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EMPLOYMENT VARIATION AND WAGE RIGIDITY: A COMPARISON OF UNION AND NON-UNION PLANTS

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Abstract

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I. INTRODUCTION

Does employment vary more over time in union than in non-union plants? The question is of fundamental importance to macroeconomic theories that explain unemployment as a result of rigid wages, typically holding the unionized sector up as the prime example.

All companies face some element of fluctuation in their product demand. When product demand falls, the derived demand for labor will also usually fall. The firm can adjust by laying off workers, by cutting wages, or by work-sharing, (hours reduction) among other ways. Where the price mechanism is frozen, one might expect greater fluctuation in quantities.

This paper compares employment variability and turnover at union and non-union plants. It asks whether employment is more unstable across years in union plants compared to non-union plants. Does this reflect differences in product demand volatility, in employer policy, or in union preferences? Do the differences correspond with those predicted by the monopoly model or the efficient contract model of unionism?

To implement these tests of labor demand at a disaggregated level, this paper uses new data on a panel of manufacturing plants between 1969 and 1981. Unlike previous analysis, I shall be directly concerned here with the second moments of individual observations: the variance over time in employment (labor demand) within plants. This will then be split into cyclical and residual components.

The following section draws hypotheses concerning employment variation from two models of industrial unions -- the monopoly union model and the efficient contract model, and compares these with what would be expected from a spot market in labor. The third section reviews some of the literature on union effects on employment and layoffs. Section IV presents some basic comparisons of employment variation and the transience of change by union status. Section V presents the main results comparing the intertemporal variation in employment across union and non-union plants. The likelihood of identifying a true union effect is strengthened first by comparing white-collar workers with blue-collar workers within plants, and second by estimating individual union effects. The sixth section is an aside that develops the implications of slower employment growth in union plants for explaining the decline in the proportion of the work force organized. Section VII examines the channels of differential employment variation, comparing termination, new-hire, and promotion rates at union and non-union plants. The eighth section presents conclusions, followed by an appendix that describes the data-set constructed for this study.

II. WAGE-SETTING, AND EFFICIENT CONTRACTS MODELS OF UNIONS

These are two major ways to think about unions that carry relevent and divergent implications for employment stability. See Oswald, 1985, for a review of these models. The first is a standard monopoly model of unions in which the union sets a fixed wage for the duration of the contract. Employment is then determined by the firm's labor demand schedule. Such rigid-wage models have been commonly used as a basis for upward sloping aggregate supply schedules. This standard model yields the testable hypothesis that employment variation is amplified in the union compared to non-union sector because of relative wage rigidity. This assumes that wages decline if labor demand falls in non-union plants. A finding of no greater employment variation is subject to at least the following additional interpretations, all still within the context of the monopoly model: 1) for some reason (such as efficiency wages or long-term implicit contracts), wages are also rigid in the non-union sector; 2) firms have foresight in their selection of products, markets and output contracts. A risk averse unionized firm realizes it is constrained by rigid factor prices, so selects itself into customer contracts with fixed output prices for positive output for a matching period. This is analogous to financial market models in which exposure to price level fluctuations is minimized by matching credit and debit maturities. 3) National unions have foresight. They select stable plants for union organization.

The second major model of unions arises from the efficient contracts literature. To take a polar case, think of the union and the company sharing rents.[1] It is in their joint interest to maximize the pool of rents available for sharing, and so to use inputs with first-best efficiency (Hall & Lilien, 1979). The quantity of labor used in production will then be identical to that called for by competitive market wages. Contract wages are then just an accounting device, and their presumed rigidity has no implication for employment variation. Since the union share of rents is taken as a pure income transfer, this strong-form efficient contract model yields the testable hypothesis that employment variation does not differ because of the presence of a union.

With homogenous union members, strong form efficiency (rent maximization) requires union members to be indifferent concerning the level of employment.

Income redistribution from employed to unemployed union members would be sufficient to accomplish this, but this is implausible because institutionally such direct redistributive measures are rarely observed (Oswald, 1984). Exceptions are found in the supplemental unemployment benefits contained in collective bargaining agreements negotiated by the United Auto Workers and the United Rubber Workers. A second more direct exception may be found in pension benefits, especially under early retirement schemes adopted during recessions.

