No. 8005
(revision of No. 8003)

UNIONISM AND THE CYCLICAL BEHAVIOR OF THE LABOR MARKET IN U.S. MANUFACTURING

by

James E. Pearce

December 1980

Research Paper

FEDERAL RESERVE BANK OF DALLAS

No. 8005
(revision of No. 8003)

UNIONISM AND THE CYCLICAL BEHAVIOR OF THE LABOR MARKET IN U.S. MANUFACTURING

by

James E. Pearce

December 1980

UNIONISM AND THE CYCLICAL BEHAVIOR OF THE LABOR MARKET IN U.S. MANUFACTURING

During the 1970's, evidence from contractions in economic activity and theoretical developments in business cycle analysis stimulated renewed interest in the cyclical behavior of the labor market. Considerable research was, and continues to be, devoted to the explanation of why movements in aggregate demand produce large fluctuations in employment and little or no change in nominal wage growth. The pattern exhibited in the 1973-1975 recession is an example of this phenomenon. Nominal wages continued to rise at a 6-8 percent annual rate while unemployment increased from 4.9 to 9.1 percent.

Most cyclical variation in employment arises from shifts in the demand for labor in the goods-producing sector of the economy, where about half the production workers belong to labor unions. Unionism has long been regarded as a source of wage rigidity,1/ so an examination of its responsibility for the pattern of sticky wages and variable employment is worth pursuing. This paper investigates the differences in the responses to cyclical shifts in demand by employers of union and nonunion labor. The empirical analysis covers the adjustments in the wage rates, weekly hours worked, and employment of production workers in manufacturing during the 1973-1975 recession. Cross section analysis of the effect of industry demand conditions on industry wage levels provides information on wage flexibility not obtainable from the usual time series methods. A comparison of the effect of the demand variable on the wages of union and nonunion workers tests the union wage rigidity hypothesis. The comparison of adjustments in hours and employment in the two environments proceeds along similar lines.

The question of how unionism affects a firm's response to demand variation falls within the domain of the recent theoretical work on the importance of contracting in explaining cyclical employment fluctuations. Traditional analysis views the setting of nominal wages in long term contracts as an important element in the explanation of why nominal shocks produce real effects. The presumed greater rigidity in the union wage should produce more cyclical variation in the employment of union members than in the employment of similar nonunion workers. Modern models of efficient contracting, which emphasize the durability of the typical employer-employee relationship, imply that employment and wages can vary independently over the life of this relationship. Thus, wage rigidity may be irrelevant to the cyclical variation in unemployment. (See the interchange between Fischer and Barro.)

The recent work has also produced some insights into possible contributors to the stickiness in wage rates. Wage inflexibility in models of efficient contracting is a consequence of infrequent job changes combined with workers' desires for stable incomes. Other analysts have concluded that adjustment costs will also cause wages to respond to shocks more slowly in labor markets in which turnover is low (Williamson and Wachter). These theories appear to be designed to explain the behavior of the labor market in manufacturing, which exhibits the low separation rates and frequent large shifts in demand stressed in this literature.

Thus, the recent theoretical work has raised the possibility that

(1) unionism may not be a primary source of wage rigidity; and

(2) the magnitude of cyclical employment variation may be independent of the response of wages to cyclical shocks.

Because these are important propositions, further discussion of the recent literature of contracting and its implications for the effects of unionism precedes the description of the empirical work. This discussion is brief, for its purpose is merely to provide a pair of alternate frameworks that will be useful in interpreting the results.

Preliminary inspection of the data suggests that union members did indeed have more rigid wages and more variable employment between 1973 and 1975. The union-nonunion wage differential rose 3 percentage points, and the percentage of union employees working overtime declined twice as much as the percentage of nonunion employees working more than usual. In May of 1975, layoff unemployment was more than 2 percentage points higher among union workers than among nonunion workers. The more detailed analysis, which controls for differences in employment changes across industries, indicates the cyclical responses of wages and hours differ sharply in the union and nonunion sectors. Smaller, but still significant, differences are found in employment adjustments.

In view of the large difference in the responses of union and nonunion wages, the similarity of the employment adjustments in the two environments is interesting. The result has a number of possible causes. First, the union and nonunion sectors are not independent of one another, so a decline in production from union plants could cause a drop in demand or a shortage of material in nonunion plants. Second, the hourly wage is not the only form of compensation, and union employers may make larger adjustments in fringe benefits than nonunion employers. Third, this may be an indication that wage and employment variation are indeed independent. Overall, the evidence indicates that unions matter, but the process through which collective bargaining influences wage, employment, and hours adjustments is not fully revealed.

I. Theoretical framework

A. Labor contracting in manufacturing

The principal feature that distinguishes contract markets from auction markets is the durable association between trading partners. In labor market analysis, hiring and training costs (Becker; Oi) and opportunities to redistribute the risk arising from future shifts in market conditions (Gordon; Baily (1974); Azariadis) receive frequent mention as incentives for employers to reduce turnover.

Recognition of the prevalence of durable associations between employers and their workers has led to a treatment of the short run labor adjustment that is, in many respects, similar to the analysis of the employment of capital in the short run. (Baily (1974) and Azariadis are well-known examples.) The employer selects a stock of workers of the optimal size and a strategy for meeting any variation in demand that may occur over the planning horizon. This strategy may include variation in the wage and in the rate of utilization of the labor stock, but changes in the stock itself are presumed to be prohibitively costly. Under these conditions, optimal adjustments in utilization are made independently of adjustments in the wage. Straightforward assumptions about the value of workers' time off the job and the risk preferences of employers and workers imply that the optimal strategy will involve variation in utilization but no variation in the wage. If the attachment between a firm and the typical employee survives several cyclical episodes, then an implication is that although wage rigidity is to be expected, it has no influence on the amplitude of cyclical variation in unemployment.2/

Since wage rigidity is irrelevant from this perspective, the strategy adopted for meeting variation in demand depends on the costs of storage and maintaining flexibility in production. Even if the wage is constant, variation in utilization causes variation in earnings that employees will accept only if they receive a compensating premium over the wage paid for employment that involves a constant earnings stream. Thus, employers facing variation in product demand have an incentive to dampen its impact on their employees. They may pursue this objective by smoothing production through the carrying of inventories, by absorbing some of the impact of production variations through the "hoarding" of labor during off-peak periods, or by simply not meeting all demand variation.

The arrangements between firms and employees in the models of Azariadis and Baily have been called contingent claims contracts, because they specify in advance the wage and utilization paths for each set of contingencies that might possibly develop over the planning horizon. Skeptics argue that such contracts are not observed in blue-collar labor markets because the cost of enforcement is too high. Workers must be able to observe their employer's compliance with the terms of any agreement, and under a contingent claims contract this would require that they monitor the course of demand. Because this is extremely costly, contracts in which the verification of compliance does not require both parties to obtain timely market information independently will dominate contingent claims contracts (Williamson, Wachter, and Harris; Calvo and Phelps). Hall and Lilien contend that contracts in blue-collar labor markets do not require close monitoring of demand by workers. The agreements tie earnings to utilization in such a manner that the employer's pursuit of profit maximization also maximizes the employees' utility. Employees can detect noncompliance by monitoring utilization, a task much less costly than monitoring demand.

