth as indicated by the parameter estimate on $\dot{p}^e$.[10] Although the coefficients are of the hypothesized signs, the relatively low Durbin-Watson statistic suggests that some underlying factors have not been properly included in the analysis and that the estimated standard errors are incorrect.

Various alternative specifications were tried to investigate the idea that the wage inflation process was somehow different in more recent years than in the earlier part of the sample. The results are presented in columns (2) through (4) of the table. A dummy variable D is defined equaling 1 after the second quarter of 1971, when wage and price controls were introduced, and 0 otherwise. The regression reported in column (2) tests if the trade-off between wage inflation and unemployment worsened in recent years, holding all other parameters constant over the entire time period, by checking the significance of the interaction term DxUR. The results indicate that, in accord with Cagan (1975) and Sachs (1980), the coefficient on UR rose over the seventies and eighties relative to the sixties suggesting that the trade-off did indeed worsen.

The regression reported in column (3) tests whether a change in the effect of expected inflation on wage growth occurred over the latter part of the sample, assuming the other parameters to be constant over the entire period. The insignificance of the coefficient estimate on Dx$\dot{p}^e$ suggests that the effect of expected inflation remained stable over the sample period. However, when both the coefficient on the unemployment rate and the expected inflation variable were permitted to vary simultaneously over time, both the coefficients on the unemployment rate and the expected inflation rate were significantly larger in the latter period than in the former as seen in column (4). The rise in the coefficient reported on the expected inflation rate is not readily interpretable due to the parameterization of the polynomial distributed lag employed. However, assuming that the parameters of the inflationary expectations process are stable over the time period investigated, the rise in the coefficient can be interpreted as a decrease in δ_2 in recent years.

The results reported in Table 2 suggest, as other researchers have noted, that the Phillips curve parameters did indeed shift over the seventies. Various other specifications were tried testing whether the shift in the parameters could be explained by the wage and price controls of the early seventies, changes in the demographic and union composition of the labor force, and changes in the competitive position of the economy as measured by the ratio of imports to Gross National Product. Although inclusion of these variables helped explain some of the apparent instability of the Phillips curve, a substantial amount was still unaccounted.

As suggested in the introduction and discussed in the previous section, the unemployment rate is determined by various other factors aside from the purely cyclical that may be in large part independent of the level of general economic activity. Thus, the Phillips curve as estimated in Table 2 may well be misspecified if the unemployment rate is not a good indicator of the performance of the economy. A more appropriate measure of labor market tightness is one that filters out the effects of such factors as sectoral shifts. The natural rate is constructed as the unemployment rate that would occur in the absence of cyclical fluctuations. Therefore, a direct way of measuring the general state of the economy is to calculate the difference between the actual unemployment rate and the natural rate, a measure which by construction reflects only cyclical variations provided that the measure of sectoral shifts is independent of the cycle.

The regression results reported in Table 3 include the age-weighted unemployment rate and the natural rate of unemployment as separate explanatory variables.[11] The unemployment rate and the natural rate have been included in unconstrained form so that a test of the hypothesis that the estimated coefficients are of equal and opposite signs may be performed. If the hypothesis is accepted, then it suggests that the difference between the actual rate and the constructed natural rate is a good indicator of economic performance, as would be expected if the sectoral shifts variable employed was independent of the cycle. Therefore, a finding of equal and opposite signs of the estimated coefficients on the two unemployment rates is consistent with a sectoral interpretation.[12]

From column (1) of the table it appears that the natural rate provides additional information independent of the level of unemployment that is relevant to the modeling of the wage inflation process. As the natural rate rises relative to the actual, cyclical unemployment becomes a less important component of the total unemployment rate. In other words, increases in the natural rate relative to the actual signal a tightening of labor market conditions due to an improvement in the general performance of the economy. As labor markets become tighter, wages rise more quickly. Similarly, given the level of the natural rate, increases in the actual unemployment rate are associated with a weakening economy and, therefore, wage inflation is lower. The low value of the F-statistic testing for equal and opposite signs of the coefficients of the actual unemployment rate and the natural rate suggests that general economic conditions are best measured by deviations of the actual from the natural rate. A significant improvement in both the R^2 and Durbin-Watson statistic occurs when the standard specification of the Phillips curve is modified to account for structural change.[13]