Union members are, however, not homogenous. In particular, they typically have (endogenous) well-defined seniority rights. Oswald (1984) uses the institutional fact that layoffs are typically in order of reverse seniority (last-in first-out) to argue that efficient contracts are on the demand curve because the median union member will be indifferent to higher levels of employment. The dynamic properties of this model are unclear -- some deus-ex-machina long-term contract enforcement mechanism is needed to keep successive median voters from marching up the demand curve until they leave only the Cheshire Cat smile on the faces of the last two union members who sell out the third for higher wages. An alternative approach makes use of the fact that recalls are also commonly by seniority, so if steady-state employment is constant and capital markets work, only the present value of life-time earnings and leisure matters, and union members are indifferent toward short-run fluctuations in employment.

III. PAST STUDIES OF UNION EFFECTS ON EMPLOYMENT STABILITY

The impact of unions on employment stability, and hence on unemployment, appears to be a balance between competing forces. Temporary layoffs are relatively more prevalent in the union sector compared to the non-union sector during cyclical down-turns because unions restrict the use of alternative adjustment mechanisms such as wage cuts or work-sharing and because the quit rate is generally lower under unionism. Medoff (1979) has calculated layoff rates in the union sector of manufacturing to be 2.2 to 4.6 times greater than in the non-union sector, and layoff unemployment 1.5 to 1.6 times greater. On this basis we would expect employment variation to fluctuate more violently in the union sector, in response to a given level of product demand variation. There is also some evidence that product demand variation itself is greater in the union sector of manufacturing. Freeman and Medoff (p. 113) cite unpublished work that estimates that the variation in shipments over the business cycle is nearly twice as large in heavily unionized as in lightly unionized industries. In contrast to the earlier work by Medoff, the more recent analysis by Freeman and Medoff is reported to find no significant difference in the response to a given change in shipments.

Under certain restrictive conditions, the mean tenure of employed workers may also contain information on the history of employement fluctuations within plants, although the connections are quite complex. For example, if workers never quit and were always discharged in reverse seniority order (last in, first out), and if plants experienced no net growth or shrinkage, then one can infer, although still in a limited sense, that a plant with a higher average tenure among its workers has experienced less employment variation. Freeman (1980) reports higher average tenure among union workers, but the precise interpretation depends on other factors such as those discussed above.

There is also evidence that permanent layoffs are no more prevalent among union than non-union workers and that unemployment due to permanent layoff is lower among union workers (Freeman and Medoff, 1984, pp. 118-120). This surprising finding will be explored further here.

The interpretation of the results to be presented here will depend critically on how rigid union wages are relative to non-union wages, and on how both of these compare to a spot market. Abowd and Card (1984, Table 3) report a set of statistics from which one can derive the variance of the annual change in the logarithm of earnings in longitudinal samples. In samples from both the PSID and the NLS, they find 24 to 38 percent greater variation in the change in wages of non-union than of union workers between 1967 and 1979. Freeman and Medoff (1984, p. 272, n. 5) report elasticities of real wages with respect to real shipments of .02 in unionized industries and of .12 in non-union industries in quarterly data between 1958 and 1975. Both studies agree in finding relatively rigid wages in the union sector, but the Freeman and Medoff results indicate that wages respond inelastically to demand shifts in both sectors.

Empirical tests of efficient contract models have begun to develop in the last few years. Most have examined the behavior of one union, the typographers. The results so far have been imprecise, mixed, and heavily dependent on auxiliary identifying assumptions. See Ashenfelter and Brown (1983), Card (1984), Carruth and Oswald (1983), Dertouzos and Pencavel (1981), Farber (1978a, 1978b), MaCurdy and Pencavel (1983), Martinello (1984), Pencavel (1983a, 1984b), and Svejnar (1982). Nevertheless, it seems fair to conclude from these studies that generally trade unions act as if both employment and wages mattered. (See Pencavel (1985) for a review.) Against this we must hold the finding (Oswald, 1984) that few unions admit to having either formal or informal agreements with employers concerning employment levels.

The overall picture suggested by this literature is of unions that are not indifferent about employment levels, and of a union sector with relatively greater temporary layoffs, relatively fewer quits, but comparable permanent layoffs. This paper will focus on union effects on long-term employment variability, as well as on promotions, new hires, and total terminations. Employment may vary more in the union sector, either because product demand is more variable, or more likely because alternative responses to a given product demand shock are relatively restricted in the union sector. The next section attempts to distinguish between these two explanations without the aid of direct measures of product demand variability or of real wage rigidity.