Under contracts of this type, the hourly wage assumes great importance. Consequently, it responds slowly to changes in market conditions, and this inflexibility contributes to variation in labor utilization (although the considerations enumerated above will also be important). Wage rigidity arises from the asymmetrical distribution of information about market conditions and the presence of appropriable quasi-rents attributable to the impediments to worker mobility. Employees resist wage cuts because they have difficulty distinguishing a genuine reduction in demand from an attempt by their employer to exploit his short run monopsony power (Feldstein (1975); Williamson and Wachter; Klein, Crawford, and Alchian). Reductions in either utilization or wages will reduce labor earnings, but hours and employment adjustments arouse less suspicion because output reductions generally accompany utilization reductions and ensure that the employer will receive less revenue after the change. Awareness of workers' disinclination to accept wage reductions deters employers from raising wages to long run equilibrium levels during periods of temporarily high demand.

Identifying which of the two types of contracts prevails in blue collar manufacturing has important consequences for theoretical predictions of the effect of unions on the response to demand variation. If contingent claims contracts prevail, the possibility that unions have little or no effect on employment variation cannot be dismissed. Under the alternative type of arrangement, the effect of unionism could be substantial. Because the issue remains unresolved, the implications of introducing collective bargaining into the analysis are explored from both perspectives in the following subsection.

B. Unionism and the response to cyclical shifts

Possible sources of greater rigidity in union wages may affect other dimensions of the response to demand variation, so they will be discussed first. Since the topic has received considerable attention elsewhere, the review here covers only the arguments that fit into the contracting framework outlined above. This restriction limits consideration to two channels of influence--the possibility that union members remain with their employers longer and the possibility that unions raise the cost of temporary wage adjustments. The discussions by Dunlop, Mitchell, and Moore and Raisian list some other potential contributors to union wage rigidity.

Several students of collective bargaining have mentioned reasons for expecting a positive association between unionism and the strength of the employer-employee attachment. Union wages are generally higher than nonunion wages, and wages and turnover are negatively related (Parsons). Nonwage considerations such as seniority systems (Rees (1959)) and grievance systems (Freeman) may also reduce turnover in union plants, and Lewis (1959) suggested that workers in occupations in which the typical employer-employee relationship is more durable will find the services of union leaders more valuable. Freeman's evidence that union members, on average, have accumulated more seniority on their current jobs than nonunion workers is consistent with these propositions.

This difference in mobility may cause employers in the two environments to view cyclical activity from different perspectives. The relative immobility of union workers may lead their employers to regard cyclical shifts as short-run disturbances that do not warrant adjustment in wages or labor stocks. Nonunion employers would be less likely to adopt a perspective in which cyclical activity is regarded as such a short run phenomenon, so their adjustment of their labor stocks will exert pressure on wages. The literature on union wage behavior does not often cite immobility as a cause of union wage rigidity, however. Dunlop may have had something of this nature in mind when he wrote (page 68), "The basic wage rate is regarded as a long-run price, usually set with an eye to noncyclical circumstances," although his statement precedes the emergence of the modern layoff models by decades. More recently, Raisian appears to be using this line of reasoning in explaining the results of his economy-wide comparisons of the cyclical behavior of union and nonunion labor markets.

The effect of collective bargaining on adjustment costs has been mentioned much more often as a source of union wage rigidity. Unionism may repress wage variation by reducing the frequency with which wages are adjusted and by reducing the amount by which wages are changed when adjustment does occur. Strikes occur most often during contract negotiations, and the desire to avoid the costs of such interruptions provides an incentive to renegotiate contracts less frequently (Mitchell). Union wage insensitivity follows directly, for the importance of temporary conditions can be expected to decline as the period over which the terms of the contracts are held constant increases (Rees (1951)). The possibility that organized workers may be able to resist wage cuts more effectively than unorganized workers has also been raised (Dunlop; Morton). If this is so, unionized employers will be less able to reduce wages in periods of temporarily low demand and, consequently, less willing to raise them in periods of temporarily high demand.

Arguments similar to those raised in the discussion of wage flexibility imply that adjustments in utilization are likely to be greater under collective bargaining. In the case in which contracts tie utilization and earnings, larger adjustments in hours and employment by employers of union labor follow directly from the higher costs of adjusting union wages. The ability of unions to raise wages provides another reason for expecting larger utilization adjustments in union plants, even under contingent claims contracting. The wage premium paid by union employers should enable them to raise their hiring standards and maintain greater variability in their production schedules relatively cheaply.

Unionism may also influence the division of adjustments in the labor input between changes in employment and changes in hours per worker. The relative strength of employees' preferences for stability in weekly hours versus continuity of employment will influence the relative costs of adjusting hours and employment through the terms of the employment contract. Because union contracts, which can be observed easily, commonly restrict the employer's freedom to impose worksharing or overtime, the temptation exists to conclude the strength of union members' relative preferences for hours stability exceeds that of nonunion workers. Avoiding extreme hours reductions through layoffs is a practice common among both nonunion and union employers, however. Futhermore, Baily (1976) has noted that layoffs prevailed over worksharing even before many workers were covered by unemployment insurance or collective bargaining contracts. Thus, no compelling reason to expect the preferences of union and nonunion workers to differ is apparent, although the wishes of union workers may be observed more strictly if the terms of explicit union contracts are more effectively enforced.

The principal avenue through which unionism is likely to reduce the cost of layoffs is its effect on the probability that a layoff will cause an employee to change jobs.3/ Although most layoff models ignore this consideration (Baily (1976) is an exception), the rehire rate varies considerably across industries (see Lilien), and the recall probability is likely to vary substantially during recessions, when layoff duration arises. All of the influences on employee mobility mentioned earlier should cause the recall probability of union members to decay relatively slowly as the length of the layoff increases, so cyclical employment adjustments should contribute less to turnover in union establishments. Unless these employers face greater hiring and training costs, this factor should make union employers less reluctant to lay employees off.

II. A description of the empirical approach

The empirical analysis in this paper compares changes in wages and utilization of union and nonunion workers in manufacturing over the 1973-75 recession. The period was an interesting one, for it contained shifts in the demand for labor resulting in large reductions in labor utilization, particularly in manufacturing. Table 1 documents the increase in unemployment among manufacturing workers due to employer-initiated separations. In addition to the rise in unemployment, the rate of growth in consumer prices accelerated sharply. In the year preceeding May of 1973, the CPI increased 6 percent. In the two subsequent 12-month periods, the index rose 10 percent and 9 percent, respectively.

The primary data for the investigation are from the Current Population Surveys (CPS) conducted in May of 1973, 1974, and 1975. The May CPS interview has several advantages for comparison of the behavior of union and nonunion sectors and for examination of cyclical labor market behavior in general. The files identify union members and hourly production workers, and they reveal the hourly wage and employment status of the people surveyed. Analysis of adjustments in hours worked is facilitated by information on "normal" weekly hours as well as hours actually worked in the survey week. The large sample with considerable industry detail is also helpful.

A. Previous empirical work

Although theory indicates that union wages should exhibit less sensitivity to cyclical demand shifts than nonunion wages, the evidence, particularly with respect to manufacturing wages, is mixed. In an economy-wide study using aggregate wage series from 1920-1958, Lewis (1963) found the union wage premium sensitive to changes in inflation but insensitive to unemployment. Ashenfelter, Johnson, and Pencavel analyzed wage behavior over the period 1954-1968 for two-digit manufacturing industries. In a Phillips equation that included interactions of a union coverage measure with inflation and unemployment variables, they found that wages in heavily covered industries displayed a significantly smaller response to both cyclical indicators. Pierson, on the other hand, ran separate Phillips equations for lightly and heavily organized manufacturing industries and found larger responses in the heavily organized sector.