Inclusion of the natural rate series constructed clearly adds to the performance of the estimating equation. If the sectoral shifts hypothesis is correct, the estimate of the coefficient on the unemployment rate in a simple model of wage inflation should increase when the natural rate has risen relative to the actual. Thus, the higher estimated effect of unemployment on wage inflation found in later years may be explained by the pronounced sectoral changes occurring over that period. When the model is modified to take into consideration the effect of these sectoral shifts on the unemployment rate, no parameter instability should result.

The parameter instability of the modified Phillips curve is examined in the regression results found in columns (2) through (4). As in the preceding analysis, a dummy variable D is defined taking on the value of 1 for the period from the third quarter of 1971 through the end of the sample period, and 0 otherwise. From column (2), once compositional changes in employment are included via the natural rate series, there is no evidence that the trade-off between inflation and unemployment has worsened in recent years given that the other parameters, namely the coefficient on inflationary expectations and the intercept, have been constrained to be constant throughout the sample period.

Stability of the coefficient estimate on expected inflation, assuming constant parameters on the other variables, is tested in column (3). The results suggest that over the earlier part of the sample, inflation and inflationary expectations were not important determinants of the wage inflation process. In fact, over this period the parameter estimate on $\dot{p}^e$ is insignificantly different from zero. However, the more recent period shows a significant positive effect of expected inflation on wage growth. One explanation may be that inflationary expectations were unimportant in earlier years because there was comparatively little variation in inflation. Thus, the effect of inflation over the first part of the sample may be picked up in large part by the constant term.

In column (4) the coefficients on both the labor market tightness variables and the expected inflation variable are permitted to vary over the sample period. Due to multicollinearity, most of the coefficient estimates are insignificant with the exception of the intercept term and expected inflation. Unreported regression results permitting the intercept to vary over the time span also indicate no significant change in the parameters of the Phillips curve.

Although wage growth appears to respond relatively more to price inflation in the latter period of the sample, the multicollinearity problems found in regression (4) make the proper interpretation difficult. For this reason the preferred model appears to be the one estimated initially in column (1). This basic regression was reestimated with the added constraint that the coefficients on the unemployment rate and the natural rate are of equal and opposite signs. The results are reported in column (5). There is remarkably little difference in the parameter estimates and the overall performance of the equation between the constrained and unconstrained cases. In summary, once sectoral shifts have been included in the analysis, there is very little support for the claim that wage growth has become less cyclically sensitive. It would appear that much of the debate about the stagflation of the seventies is in large part attributable to a failure to distinguish among the sources of the underlying disturbances to the economy.

The coefficient estimates of expected inflation and the actual and natural rates of unemployment are surprisingly robust to alternative specifications. In addition to modifying the Phillips curve specification by concentrating on an alternative measure of labor market tightness, several other variables were also examined. Gordon (1977) and Perry (1980) suggest that the changing demographic composition of the labor force has had an effect upon wage growth. Specifically, as the proportion of traditionally low-wage earners such as females, nonwhites, and youths rises, wage growth declines.

The effects of such demographic changes on wage inflation are analyzed in the regression results reported in Table 4, columns (1) through (4). The demographic variables included are respectively the fraction of the labor force that is female, nonwhite, and under age 25 calculated as a one quarter deviation from an eight-quarter moving average trend with geometrically declining weights. None of these three demographic variables is either independently or jointly significant. In fact the sign of the estimated coefficient on Nonwhite is of the opposite sign from that which is hypothesized.