IV. EMPLOYMENT VARIATION IN UNION AND NON-UNION PLANTS

Is employment more stable in non-union plants than in union plants? The evidence to be presented in this section offers some support for this hypothesis, although many of the differences are small. This section presents some basic evidence of size, industry and transient effects.

California manufacturing plants are observed over 5 consecutive years in rolling time-frames between 1969 and 1981. The standard-deviation of total employment within a plant across time is 714 in union plants, greater than the 521 observed in non-union plants (see Appendix for description of data). Only plants with positive total employment for five years are included in the sample, so employment variation due to births or deaths is not measured here. If the type of variation union plants face is more likely to drive them out of business, evidence drawn from this sample may give a downward biased picture of total employment change in union plants.

Mathematically, the standard deviation increases with scale ($VAR(\lambda x) = \lambda^2 VAR(x)$), so these measures are contaminated with scale effects. Since the union plants average 1539 employees over these years, compared to 1246 in the non-union plants, the (scale-free) coefficient of variation is also marginally greater in the union plants, .46 compared to .42. The coefficient of variation is rather clumsy to use, so the remainder of this paper shall use the logarithm of size. This has two advantages. First, it is more natural to think of plants changing in equal proportion to their size than in equal absolute numbers. Second, $VAR(\log(\lambda x)) = VAR(\log x)$ where λ is a constant, so the logarithm measure is scale free.

Variation in employment within union plants compared to non-union plants may be examined at a basic level by calculating the correlation of employment within plants over time. Table 1 presents such correlation matrices of the logarithm of plant size over 5 consecutive years separately for union and non-union plants. This table shows that employment variation has been slightly greater in union plants than in non-union plants within the manufacturing sector of one state. The correlation of plant sizes in neighboring years ranges from .97 to .99 in non-union plants. It is nearly identical in union plants, .97 to .98. The differences are also negligible over 5 years. The correlation between (early) year 1 and (recent) year 5 employment is .93 among non-union plants, compared to .92 among union plants. Over the same 5 years absolute employment in the non-union plants grows by 17%, compared to only 1% in the union plants,[2] a finding we shall return to.

The median and mode period of observation is from 1972 (year 1) to 1976 (year 5) in both the union and non-union subsamples. The distribution is shown in Table 4. Obviously, as one would expect with any sort of business cycle, the correlation matrix is not stationary but rather the correlations vary by calendar year. For example, the lowest first order correlation is found between year 2 and year 3 -- corresponding to 1973 and 1974 -- in the union sub-sample (but not in the non-union subsample). Note also that the higher order correlations are close to integer multiples of the first order correlations, as would be predicted by a first-order autoregressive process. This is inconsistent with a simple fixed effect model which would predict identical correlations for every pair of years. The usefulness of past plant size for predicting future size decays as time passes.

Although union plants grow slower than non-union plants, they appear to experience slightly greater employment variation. As time passes, the correlation of current employment within plants with employment in any given past year declines, and it declines faster in union plants than in non-union plants. The slower growth of union plants may be due to higher wages. The greater employment variation may well reflect more rigid wages in the union sector, since it is unlikely that the variability of product demand differs greatly across union and non-union plants within the manufacturing sector in the one state considered here.

Stronger empirical support for this view could be developed by directly measuring product demand variation in the two sectors. An alternative approach examines highly disaggregated sectors whose products are perhaps more likely to face highly correlated demands across plants. Within a detailed industry, for a specific region and year, do we still observe greater variation in union plants? Table 2 replicates the correlation matrices of Table 1 for residuals from regressions of the logarithm of plant size on industry, region, and calendar year. This helps to isolate the union effect. The correlation between plant size 5 years apart is .87 in the union plants, less than the .93 observed in non-union plants. Within disaggregated industries, regions, and periods of time, the variation of employment over time is greater in union than in non-union plants. One would also like to see other characteristics of these plants controlled for and other measures of employment variation, as we shall turn to in the next section. It is possible, but unlikely, that this difference in employment variation could be explained by greater product demand variation faced by the union plants in this sector. It is more likely that this greater employment variation reflects either 1) more rigid wages and staffing procedures in the union sector that result in greater employment variation for a given product demand variation, or 2) the effects of omitted plant characteristics, such as size.