Although there have been other aggregate studies (see the survey by Moore and Raisian), the recent work has used micro data. Raisian, in an economy-wide analysis of the period 1967-1974, found union wages less sensitive to unemployment than nonunion wages. But in a similar study, Moore and Raisian found little difference with respect to union status when they restricted the sample to manufacturing workers. Mitchell, in an analysis of wage data from individual manufacturing establishments, found evidence of union wage rigidity with respect to both inflation and unemployment.

Recent research has produced evidence consistent with the hypothesis that the employment of union workers is more susceptible to interruption by layoff than the employment of nonunion workers. Feldstein (1978) found union members had a higher layoff unemployment rate in March of 1971, which was a recession year. Medoff obtained a positive coefficient for union coverage in a model analyzing variation in average layoff rates across state by industry cells. Raisian found the relationship between changes in annual weeks worked and changes in industry unemployment rates over the 1967-1974 period to be stronger for union members than for other workers. Feldstein and Raisian used microdata that distinguished union from nonunion workers, and their samples included workers from all industries. Medoff confined his analysis to layoffs in manufacturing industries.

The interpretation of the existing evidence on the relationship between unionism and the cyclical behavior of wages and employment contains some ambiguity. Most of this evidence consists of relationships between union coverage and changes in wages or employment estimated from aggregate data. Such models generally encounter particular difficulty in distinguishing behavioral differences induced by collective bargaining from differences attributable to environmental characteristics associated with unionism. The use of aggregate unemployment to measure differences in business conditions adds another weakness, for it does not capture differences in demand changes across industries or occupations. Studies that used micro data and did not rely on an aggregate unemployment rate, such as those of Feldstein and Raisian, drew their evidence from samples that contained workers from all sectors of the economy and procedures that employed imprecise controls for differences across industries.

B. The data and method

The evidence in this paper is drawn from a sample in which industry variables from establishment data have been merged with observations on individual workers in manufacturing from the CPS. The sample contains only white male production workers who are paid by the hour. The restriction of the sample should reduce differences in unobserved attributes that differentiate union and nonunion workers, and the industry variables should capture some of the effects of differences in organized and unorganized plants that would be ignored if the CPS information was not supplemented. Although the possibility cannot be dismissed that any union-nonunion differences in the response to demand variation implied by the resulting estimates arise from selectivity or omitted variable bias, the evidence here should contain less such contamination than the results of previous work.

The primary analytical tool is multiple regression, in which the unit of observation is the individual worker. Three models are estimated; their dependent variables reflect the individual's wage rate, hours worked, and employment status in the May reference week for the CPS of the relevant year. The right side of the equations contain individual characteristics from the CPS record and industry variables attached to the CPS record. Because the principal source of differences in the sensitivities of employees' earnings to the business cycle is likely to be the cyclical volatility of the employer's demand, the discussion of the regressions focuses on the coefficients of a variable intended to capture variance in demand conditions across industries and over time. Its value is the difference between E_{it}, the log of employment for industry i in May of year t, and a forecast of E_{it} based on the trend in industry employment over the 7 1/2-year period ending in May of year t-1. The CPS industry code is sufficiently detailed to identify 73 industries within manufacturing to which employment series from the BLS Employment and Earnings publication could be matched.

Higher values of this prediction error correspond to higher levels of residual employment, from which greater excess demand is inferred. The variable is thus expected to be positively related to wages and utilization. The models are of the form

$$y = a_1P + a_2D(U)P + bX,$$

where P is the prediction error, D(U) is the union membership dummy, X is a vector of other variables, and a and b are coefficient vectors. a_1 is the relationship between the prediction error and the dependent variable for nonunion workers; a_2, the interaction coefficient, indicates the difference between the estimates for union and nonunion workers, and a_1+a_2 is the relationship between the prediction error and the dependent variable for union members. The union wage rigidity hypothesis implies a negative interaction coefficient in the wage equation, and the hypothesis that utilization is more sensitive in the union sector than the nonunion sector anticipates a_1 and a_2 will have the same signs in the employment and hours models.

The use of employment series to construct the measure of demand shifts raises some econometric issues. One problem is that the prediction error may contain some of the response that its coefficient is intended to measure. Since hours adjustment is a substitute for employment adjustment, groups that have more flexible hours will tend to have smaller fluctuations in employment, so the estimate of the hours response will be biased toward zero. The estimated wage response will contain a similar bias if adjustments in wages and utilization are substitutes. The use of the employment-based variable assumes that variance in the magnitude of shifts in the demand for labor is very large relative to the variance in the response to such shifts.

Another problem arises in the employment status regressions, in which the dependent variable indicates whether the individual was working during the survey reference week. Including the prediction error for industry employment among the explanatory variables of such an equation amounts to putting variants of the same concept on both sides of the equal sign, so the response estimate is likely to contain serious bias. Interest in this paper centers on the distribution of changes in unemployment rather than their magnitude, however. Thus, only the relationship between the responses of union and nonunion unemployment is important. The regressions containing the prediction error are not likely to indicate a significant difference in these responses if they are in fact equal, so the variable is adequate for this limited purpose.

C. Specific models and descriptive statistics

Section III discusses regression analysis of the behavior of real wages. The log of the straight-time hourly wage, deflated by the Consumer Price Index, is the dependent variable. The model assumes that human capital accumulation as well as industry and area characteristics determine the long-run equilibrium wage for each individual. Short-run conditions, captured by the prediction error, cause departures from long-run values. Each industry is described by the amount of value added per establishment, union coverage, and the volatility of employment.4/ The last variable is approximated by the sum of squared residuals from a quadratic trend regression of the log of industry employment over the years 1958-1975. Section IV discusses two regressions, one for employment status and the other for hours worked. Both use a qualitative dependent variable with values ranging from one to three. The estimates reported are partial derivatives obtained from maximum likelihood estimation of conditional logit models.

The discussion in the previous section emphasized short-run employment adjustments via temporary layoffs and rehires, so examination of the cyclical behavior of temporary layoff unemployment is the principal element of the empirical analysis of employment variation. Focusing on layoff unemployment alone, however, would be misleading. As Table 1 illustrates, year-to-year variation in unemployment due to permanent discharges was larger in the nonunion sector than in the union sector.

This difference probably reflects the joint influences of the types of employers that become unionized and the constraints introduced by unions. The adjustment of employment through layoffs and rehires, rather than through discharges and new hires, is a practice generally associated with larger enterprises, which have self-contained "internal labor markets," and most of these employers are unionized. (See Wachter and Williamson and the references therein.) In addition, union contracts may make permanent dismissals more costly in order to protect the gains union workers achieve through collective bargaining.

The pattern in Table 1 raises the point that the larger rise in union layoff unemployment might reflect a substitution of layoffs for discharges, with little difference in the overall employment adjustment in the two sectors. To investigate this possibility, the dependent variable in the employment status model, which compares the employed with the unemployed, distinguishes people who expect recall from those who do not. Those unemployed via voluntary separation and those employed but not working for reasons other than layoff were excluded from the sample. The model contains the value added per establishment variable and a union membership dummy to control for the effects of employer size and collective bargaining on the layoff-discharge decision.