In addition to standard demographic variables, wages and wage growth are thought to be related to the ability of unions to increase wages above the competitive equilibrium level. The decline in union membership in recent years has been well-documented.[14] If declining membership implies that wages are permitted to return to their lower equilibrium levels, then the coefficient on union membership in a regression equation relating unionism to wage growth should be positive. As suggested in Lawrence and Lawrence (1985), however, the decline in union membership may trigger wage increases rather than decreases as the remaining membership attempts to maximize rents in the face of a permanent decline in product demand. Thus, the coefficient on unionism in a wage growth regression may be negative.

Union power is proxied by the fraction of nonagricultural employment belonging to a union calculated as a one quarter deviation from an eight-quarter moving average trend with geometrically declining weights.[15] The results reported in column (5) of Table 4 show an insignificant effect of union membership changes on wage growth.

Finally, the effect of wage and price controls on wage growth was examined by creating a dummy variable (WPC) equaling 1 when wage and price controls were in effect, and 0 otherwise. The regression results reported in column (6) show an insignificant response of wage growth to the wage and price controls of the early seventies.

The failure of other variables to have a significant effect upon wage growth once sectoral shifts have been accounted for is rather puzzling since other researchers, such as Perry (1980), have found these other effects to be significant in studying wage dynamics. One possible explanation is that changes in these other factors may cause or be related to adjustments in the distribution of employment across industries. Therefore, the measure of the natural rate may proxy for those other factors affecting wage growth. For example, the decline in union membership is predominantly a phenomenon encountered in manufacturing industries and other traditional union strongholds. These industries have all been undergoing a long term decline in employment share. Thus, it is quite possible that the shift in employment captured by the measure of sectoral redistribution and the decline in union membership are simply two different aspects of the same phenomenon. Therefore, it is perhaps unsurprising that union membership should exert little influence on wage growth once sectoral shifts have been included in the analysis.

Section IV: Conclusions

The worsening of the trade-off between inflation and unemployment in recent years is illusory. Standard specifications of the Phillips curve employ the unemployment rate as a proxy for the general level of economic activity. However, the large amount of sectoral realignment occurring in the seventies challenges the traditional linkage of unemployment to economic activity. Thus, in times of structural adjustment the unemployment rate is not a good indicator of economic conditions associated with the business cycle.

High unemployment may occur even though aggregate demand is relatively high since sectoral shifts force people to adjust and, in some instances, endure a spell of unemployment. Once the unemployment rate has been adjusted to filter out the effects of compositional changes in the structure of employment, a better measure of the general level of economic activity results that suggests that the cyclical sensitivity of wages has not changed over time.

The results reported here are consistent with the sectoral shifts hypothesis of unemployment. If sectoral shifts are truly independent of the cycle, then the difference between the actual rate of unemployment and the constructed natural rate series is a superior measure of labor market activity to the unadjusted unemployment rate. In accord with this interpretation, the regression analysis suggests that the actual and natural rates of unemployment have equal and opposite effects upon the wage inflation process.

The statistical insignificance of other variables usually found to be a determinant of wage inflation and the robustness of the results suggests that the measure of the natural rate captures a great deal of information. The appropriate interpretation of the results is somewhat difficult, however. One possibility is that sectoral shifts directly cause unemployment and therefore have an effect upon wage growth. Alternatively, these sectoral shifts may simply be symptomatic of some other fundamental underlying phenomenon. For this reason a more structural approach may be warranted. One suspects that wage growth is in fact related to a wide spectrum of variables such as unionization and import competition that a more aggregated, reduced form approach is unable to ascertain.