Do the changes in plant size observed here largely reflect trend growth or transient shocks and measurement error? There appears to be little trend in growth rates across years relative to transient shocks, in the sense that plants with above average growth in one year don't enjoy similar success in later years. The correlation matrices of growth rates across time within plants in the union and non-union subsamples are presented in Table 3. At best, half of the correlations are significant, and all the significant correlations are negative. The best bet after seeing a plant grow is that it will shrink (relative to group average). (See Leonard (1984b) for further evidence.) There is some suggestion in Table 3 that union plants take longer to turn around.

This new evidence comparing employment volatility in the union and non-union sectors also suggests that past evidence indicating higher tenure for union than for non-union workers may tell only part of the story. How can it be true that union workers have higher tenure than non-union workers, yet employment undergoes greater fluctuations in union plants? The data on tenure refers only to uncompleted spells of employment. A conceptually distinct measure of tenure would compute the average duration of employment for all workers who had been employed in a plant. While this distinction has been fruitfully applied in the past to unemployment durations (see Salant, Kaitz, Akerlof and Main), it has not been applied to the union tenure effect. While it may well be true that the average worker currently employed in a union plant has greater tenure than his non-union counterpart, it may also be true that the average completed spell of employment observed in the union sector is shorter than its non-union counterpart. The workers who are not permanently displaced by the greater employment variation we observe in the union plants remain on their jobs and accumulate the higher tenure usually observed. The workers who are displaced are not counted in these uncompleted spell measures of tenure, but, by the nature of union seniority clauses, would contribute to lower completed spell measures of tenure.

V. REGRESSION TESTS OF THE UNION EFFECT ON EMPLOYMENT VARIATION

The results of the last section have the weaknesses as well as the strengths of simplicity. What little additional employment variation there is at union plants is of a nature that may not readily be explained by greater variation in demand for the products of unionized plants. The weakness of such simple tests is their failure to control for other differences between union and non-union plants -- size in particular. This section presents regressions of two different measures of intertemporal employment variation on a vector of detailed plant characteristics and first finds that total employment in union plants, measured from year to year, is generally as stable as in non-union plants, *ceteris paribus*. A second measure splits this variation more finely into cyclical changes and a residual, and finds less cyclical and residual variation in the union sector. This effect is then compared across white-collar and blue-collar workers, and across individual unions.

The first problem is to disentangle the union growth effect from the union variation effect. A plant in steady growth or decline will show a greater raw variance than one with constant employment. Since union plants have lower growth rates than non-union plants, it is desirable to compare the variation in employment about trend to differentiate from trend growth effects. Let S_{it} be the logarithm of total employment in plant i in year t. The raw variance is of course:

(1) $$V_i = \Sigma_{t=1}^{5}(S_{it} - \bar{S}_i)^2/4$$

Now consider the N time-series regressions:

(2) $$S_{it} = \alpha_i + \beta_i T_t + e_{it} \qquad i = 1,N \quad t = 1,5$$

where T = time trend [-2, -1, 0, 1, 2], and the errors are assumed to be independently and identically distributed (no serial correlation). Detrended size is simply the residual from this regression, or:

(3) $$\tilde{S}_{it} = S_{it} - [\hat{\alpha}_i + \hat{\beta}_i T_t]$$

Note that an additional degree of freedom for each observation has been used up in detrending. The variance of detrended employment is then:

(4) $$\tilde{V}_i = \Sigma_{t=1}^{5}\tilde{S}_{it}^2/3$$

The square root of this gives a measure of the mean proportional detrended change.

Union plants do differ from non-union plants in a number of major aspects besides unionism itself, as Table 4 shows for the California manufacturing sample. Detrending should reduce the variance of non-union plants more than that of union plants because union plants' unweighted annual average growth rate is only 3.7%, compared to the non-union 5.6%. Also of particular concern here, the sample union plants are more than half again as large as the non-union plants. This is important because past studies of employment variation suggest that changes are proportional to absolute size. With such scale effects, the variance of the logarithm of size need not increase with size even though the variance of absolute size does. The union plants are also more likely than non-union plants to be part of larger multi-establishment companies. The proportions of craftsmen, operators, laborers and service workers, and of clerical workers are higher in union plants. The union plants are relatively more likely than non-union plants to be found in the two major SMSAs of Los Angeles and San Francisco than in outlying, more recently developed areas. The union plants are also more likely to be found in the following industries: printing, glass, machinery, electrical equipment (except SIC 367), and transportation equipment. This sample includes more than 700,000 employees, or more than a third of all the employees in California manufacturing. The effects of shocks that are common across plants such as the business cycle, may be interpreted in the context of the market, rather than the partial equilibrium of a single plant. In both sectors, the median and mode period of observation is from 1972 to 1976. This period includes the turbulent time after the first "oil shock" -- a classic period in which to examine the effects of rigid wages.