The dependent variable in the hours adjustment model is somewhat unusual. Table 2 describes the distribution of a variable designated "excess hours," which is the difference between actual and normal weekly hours expressed as a percentage of normal hours. The stability across years exhibited by the means of the positive and negative values of this variable is remarkable. Presumably, it reflects the presence of upper and lower bounds on hours worked implied by models such as Baily's (1977). This property complicates the analysis, for a regression using excess hours as the dependent variable yields prediction error coefficients with the wrong signs. The table indicates that most year-to-year variation in the variable arises from variation in the proportions of people working more or fewer hours than normal. Therefore, the hours adjustment model reported in Section IV uses a qualitative dependent variable whose value is determined by the sign of excess hours.

Table 3 contains some statistics indicating how the industry measures vary across CPS industry groups. The prediction errors show a very large decline in residual employment over the period. Employment was above the level predicted from past trends by an average of 13.5 percent in May of 1973, but by 1975 it had fallen to 11 percent below trend. A comparison of the means of the shift and volatility measures in the durable and nondurable sectors indicates that employment has been more stable in nondurables by a factor of about three. Establishments in nondurables are also slightly smaller and less heavily unionized.

Table 4 shows the distributions of the industry variables after their attachment to the records of the CPS. The unit of observation here is the individual worker. The means of the shift measures for union and nonunion workers reveal that the concentrations of workers in industries that experienced large reductions in employment were about the same for both union members and nonunion employees. Thus, any union-nonunion differences in utilization adjustments are not so likely to be a consequence of greater union penetration in the more volatile durable goods sector. The table does indicate a marked difference in the concentrations of the two types of workers in industries characterized by large plants, however.

III. Adjustments in Wage Rates

The wage equations discussed in this section contain a mixture of standard and less common variables. In addition to the industry measures described earlier, the equations include age and its square, years of schooling, and sets of dummies capturing variance attributable to differences in marital and union status, area population density, region, occupation, and state laws regarding the establishment of the union shop.5/ The industry variables and the right-to-work dummy are interacted with the union membership dummy. Table A-1, appended, contains the coefficients of the variables that are not of primary concern in this paper. Space does not permit a thorough discussion of these estimates, but a few are of sufficient interest to note briefly.

The positive coefficients for the union coverage variable and its interaction with the union membership dummy indicate that wages of both union and nonunion workers are higher in the more heavily organized industries and that the union-nonunion wage differential is a positive function of coverage. These relationships are consistent with conventional wisdom, which holds that union power is positively related to coverage and that unionism raises nonunion wages through threat or spillover effects. Other factors could explain the coefficients, but the issue will not be pursued further here. The right-to-work and establishment size coefficients indicate nonunion wages are higher in larger establishments and in states that allow the union shop, but union wages appear to be almost independent of these variables. Although this, too, has more than one possible explanation, the threat effect investigated by Rosen (1969) is an attractive candidate. Union coverage is higher in union shop states and industries characterized by large plants. Therefore, the estimates suggest that nonunion employers in greater jeopardy of being organized pay higher wages than other nonunion employers in order to reduce the appeal of unionism.

Table 5 displays the coefficients of variables included to capture the effects of cyclical activity. Three specifications were estimated to investigate the sources of year-to-year wage variation, which is assumed to arise solely from the combined effects of productivity trends and the rises in unemployment and inflation. The union wage rigidity hypothesis implies that the rise in unemployment should have reduced growth in nonunion wages relative to growth in union wages, but the rise in inflation should have at least partially offset that reduction. Therefore, the net change in the union-nonunion differential cannot be predicted. The estimates in column 1 measure the net changes in union and nonunion wages, and the estimates in column 2 isolate the component of the change attributable to the rise in unemployment. The specification in the third column assumes all between-year variance in real wages was due to this factor; it is included simply to examine the performance of the prediction error variable.

Looking first at the bottom three rows of coefficients, the estimates offer no evidence that historic employment variability influences long-run wage levels. The estimated elasticities with respect to the measure of stochastic employment variability are small and positive for nonunion workers and small and negative for union workers. In only one case is an estimate larger than its standard error. This finding is curious, because authors developing layoff models have emphasized the proposition that employers having reputations for extensive variation in labor utilization must pay higher wages to attract workers. The small elasticities may be a consequence of the reduction of variance in risk though unemployment insurance, misspecification of the model, an inappropriate measure of unemployment risk, or the restriction of the sample to manufacturing workers. Consequently, the estimates do not necessarily indicate that the proposition does not hold.6/

The attempt to measure short-run responses in real wages was more successful. The specification reported in the first column contains a set of year dummy variables and the interaction of this set with the union membership dummy. This permits estimation of separate intercepts for the union and nonunion wage functions in each year. The 1974 and 1975 dummy coefficients indicate that the intercept of the nonunion wage function shifted downward about 4.5 percent in the two years following May 1973. The coefficients of the interaction variables (D(1973)xD(U), et cetera) allow the inference of a separate union-nonunion wage differential in each year. The coefficient estimates imply this differential rose almost 3 percentage points over the period, so the downward shift in the union wage function was less than 2 percent.

The specification reported in column 2 adds the industry employment prediction error and its interaction with the union membership dummy to capture the effect of the shifts in industry demand. The prediction error coefficient is positive and significant, implying procyclic variation in the real wages of nonunion workers. The interaction coefficient is negative, indicating that the shifts had a smaller impact on the real wages of union members. The sum of these two coefficients, labeled "a_1+a_2" in the table, suggests slight procyclic variation in union real wages. The third column reports the estimates from an equation containing the prediction error variables without the year dummies. The pattern of the coefficients is similar to that of column 2.

The addition of the prediction errors has a striking effect on the year dummy coefficients. The estimates indicate that nonunion wages, once they have been adjusted for the effects of the demand shift, did not change in the first year and rose about 3 percent in the second year. Accordingly, the rising pattern exhibited by the year-union interaction coefficients in the first column reverses. The estimates in column 2 imply that the union-nonunion wage differential would have declined 3.5 percentage points over the period if industry demand had not fallen, and the influences of trend and rising inflation would have held union real wages essentially constant.

The precise magnitudes of the estimates of the separate influences of inflation and unemployment cannot be regarded with a great deal of confidence, because the estimates are probably sensitive to the choice of procedure used to forecast industry employment, E_{it}. Nevertheless, the general character of the results is consistent with the hypothesis that nominal union wages are less sensitive to short-run disturbances than nominal wages of nonunion workers. The response of nonunion wages to industry demand shifts is admittedly small, but it is highly significant, and the response of union wages to these shifts is a fraction of the nonunion response. Furthermore, if the trends in union and nonunion wages can be assumed to be identical, then the pattern of the year dummy coefficients in column 2 indicates that the rise in the growth of consumer prices in the year following May of 1973 had a larger and more persistent effect on real union wages than real nonunion wages. This suggests that nominal union wages responded more slowly to unforeseen changes in inflation than nominal wages of nonunion workers.

In their paper using a similar approach to capture the effect of cyclical shifts on wage rates, Smith and Welch note that if the employment residuals do in fact capture cycles in cross section, then the qualitative characteristics of the estimates from the pooled sample should be retained when the model is estimated over individual year cross sections. The results of this exercise, presented in Table 6, are encouraging. Procyclic variation in the wages of both union and nonunion workers is observed in all three years, and the interaction coefficient is negative in each case. Thus, the prediction error does appear to be a moderately consistent measure of cyclical activity even in single-year cross sections.7/

IV. Employment and hours adjustments

The logit models reported in this section contain fewer explanatory variables than the wage regressions. The equations here include age and its square, years of schooling, value added per establishment, dummies for union membership and occupation, and the industry employment prediction error and its interaction with the union dummy. The models also contain a lagged prediction error--the average of the forecast errors for December through May of year t-1. This variable was added to capture the effect of changes in the stocks of employees that may have occurred in the recent past. Current utilization should be negatively related to past adjustments. For example, an exceptionally large amount of hiring in year t-1 should lead to a smaller fraction of employees working overtime and a larger fraction of employees on layoff in year t.