Finally, the sectoral shifts hypothesis and its relation to the wage inflation process has potentially important insights to yield for the role of government policy. First, and perhaps most obviously, the sectoral shifts hypothesis suggests that a countercyclical policy may be hard to implement because of the difficulty in interpreting movements in the unemployment rate. An expansionary policy adopted at the wrong time in the cycle may serve only to increase inflation rather than decrease the unemployment rate. Secondly, traditional monetary and fiscal policy effectiveness may depend in large part on the composition of unemployment. Such policies seem better suited to dealing with unemployment that is cyclical in nature but may prove to be largely ineffective in reducing unemployment that is attributable to long term sectoral shifts. Therefore, a combination of monetary and fiscal stimuli combined with industrial policy may be a more effective means of dealing with the problem of unemployment. Finally, although the Phillips curve relations estimated here are reduced form expressions, they suggest that there is a trade-off that policymakers can exploit between cyclical unemployment and the inflation rate. Furthermore, government policy may be nonneutral across industries. For example, monetary policy is likely to have a larger effect upon those industries that are relatively more interest-sensitive. Changes in policy may therefore affect the natural rate of unemployment and, hence, the nature of the trade-off between unemployment and inflation.

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[1] Recent works on the topic of sectoral shifts and the unemployment rate include Lilien (1982a, 1982b), Neumann and Topel (1984), Loungani (1985), Davis (1986a, 1986b), and Abraham and Katz (1986).

[2] The likelihood ratio is $\lambda = \prod_{j=1}^{8}(\hat{\sigma}_j^2)^5 / [\sum_{j=1}^{8} \hat{\sigma}_j^2 / 8]^{40}$ where $\hat{\sigma}_j^2 = \sum_{i=1}^{I} s_{ij}\left(\frac{\Delta e_{ij}}{\Delta e_j} - s_{ij}\right)^2$,
j = 1, ..., 8 indexes the cycles, and s_{ij} is the employment share of the i^{th} industry at the peak of the j^{th} cycle.

[3] While some economists have attempted to identify specific cycles as driven by supply shocks and others as linked to such factors as the Korean and Vietnam Wars, it seems rather arbitrary to define the resultant changes in the distribution of employment as a purely cyclical phenomenon. For example, in a general equilibrium model, factor price shocks, such as those observed in the seventies, undoubtedly affect some industries more severely than others. While aggregate output falls signaling a recession, the distribution of employment also changes due both to the direct effects of the factor price disturbance on production and the indirect effects of lower income on the level and distribution of demand. Thus, to argue whether the observed change in the distribution of employment across industries is the result of purely sectoral versus cyclical factors is in many ways misleading. Perhaps a more approporiate way of conceptualizing the distinction is to define the direct effects of such factor price shocks as sectoral while the indirect effects are cyclical--much as the effect of changes in price on product demand are decomposed into income and substitution effects.

[4] Even if some sectoral disturbances are transitory, for the purposes of modeling unemployment it is reasonable that only the permanent shocks are of importance since recognized temporary deviations imply a smaller benefit from adjustment.

[5] Industries included in the computations include government, construction, mining, durable manufacturing, nondurable manufacturing, transportation and public utilities, services, wholesale trades, retail trades, and finance, insurance, and real estate. See Neumann and Topel (1984) for computational details.

[6] In contrast, the series employed in Neumann and Topel (1984) extends only through 1977.

[7] Although the approach employed here follows that of Neumann and Topel (1984), the purposes are quite different. Neumann and Topel are concerned with the determinants of the geographical dispersion of unemployment rather than measures of labor market tightness and their relation to wage growth.

[8] As in Neumann and Topel (1984), the effect of transitory changes in the distribution of employment across industries was insignificant and therefore is unreported here.

[9] See Maddala (1977) for details.

[10] Several other lag structures were tested that are less restrictive. However, the regression results did not improve significantly and are unreported here due to space limitations. In addition, the estimation was also performed with the unemployment rate entered in inverse form. These regressions in general had lower Durbin-Watson statistics and lower values of R^2.

The estimation was carried out using a one-step procedure. Therefore, the coefficient δ_2 is not identified without further restrictions. The product of δ_2 and a_2 is found to be negative. Hence, either $\delta_2, b_i > 0$ or $\delta_2, b_i < 0$. The former interpretation is more appealing as higher expected price inflation is thought to be associated with greater wage growth and, intuitively, higher past inflation should lead to larger current expectations of inflation.