Tables 7 and 8 show estimated derivatives dp_i/dx_j-- where p_i is the probability of observing an individual in state i and x_j is independent variable j--for i=1 (on layoff in Table 7 and working overtime in Table 8), i=2 (unemployed via discharge in Table 7 and working fewer hours than normal in Table 8) and all j. In the employment status model, the estimates are used to compare the employed with the unemployed and also to compare those who expect recall with those who must find new jobs. In the hours adjustment model, the derivatives identify some of the characteristics associated with variation in the workweek and also captures any asymmetries in hours variation. Derivatives in a logit model are functions of the values of the probabilities, and the derivatives in the tables here were computed using values of p_i obtained by evaluating the logit functions. Sample means were used for all independent variables except the prediction errors, which were set to zero to simulate a steady-state. The values of these probabilities are listed at the bottom of each table. The Appendix contains the logit coefficient estimates and an outline of the procedure for computing logit derivatives.

A. Employment status results

The negative contemporaneous prediction error derivatives in Table 7 reflect the negative association between levels of residual employment and rates of industry unemployment. The absolute values of the derivatives indicate how a small reduction in residual employment would be distributed between union and nonunion unemployment and between layoffs and discharges. The negative interaction derivative for layoffs reveals that layoffs of union workers account for more of the reduction in residual employment than do layoffs of nonunion workers. The trivial interaction derivative in the discharge column indicates that a decline in residual employment contributes equally to discharge unemployment in the union and nonunion sectors. A comparison of the derivatives across equations shows that deviations from trend cause larger changes in layoff unemployment than discharge unemployment among both union and nonunion workers. These findings are generally consistent with expectations based on the considerations covered in Section I. Layoffs dominate cyclical employment variation,8/ and layoff unemployment among union members is more sensitive to between-year shifts in employment than nonunion layoff unemployment. The principal surprise it is the rather large contribution of discharges to cyclical variation in unemployment among union members.

Most of the remaining derivatives are also consistent with the previous discussion. The adjustment in the current year is negatively related to residual employment in the previous year. The layoff-discharge ratio is a positive function of establishment size, union membership, and schooling.9/ The schooling effect may reflect the positive correlation between levels of general and specific skills outlined by Oi. Older workers and those higher in the occupational hierarchy are less likely to be disemployed, but these factors have no effect on the layoff-discharge ratio.10/ The trivial relationship between schooling and layoff unemployment was unexpected, but it may be a consequence of the sample's homogeneity.

B. Hours adjustment results

In the hours adjustment estimates in Table 8, the signs of the contemporaneous prediction error derivatives reflect the positive relationship between residual employment and hours worked per week. The magnitudes of these derivatives indicate the size of changes in the percentages of employees working more and fewer hours than normal that accompany a 1-unit deviation of employment from trend. The union-prediction error interaction derivative in each column has the same sign as the contemporaneous prediction error derivative in that column, so employment deviations are associated with adjustments in hours worked for a larger percentage of employed union workers. This is true for decreases in hours worked as well as increases. Comparing derivatives across columns indicates that cyclical variation in the percentage of employees working overtime is much larger than cyclical variation in the percentage of employees working fewer hours than usual. Although this could have been anticipated from the figures in Table 2, the result is still somewhat surprising in view of the premium required for overtime.

The remaining estimates in the hours adjustment model also reveal some interesting information. Most notable is the negative relationship between hours in the current period and residual employment in previous periods; the influence runs in the anticipated direction, but it is unexpectedly strong. The asymmetric relationships between age, occupation, and hours adjustments are also curious; older workers and craftsmen are less likely to experience hours reductions, but the frequency of overtime work is apparently independent of these variables. The schooling derivatives suggest that workers with more skill are less likely to have their workweek shortened and more likely to work overtime, although the estimates for these variables could also indicate that skilled workers underreport "normal" weekly hours. The percentage of workers experiencing any adjustment in the workweek, either upward or downward, declines as establishment size increases. This may reflect limitations in the flexibility of operating large assembly lines, or it may arise from the more bureaucratic lines of authority in large plants.

C. Summary

The two logit models indicate that adjustments in labor stocks are an important component of manufacturers' strategies for meeting year-to-year variation in demand. Discharges contributed significantly to the 1973-75 rise in manufacturing unemployment, and the lagged prediction error derivatives reveal a significant relationship between adjustments in labor stocks in year t-1 and adjustments in utilization in year t.

Comparisons of the union and nonunion derivatives in Tables 7 and 8 provide limited support for the hypothesis that the smaller adjustments in the wages of union workers led to larger adjustments in their utilization. Both layoff unemployment and overtime rates among union workers were more responsive to the forces that produced the 1973-75 declines in residual employment in manufacturing industries. The comparisons reveal an incongruity, however. The union-nonunion differences in the changes in wages and hours were very large, while the employment adjustments in the two sectors were somewhat similar. The adjustment costs theory of wage rigidity does not indicate how much employment stability a given increase in wage flexibility buys, but the estimates in Table 7 suggest it is not much. Whether this is a sign that the stability of wages and utilization are not substitutes over cycles--or an indication that the tradeoff is between variation in wages and hours--cannot be determined. Nevertheless, the results suggest that adjustment costs are important determinants of cyclical hours variation.

V. Concluding remarks

The results of this analysis of the 1973-1975 period illustrate the features that have dominated most discussions of cyclical labor market behavior. Nominal wage growth was stable, even in the face of large changes in inflation and unemployment, and there is little evidence that employers adopted worksharing arrangements to reduce layoffs. On the other hand, the estimates suggest that the standard characterization can be overdrawn. Wage adjustments did occur, and on average they were larger in the industries facing larger employment adjustments. Furthermore, the workweek showed flexibility in the upward direction during the period of extremely high labor force utilization, and that flexibility dampened the subsequent employment adjustment when utilization fell sharply.

The data revealed two exceptions to conventional wisdom. The first is that hours adjustments appear to be more frequent and larger than is commonly believed to be the case. Table 2 shows that the small and stable mean for excess hours conceals considerable hours variation for individual workers. About 30 percent of manufacturing employees were not at work the usual number of hours during the May survey weeks in each of the three years, and the average departure from the norm was not trivial. A great deal of this variation is not cyclical, and the high proportion of employees working overtime in May 1975, when unemployment was near 9 percent, suggests seasonal factors and random fluctuations may be sources of hours adjustments equal in importance to cyclical activity.

The data also are inconsistent with the popular notion that reductions in nominal wage growth are achieved less easily than increases. Although the Nixon price control program may have restricted the response of wages to demand increases early in the period, the relationships between residual employment and wages are stronger for later years when employment in most of the industries was below trend. The rather weak response to the 1974 jump in the price level also undermines support for the proposition that the flexibility of wage growth is asymmetrical.

The results indicate that close examination of disaggregated data can reveal much about cyclical behavior of the labor market and about the role of labor unions. Conclusions about cyclical behavior from the evidence here must be tentative, because the data cover only the contraction phase of one business cycle. Two additional issues that could be approached along these lines are differences across industries in lags in wage responses to a change in the inflation trend and the effects of the aggregate economic conditions on an employer's response to industry demand shifts. Both issues were ignored in the preceeding analysis, but they could be readily incorporated in an analysis of similar data covering a longer period. Factors other than unionism probably affect wage flexibility, and an employer's response to a change in demand can be expected to depend on whether the employers in other industries are experiencing similar conditions. With regard to union-nonunion differences, the degree to which one will find the paper's evidence persuasive will depend on his evaluation of the analytical method. The approach has weaknesses--

potentially important within-industry variance is not captured--but evidence from which conclusions can be drawn more confidently requires detailed data on the behavior of individual employers. The estimates indicate that such data, if available, would show unionism has a pronounced effect on the response to demand variation.

FOOTNOTES

1. "Wage rigidity" refers to the situation in which changes in wages are "in the right direction but too small to make actual wages equal to their equilibrium counterparts," (Lewis (1963), p. 213).

2. Hall suggests that the wage rate might be viewed as an instalment payment to the employee when workers typically remain with one employer for a "long" period (on the order of ten years).

3. "Layoff" here refers to an employer-initiated separation that is temporary ex ante--the employer will attempt to recall the employee eventually, but he is free to seek another permanent job in the meanwhile. Employer-initiated separations that the employer intends to be permanent from the outset will be called "discharges." The BLS refers to workers in either situation as "job losers."

4. The value added variable, included to capture the effect of establishment size, was taken from the 1972 Census of Manufacturers. It was merged with the CPS data on a geographic (9 Census divisions) as well as industry basis. The union coverage rates, which refer to white, full-time production workers, were computed from the May CPS files for 1973-1976. The volatility measure was computed from Employment and Earnings data.

5. If the respondent did not reside in one of the most populous states, the geographic information in the CPS placed him in a state group rather than an individual state. If the group contained both right-to-work and other states, then the value assigned the right-to-work variable is the probability he lived in a right-to-work state, based on his two-digit industry.

6. Abowd and Ashenfelter have investigated this issue more thoroughly. They do find evidence of a wage premium for employment in jobs with anticipated layoff unemployment.

7. Single year cross-sections provide an attractive alternative to the conventional time series procedures for analyzing the impact of wage controls. In addition to affecting economy-wide average wages, controls could dampen the response to short run movements in demand. Applying this reasoning to the results in Table 6, where the power of the prediction error is smaller in 1973 than in 1975, is tempting, because the 1973 wages may have been influenced by the final phases of the Nixon price control program. But the lack of data covering the period before and during the Nixon controls makes any inferences using this approach highly suspect.

8. Lilien found unemployment among job changers contributes more to increases in manufacturing unemployment during recessions than temporary layoffs, but he defines temporary layoffs as those separations that actually end in recall, whereas here a separation is considered to be temporary if the employee anticipates recall.

9. The significance of the union dummy is considerably higher when the establishment size variable is omitted.

10. Although one might anticipate that seniority would offer a union worker more protection than a nonunion worker, the data do not support that hypothesis. Addition of interactions of age and its square with the union dummy increased the overall chi square statistic by 2.1, which is insignificant at the 10 percent confidence level.

REFERENCES

Abowd, John M., and Orley Ashenfelter. "Anticipated Unemployment, Temporary Layoffs, and Compensating Wage Differentials." NBER Conference Paper No. 37, June 1980.

Ashenfelter, Orley, George E. Johnson, and John Pencavel. "Trade Unions and the Rate of Change in Money Wages in U.S. Manufacturing Industries." Review of Economic Studies 34 (January 1972): 27-54.

Azariadis, Costas. "Implicit Contracts and Underemployment Equilibria." Journal of Political Economy 83 (December 1975): 1,183-1,202.

Baily, Martin Neil. "Wages and Employment under Uncertain Demand." Review of Economic Studies 41 (January 1974): 37-50.

____________. "Contract Theory and the Moderation of Inflation by Recession and Controls." Brookings Papers on Economic Activity (3:1976): 585-622.

____________. "On the Theory of Layoffs and Unemployment." Econometrica 45 (July 1977): 1,043-1,063.

Barro, Robert J. "Long-term Contracting, Sticky Prices, and Monetary Policy." Journal of Monetary Economics 3 (July 1977): 305-316.

Becker, Gary S. "Investment in Human Capital: A Theoretical Analysis." Journal of Political Economy 70 (Supplement: October 1962): 9-49.

Calvo, Guillermo A., and Edmund S. Phelps. "Employment Contingent Wage Contracts," in Stabilization of the Domestic and International Economy, pp. 160-168. Edited by Karl Brunner and Allan H. Melter. Carnegie-Rochester Series on Public Policy (A Supplementary Series to the Journal of Monetary Economics), Vol. 5, 1977.

Dunlop, John T. Wage Determination Under Trade Unions. New York: Augustus M. Kelley, Inc., 1950.

Feldstein, Martin S. "The Importance of Temporary Layoffs." Brookings Papers on Economic Activity (3:1975): 725-744.

____________. "The Effect of Unemployment Insurance on Temporary Layoff Unemployment." American Economic Review 68 (December 1978): 834-846.

Fischer, Stanly. "'Long-term Contracting, Sticky Prices, and Monetary Policy': A Comment." Journal of Monetary Economics 3 (July 1977): 317-324.

Freeman, Richard B. "The Effect of Unionism on Worker Attachment to Firms." Journal of Labor Research 1 (Spring 1980): 29-62.

Gordon, Donald F. "A Neo-Classical Theory of Keynesian Unemployment." Economic Inquiry 12 (December 1974): 431-459.

Hall, Robert E. "Employment Fluctuations and Wage Rigidity." Brookings Papers on Economic Activity (1:1980): 91-142.

Hall, Robert E. and David M. Lilien. "Efficient Wage Bargains under Uncertain Supply and Demand." American Economic Review 69 (December 1979): 868-879.

Klein, Benjamin, Robert G. Crawford, and Armen A. Alchian. "Vertical Integration, Appropriable Rents, and the Competitive Contracting Process." Journal of Law and Economics 21 (October 1978): 297-326.

Lilien, David M. "The Cyclical Pattern of Temporary Layoffs in United States Manufacturing." The Review of Economics and Statistics 62 (February 1980): 24-31.

Lewis, H. Gregg. "Competitive and Monopoly Unionism," in The Public Stake in Union Power, pp. 181-208. Edited by Philip D. Bradley. Charlottesville: University of Virginia Press, 1959.

________________. Unionism and Relative Wages in the United States. Chicago: University of Chicago Press, 1963.

Medoff, James L. "Layoffs and Alternatives under Trade Unions in U.S. Manufacturing." American Economic Review 69 (June 1979): 380-395.

Mitchell, Daniel J. B. Unions, Wages, and Inflation. Washington, D.C.: The Brookings Institution, 1980.

Moore, William J. and John Raisian. "Cyclical Sensitivity of Union/Nonunion Relative Wage Effects." Journal of Labor Research 1 (Spring 1980): 115-132.

Morton, Walter A. "Trade Unionism, Full Employment, and Inflation." American Economic Review 40 (March 1950): 13-39.

Oi, Walter Y. "Labor as a Quasi-Fixed Factor." Journal of Political Economy 70 (December 1962): 538-555.

Parsons, Donald O. "Models of Labor Market Turnover: A Theoretical and Empirical Survey," in Research in Labor Economics, pp. 185-224. Edited by Ronald G. Ehrenberg. Greenwich, Connecticut: JAI Press, 1977.