A two-step procedure which is unreported in the text was also performed. First, the parameters of the polynomial $b(L)$ were estimated by an unrestricted regression of current price inflation on eight lagged values. The sum of the coefficients was 0.88 with an estimated standard error of 0.07. This model was used to predict actual inflation which was then used as a regressor in estimating equation [2]. The estimate of δ_2 calculated in this way is 0.69 with a standard error of 0.07. Because of the errors in variables problem that occurs, this estimate of δ_2 is more properly thought of as a lower bound. By calculating the variance-covariance matrix of the observations, the coefficent estimate on the unemployment rate is also found to be biased towards zero.

[11] The age-weighted unemployment rate, rather than the actual civilian unemployment rate, is used in the analysis since the estimates of the natural rate were constructed from the regression results on this dependent variable.

[12] The errors in variables problem is more complicated for this model than for the previous one. Both expected price inflation and the natural rate of unemployment are measured with error. Therefore, the bias introduced is a complicated nonlinear function of the variance-covariance matrix of the observations. Theil (1961) finds that in the case of two variables measured with error, the direction of bias depends upon the correlation between the true values of expected inflation and the natural rate. If the correlation is zero, then the OLS parameter estimates are biased towards zero with the magnitude depending upon the ratio of the variance of the measurement error to the variance of the true values. If the true natural rate and true expected inflation are correlated, then the OLS estimates may either underestimate or overestimate the true parameters of the model. The estimated correlation coefficient between the two constructed series is 0.438. Using this as an estimate of the true correlation, the OLS estimates are more severely biased towards zero than in the case where only one variable is measured with error.

[13] The estimate of the coefficient on expected inflation is 0.42 with an estimated standard error of 0.09. However, it should be noted that the OLS estimate is likely to be biased downwards and may be considered a lower bound.

[14] See Neumann and Rissman (1984) for a discussion.

[15] Union membership and the sum of union and employee association membership are found annually in the BLS' *Directory of National Unions and Employee Associations* through 1978. The sum of the two is continued in the BNA's *Directory*

of U.S. Labor Organizations. The data were used to construct a union membership series through 1981. Quarterly observations were computed by linear interpolation of the annual data.

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Figure 1
Permanent and Temporary Changes in the Distribution of Employment Across Industries