Pierson, G. "The Effect of Union Strength on the U.S. Phillips Curve," American Economic Review, 58 (June 1968), 456-467.

Raisian, John. "Cyclic Patterns in Weeks and Wages." Economic Inquiry 17 (October 1979): 475-495.

Rees, Albert. "Some Non-Wage Aspects of Collective Bargaining," in The Public Stake in Union Power, pp. 124-142. Edited by Philip D. Bradley. Charlottesville: University of Virginia Press, 1959.

______________. "Postwar Wage Determination in the Basic Steel Industry." American Economic Review 41 (June 1951): 389-404.

Rosen, Sherwin. "Short-Run Employment Variation on Class-I Railroads in the U.S., 1947-1963." Econometrica 36 (July-October 1968): 511-529.

______________. "Trade Union Power, Threat Effects and the Extent of Organization." Review of Economic Studies 36 (April 1969): 185-196.

Smith, James P. and Finis Welch. "Local Labor Markets and Cyclic Components in Demand for College Trained Manpower." Annales de l'INSEE 30-31 (April-September 1978): 599-630.

Theil, Henri. "A Multinomial Extension of the Linear Logit Model." International Economic Review 10 (October 1969): 251-259.

Williamson, Oliver E., and Michael L. Wachter, "Obligational Markets and the Mechanics of Inflation." Bell Journal of Economics 9 (Autumn 1978): 549-571.

Williamson, Oliver E., Michael L. Wachter, and Jeffrey E. Harris. "Understanding the Employment Relation: The Analysis of Idiosyncratic Exchange," Bell Journal of Economics 6 (Spring 1975): 250-278.

TABLE 1

LAYOFF AND DISCHARGE UNEMPLOYMENT IN MANUFACTURING

	Percent Unemployed Due to:	
	Temporary Layoff	Permanent Discharge
Nonunion		
1973	0.7	1.7
1974	1.0	1.4
1975	4.6	6.2
Union		
1973	0.7	1.1
1974	1.6	0.9
1975	7.0	3.2

Notes: Population restricted to hourly white male production workers who were either employed or job losers.

SOURCE: May Current Population Surveys.

TABLE 2

EXCESS HOURS IN MANUFACTURING

	Mean	Standard Deviation	Percent Positive	Mean of Positives	Percent Negative	Mean of Negatives
Nonunion						
1973	1.51	14.1	23.0	18.7	13.2	-21.1
1974	0.18	14.2	19.4	17.4	13.5	-23.7
1975	0.34	14.5	17.6	19.8	13.5	-23.3
Union						
1973	2.87	15.2	25.3	21.1	10.9	-22.7
1974	2.04	13.7	20.8	20.7	10.5	-21.6
1975	0.02	13.0	14.2	20.0	13.2	-21.6

Notes: All figures are percentages. Population restricted to white male hourly production workers.

SOURCES: May Current Population Surveys.

TABLE 3

STATISTICS ON INDUSTRY VARIABLES

A. ALL MANUFACTURING (73 INDUSTRIES)

	Mean	Standard Deviation	Minimum	Maximum	Number Positive	Mean of Positives	Mean of Negatives
Prediction Errors:							
1973	.135	.125	-.075	.612	68	.149	-.048
1974	.070	.117	-.449	.464	59	.100	-.056
1975	-.111	.111	-.345	.194	11	.069	-.143
Index of Stochastic Employment Variability	.660	.882	.012	5.30			
Value Added per Establishment ($10M)	.264	.421	.001	3.99			
Union Coverage	.519	.162	.095	.859			

B. MAJOR SUBSECTORS

	Durables (38 Industries)		Nondurables (35 Industries)	
	Mean	Standard Deviation	Mean	Standard Deviation
Prediction Errors:				
1973	.194	.135	.072	.074
1974	.092	.148	.047	.064
1975	-.153	.111	-.065	.092
Index of Stochastic Employment Variability	.981	1.07	.312	.407
Value Added per Establishment ($10M)	.275	.437	.250	.403
Union Coverage	.539	.180	.498	.139

SOURCES: Employment and Earnings
Census of Manufacturers
May Current Population Surveys

TABLE 4

STATISTICS ON INDUSTRY VARIABLES AFTER ATTACHMENT TO CPS RECORDS

	Mean	Standard Deviation	Percent Positive	Mean of Positives	Mean of Negatives
A. NONUNION WORKERS					
Prediction Errors:					
1973	.165	.129	95	.176	-.052
1974	.071	.114	78	.106	-.056
1975	-.137	.108	11	.063	-.163
Index of Stochastic Employment Variability	.665	.856			
Value Added per Establishment ($10M)	.243	.440			
Union Coverage	.484	.151			
B. UNION WORKERS					
Prediction Errors:					
1973	.151	.122	95	.162	-.045
1974	.077	.105	78	.108	-.028
1975	-.125	.105	12	.060	-.152
Index of Stochastic Employment Variability	.728	.785			
Value Added per Establishment ($10M)	.563	.830			
Union Coverage	.613	.155			

SOURCES: Employment and Earnings
Census of Manufacturers
May Current Population Surveys

TABLE 5

ESTIMATES FROM WAGE EQUATIONS
POOLED SAMPLES FOR 1973-75

	1	2	3
Prediction Error [a_1]	--	.269 (7.7)	.201 (8.6)
(Prediction Error)xD(U) [a_2]	--	-.207 (4.6)	-.144 (4.7)
[a_1+a_2]	--	.062 (2.1)	.057 (2.9)
D(1974)	-.021 (2.1)	.003 (0.3)	--
D(1975)	-.046 (4.8)	.035 (2.4)	--
D(1973)xD(U)	.071 (3.3)	.099 (4.2)	--
D(1974)xD(U)	.078 (3.6)	.087 (3.9)	--
D(1975)xD(U)	.100 (4.5)	.064 (2.9)	--
D(U)(Union Dummy)	--	--	.082 (4.0)
ln(Demand Variability Index) [b_1]	.009 (2.3)	.004 (1.0)	.005 (1.3)
[Ln(Demand Variability Index)]xD(U) [b_2]	-.014 (-2.9)	-.01 (2.1)	-.012 (2.4)
[b_1 + b_2]	-.005 (1.7)	-.007 (2.2)	-.007 (2.3)
R^2	.3523	.3565	.3559
Standard Error of Estimate	.2384	.2376	.2377

Number of Observations: 9,596
Mean of Dependent Variable: 1.075

Notes: Absolute t-ratios in parentheses. Remaining coefficients in Table A-1.

SOURCES: Employment and Earnings, Census of Manufacturers, and May Current Population Surveys

TABLE 6

ESTIMATES OF COEFFICIENTS OF
PREDICTION ERROR IN WAGE EQUATIONS:
SEPARATE SAMPLES FROM 1973-75

	1973	1974	1975
Prediction Error [a_1]	.328 (4.8)	.270 (4.3)	.449 (6.1)
(Prediction Error)xD(U) [a_2]	-.228 (2.4)	-.152 (1.9)	-.325 (3.5)
[a_1+a_2]	.100 (1.6)	.118 (2.2)	.124 (2.2)
R2	.3382	.3627	.3754
Partial R2 (both variables)	.0073	.0079	.0133
Standard Error of Estimate	.2398	.2312	.2413
Mean of Dependent Variable	1.09	1.07	1.06
Number of Observations	3292	3226	3078

Notes: Absolute t-ratios in parentheses:
These equations also contain the demand variability index and all the variables listed in Table A-1.