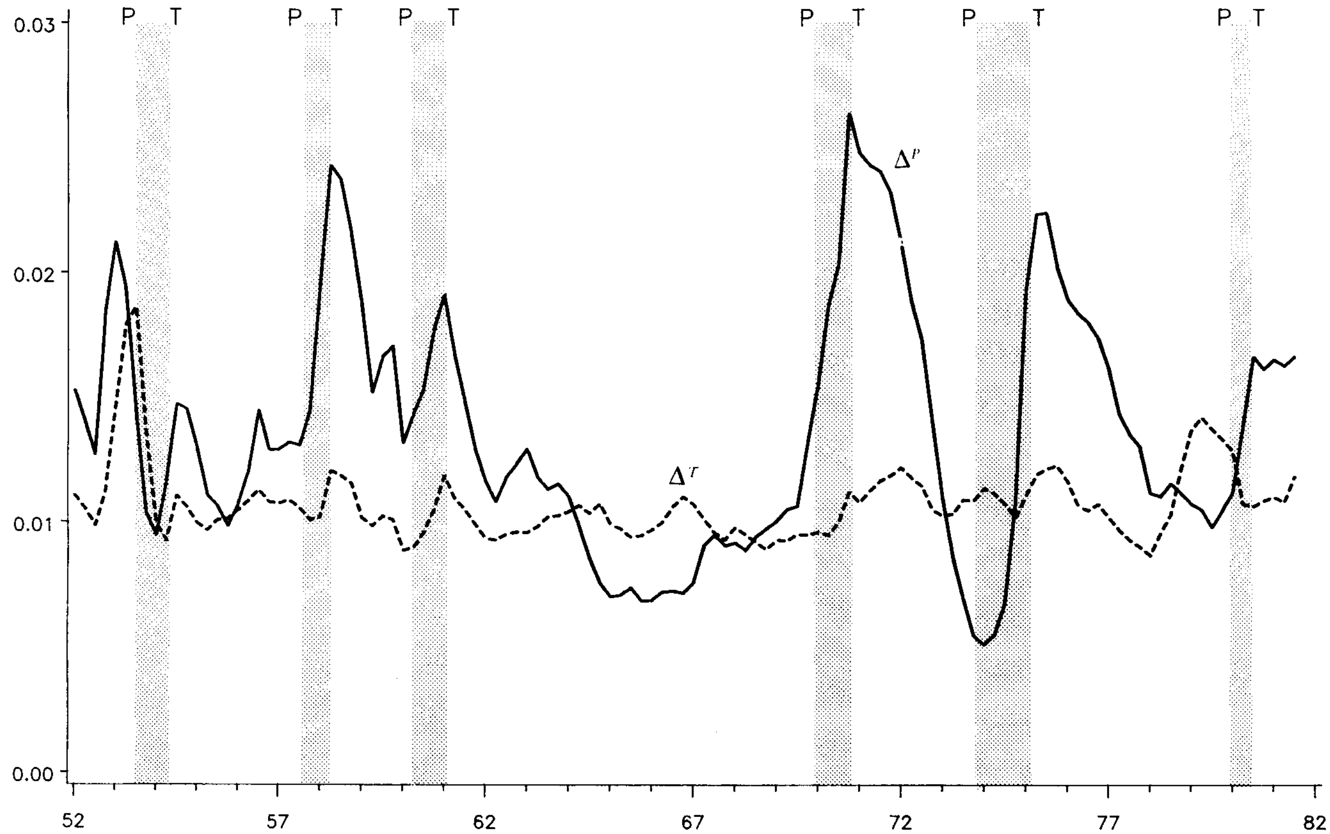

Figure 2
The Natural and Fixed-Weight Unemployment Rates
1958-1981

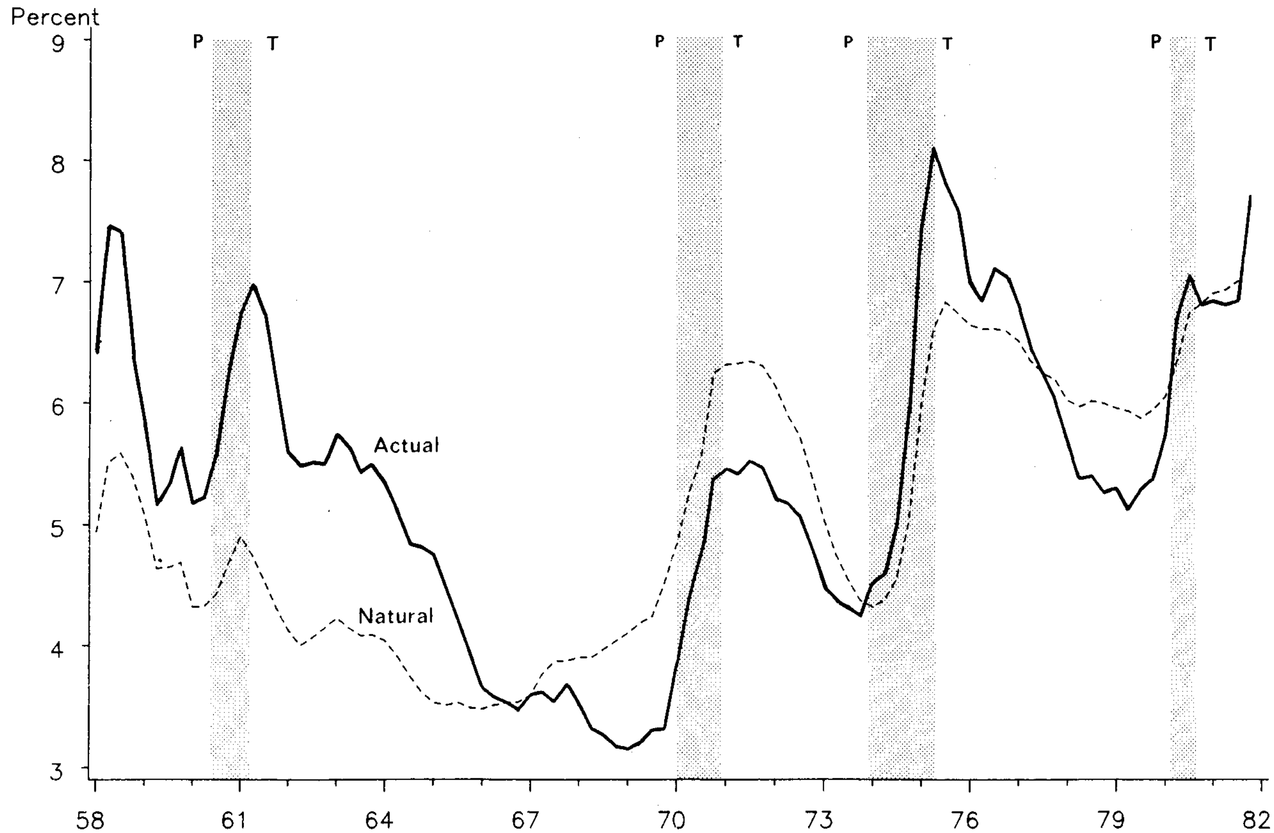

Table 1
Industry composition of durable manufacturing employment changes from peak to trough
Net of employment share

	48:4-49:4	53:3-54:2	57:3-58:2	60:2-61:1	69:4-70:4	73:4-75:1	80:1-80:3	81:2-82:4	$\hat{\sigma}^2_i$
Lumber and Wood Products	-0.051	-0.035	-0.025	0.027	-0.031	0.070	0.041	-0.004	0.0018
Furniture and Fixtures	-0.029	-0.006	-0.016	0.005	-0.019	0.038	0.015	-0.012	0.0005
Stone, Clay and Glass	-0.025	-0.025	-0.029	-0.006	-0.038	0.011	0.018	-0.003	0.0004
Primary Metals	0.150	0.028	0.059	0.124	-0.015	-0.034	0.085	0.087	0.0042
Fabricated Metals	-0.008	-0.002	-0.024	-0.002	-0.002	0.014	0.032	0.018	0.0003
Machinery, Except Electrical	0.052	-0.026	0.030	-0.030	-0.029	-0.163	-0.092	0.055	0.0056
Electric and Electronic Equipment	-0.017	0.029	-0.029	-0.102	-0.034	0.067	-0.042	-0.088	0.0031
Transportation Equipment	-0.031	0.051	0.075	0.011	0.196	0.030	-0.001	0.000	0.0050
Instruments and Related Products	-0.010	-0.006	-0.017	-0.019	-0.013	-0.039	-0.057	-0.043	0.0003
Miscellaneous Manufacturing	-0.032	-0.009	-0.024	-0.008	-0.014	0.005	0.001	-0.010	0.0001
$\hat{\sigma}^2_j$	0.0044	0.0010	0.0020	0.0038	0.0073	0.0062	0.0031	0.0029	-

SOURCE: *Employment and Earnings,* various issues.

Table 2
Parameter instability of the standard Phillips curve specification
(1960:2 to 1981:3)

	(1)	(2)	(3)	(4)
UR	-0.0013 (3.312)	-0.0020 (5.091)	-0.0016 (3.658)	-0.0018 (4.730)
$\dot{p}^e$	-0.0051 (9.692)	-0.0033 (4.970)	-0.0037 (3.507)	-0.0054 (5.373)
c	0.0144 (7.900)	0.0186 (9.447)	0.0167 (7.084)	0.0161 (7.652)
DxUR	-	0.0009 (4.052)	-	0.0017 (4.720)
$Dx\dot{p}^e$	-	-	-0.0014 (1.520)	0.0038 (2.733)
R^2	0.562	0.635	0.574	0.666
DW	1.469	1.784	1.506	1.959

NOTE: The dependent variable is the first difference in the logarithm of Average Hourly Earnings reported in various issues of *Employment and Earnings* by the BLS. T-statistics are in parentheses.

Table 3
The Phillips curve and sectoral shifts
(1960:2 to 1981:3)

	(1)	(2)	(3)	(4)	(5)
WUR	-0.0024 (5.593)	-0.0025 (4.625)	-0.0031 (6.463)	0.0000 (0.011)	-0.0024 (5.628)
Natural Rate	0.0026 (3.671)	0.0018 (2.179)	0.0031 (4.370)	-0.0026 (0.974)	-0.0024 (5.628)
$\dot{p}^e$	-0.0031 (4.678)	-0.0022 (2.531)	-0.0004 (0.321)	-0.0096 (2.231)	-0.0033 (8.409)
c	0.0097 (4.359)	0.0135 (5.601)	0.0123 (5.290)	0.0148 (5.938)	0.0104 (14.722)
DxWUR	-	0.0000 (0.003)	-	-0.0024 (1.299)	-
DxNatural Rate	-	0.0007 (0.607)	-	0.0048 (1.847)	
$\mathrm{Dx}\dot{p}^e$	-	-	-0.0024 (2.821)	0.0077 (1.757)	-
R^2	0.641	0.684	0.673	0.696	0.641
DW	1.793	2.045	1.975	2.096	1.787
F test	0.749	1.269	0.0005	4.372	-

NOTE: The dependent variable is the first difference in the logarithm of Average Hourly Earnings reported in various issues of *Employment and Earnings* by the BLS. T-statistics are in parentheses. The F statistic reported tests for equal and opposite signs of the parameters on WUR and the Natural Rate.

Table 4
Additional Phillips curve estimates

	(1)	(2)	(3)	(4)	(5)	(6)
WUR	-0.0024 (5.587)	-0.0025 (5.647)	-0.0026 (5.698)	-0.0027 (5.780)	-0.0025 (5.494)	-0.0023 (4.938)
Natural Rate	0.0024 (3.262)	0.0025 (3.411)	0.0029 (3.884)	-0.0025 (3.102)	-0.0025 (3.533)	0.0024 (2.926)
$\dot{p}^e$	-0.0032 (4.693)	-0.0032 (4.756)	-0.0024 (2.804)	-0.0027 (3.028)	-0.0031 (4.618)	-0.0033 (4.518)
Female	-0.0590 (0.722)	-	-	-0.0836 (1.004)	-	-
Nonwhite	-	0.0448 (0.959)	-	0.0540 (1.130)	-	-
Youths	-	-	-0.0405 (1.242)	-0.0410 (1.253)	-	-
Union	-	-	-	-	-0.0242 (0.623)	-
WPC	-	-	-	-	-	0.0008 (0.597)
c	0.0111 (3.756)	0.0101 (4.461)	0.0111 (4.457)	0.0136 (4.132)	0.0100 (4.369)	0.0010 (4.372)
R^2	0.644	0.645	0.648	0.652	0.643	0.643
DW	1.805	1.841	1.822	1.906	1.795	1.804
F test	0.0003	0.006	0.184	0.0027	0.0026	0.0085

NOTE: The dependent variable is the first difference in the logarithm of Average Hourly Earnings reported in various issues of *Employment and Earnings* by the BLS. T-statistics are in parentheses. The F statistic reported tests for equal and opposite signs of the parameters on WUR and the Natural Rate.