SOURCES: <u>Employment and Earnings</u>
<u>Census of Manufacturers</u>
May Current Population Surveys

TABLE 7

DERIVATIVES FROM EMPLOYMENT STATUS MODEL

	(1) Layoff	(2) Discharge
Age	-.047 (6.8)	-.050 (7.2)
Schooling	-.011 (0.3)	-.100 (3.1)
D(Craftsman)	-.489 (2.8)	-.387 (2.2)
Ln(VAE)	.252 (4.1)	-.202 (3.0)
D(U)	.340 (1.7)	-.330 (2.0)
Lagged Prediction Error	1.54 (2.1)	1.57 (2.1)
Contemporaneous Prediction Error [a_1]	-6.63 (8.0)	-4.24 (6.7)
(Contemporaneous Prediction Error) x D(U) [a_2]	-2.46 (2.4)	-1.64 (0.2)
[a_1 + a_2]	-9.09 (14.3)	-4.41 (6.2)
Steady State Probability	.013	.012

Notes: Derivatives have been multiplied by 100.
VAE is value added per establishment.
Absolute t-ratios in parentheses.

SOURCES: See Table A-2.

TABLE 8

DERIVATIVES FROM HOURS ADJUSTMENT MODEL

	(1) Overtime	(2) Short Time
Age	-.010 (0.2)	-.162 (5.8)
Schooling	.563 (3.0)	-.362 (2.7)
D(Craftsman)	.319 (0.4)	-1.27 (1.9)
Ln(VAE)	-.737 (2.1)	-.717 (2.7)
D(U)	-.212 (0.2)	-.535 (0.8)
Lagged Prediction Error	-15.6 (4.1)	3.9 (1.4)
Contemporaneous Prediction Error [a_1]	11.5 (2.9)	-1.05 (3.8)
(Contemporaneous Prediction Error) x D(U) [a_2]	12.9 (2.5)	-4.35 (1.2)
[$a_1 + a_2$]	24.3 (7.4)	-5.4 (2.1)
Steady State Probability	.220	.115

Notes: Derivatives have been multiplied by 100.
VAE is value added per establishment.
Absolute t-ratios in parentheses.

SOURCES: See Table A-3.

APPENDIX

The multinomial logit model, introduced by Theil, is of the form

$$(1)\quad \ln \frac{P_i}{P_k} = B_i X + \qquad (i = 1, \ldots, k\text{-}1),$$

where the dependent variable assumes integer values from 1 through k. X is the vector of independent variables, P_i is the probability that the dependent variable is equal to i, and the B_i are vectors of coefficients with elements b_{ij}. The probabilities must sum to 1, so the individual P_i and the independent variables are related by

$$(2)\quad P_i = \begin{cases} 1/(1+{}^{k-1}_{\;1}S_m) & \text{if } 1 = k \\ P_k S_i & \text{if } i < k, \end{cases}$$

where

$$S_i = \exp\,(B_i X) \qquad (i = 1, \ldots, k\text{-}1).$$

The derivative of P_i with respect to the j^{th} independent variable (denoted P_{ij}) is then

$$(3)\quad P_{ij} = P_i(b_{ij} - {}^{k-1}_{\;1}P_i b_{ij}).$$

Sample frequencies are commonly used for the probabilities P_i. A set of steady-state probabilities is desired here, however, and these can be obtained from equations (2) by assigning suitable values to the independent variables and using the logit coefficients reported in Tables A-2 and A-3. To approximate the steady-state, the prediction error variables are set to 0 and the remaining variables are set to their means. The establishment size mean is 0.62, the mean age is 38, and the mean years of schooling is 12.

The age derivative for both groups and the contemporaneous prediction error derivative for union workers required computation of linear combinations. The form depends on the value of age, and the derivatives reported in Section IV are for a worker 38 years old. The contemporaneous prediction error derivatives for union workers are computed using the sums of the logit coefficients of this variable and the logit coefficients of the interaction with the union dummy.

TABLE A-1

COEFFICIENTS OF REMAINING VARIABLES FROM WAGE REGRESSIONS[a]

Schooling	.027 (25.0)
Age	.033 (19.7)
Age Squared	-3.62×10^{-4} (17.4)
D(Never Married)	-.054 (4.3)
D(Married, Spouse Present)	.044 (4.3)
D(SMSA > 1 million)	.038 (5.5)
D(SMSA < .25 million)	-.067 (10.6)
D(Northeast)	-.097 (11.7)
D(North Central)	-.021 (2.7)
D(South)	-.085 (8.5)
D(Laborer)	-.077 (8.2)
D(Operative)	
D(RTW)[b]	-.064 (5.8)
D(RTW) x D(U)	.059 (4.9)
% Union	.113 (4.1)
(% Union) x D(U)	.102 (2.8)
Value Added per Establishment	.085 (8.8)
(Value Added per Establishment)xD(U)	-.065 (6.0)

Notes: [a]These are the coefficients of the remaining variables from the regression reported in column 3 of Table 5. Absolute t-ratios in parentheses.

[b]D(RTW) is a dummy whose value is 1 if the individual resides in a state that outlaws the union shop.

TABLE A-2

LOGIT COEFFICIENTS FOR EMPLOYMENT STATUS MODEL

	Layoff	Discharge	Chi-Square
Constant	-.313 (0.4)	1.42 (1.8)	--
Age	-.176 (4.7)	-.192 (4.7)	42.6
Age Squared	1.83×10^{-3} (3.8)	1.97×10^{-3} (3.7)	26.7
Schooling	-9.58×10^{-3} (0.3)	-.083 (3.2)	10.0
D(Craftsman)	-.387 (2.8)	-.327 (2.3)	12.7
D(U)	.263 (1.7)	-.271 (2.0)	7.0
Ln(VAE)	.196 (4.0)	-.166 (2.9)	26.0
Lagged Prediction Error	1.22 (2.1)	1.32 (2.2)	8.7
Contemporaneous Prediction Error	-5.24 (8.1)	-3.60 (6.8)	106.3
(Contemporaneous Prediction Error)xD(U)	-1.93 (2.4)	-.162 (0.2)	5.7
Sample Proportions	.026	.022	--

Number of Observations: 12,044
Overall Chi Square: 606.82

Notes: VAE is value added per establishment. Figures in parentheses are absolute t-ratios.

TABLE A-3

LOGIT COEFFICIENTS FOR HOURS ADJUSTMENT MODEL

	Overtime	Short Time	Chi-Square
Constant	-2.06 (6.1)	-.462 (1.1)	--
Age	.038 (2.3)	-.027 (1.4)	8.8
Age Squared	-5.38×10^{-4} (2.6)	1.35×10^{-4} (.5)	7.9
Schooling	.029 (2.5)	-.028 (2.1)	13.4
D(Craftsman)	.0001 (0.0)	-.125 (1.9)	3.8
D(U)	-.021 (0.4)	-.058 (0.9)	0.8
Ln(VAE)	-.055 (2.7)	-.084 (3.2)	14.9
Lagged Prediction Error	-.883 (3.9)	.158 (0.6)	17.3
Contemporaneous Prediction Error	.678 (2.9)	.066 (0.2)	8.5
(Contemporaneous Prediction Error)xD(U)	.713 (2.4)	-.248 (0.7)	6.8
Sample Proportions	.204	.122	--

Number of Observations: 10,700
Overall Chi Square: 163.84

Notes: VAE is value added per establishment. Figures in parentheses are absolute t-ratios.