Direct Tests of Variation

Table 5 presents regressions of the raw and detrended intertemporal variation in the annual logarithm of total employment within plants on a vector of individual plant characteristics including union status, size and its square at initial year of observation, annual growth rate, corporate structure, occupational structure, 8 dummies for period of observation, 17 dummies for 2 or 3 digit SIC industry and 5 dummies for region.

There is no significant difference in Table 5 in the employment stability of union and non-union plants. Measured in terms of year to year variation in total employment, union plants -- despite their presumed greater wage rigidity -- are just as stable as non-union plants. By this measure, the presumed freezing of the price adjustment mechanism by unions does not result in greater quantity fluctuation.

The natural result that size changes tend to be in proportion to size can be seen in Table 5 from the finding that the coefficients on size and its square are insignificant. Growth still adds strongly and significant to variance, even to the variance of the detrended logarithm of employment.

Most of the other plant characteristics have little significant impact. In contrast to other findings in a national sample, but as predicted by models with perfect capital markets, corporate structure makes little difference. As one would expect from Oi-type models, non-clerical, white-collar employment is relatively stabilized compared to blue-collar employment. The only significant industry or region effects (not shown) are higher variances in transportation equipment, and in San Jose (a notably febrile market).

What of the formidable oil shock itself? Does the battering it gave the U.S. economy show itself in greater employment instability within plants after 1973? Curiously, there is no strong evidence of extraordinary post-shock turbulence here. Rather, the periods of greatest employment instability are the observations from 1969 to 1973, and from 1977 to 1981 (second oil shock). Within plant employment variation actually declines significantly in between. This is both unexpected and unexplained. If this finding is not a fluke, it may carry implications for thinking about the microeconomic response to macroeconomic shocks.

Pooled Time-Series Cross-Section Results on Change in Logarithms

An alternative specification fits more comfortably into the time-series cross-section framework and also allows for a more finely divided examination of the the different cyclical sensitivities in union and non-union plants. Like all random walk models, of which this is a variant, this process implies that the variance of the size distribution increases without limit as time passes. A convenient specification to be estimated separately for union and non-union plants is:

$$d_{it} = \alpha + b_1 DGNP_t + b_2 DGNP_{t-1} + b_3 \bar{S}_i + \tilde{Z}_i \tilde{b} + e_{it} \tag{5}$$

where

$d_{it} = S_{i,t} - S_{i,t-1}$ $\qquad t = 2, 5$

S_{it} = logarithm of employment in plant i in year t

$DGNP_t$ = change in log of real gross national product between year t and year t-1.

$\bar{S}_i = (\Sigma_{t=1}^{5} S_{i,t})/5$

$\tilde{Z}_i$ = A vector of plant characteristics in the initial year of observation (or time invariant), including industry, region, occupational and corporate structure.

e_{it} = error term, assumed serially and cross-sectionally uncorellated, and normally distributed with mean zero.

This is a regression, using pooled time-series cross-section data, of the first difference in the logarithm of size (the logarithm of the annual growth rate) of the i'th plant in between year t and year t-1 on a set of detailed plant characteristics. These controls include the mean logarithm of size, a set of industry and region dummies, controls for occupational and corporate structure, and cyclical indicators.

This equation yields a test on comparative variation when estimated separately for the union and non-union subsamples, by inspection of the residual variances. If the detrended level of union employment is more variable, then so must be the growth rate, and so the residual variance from a regression of the logarithm of growth rates on other plant characteristics. If the standard error of the estimate is substantially larger in the union subsample, the inference drawn is that union employment is more volatile than non-union employment. Of course, proof by residual is among the weakest of proofs, so little weight should be given to such evidence.

Relative cyclical sensitivity is tested by comparing the response across sectors of changes in plant level employment to changes in real gross national product. Suppose the true model were:

$$d_{it} = \alpha + \delta_1\left(\frac{DGNP_t - DGNP_{t-1}}{2}\right) + \delta_2\left(\frac{DGNP_t + DGNP_{t-1}}{2}\right) + b_3\bar{s}_i + \tilde{z}_i\tilde{b} + e_{it} \tag{6}$$

Here plant level employment growth is a function of both the trend growth in GNP ($DGNP_t$+$DGNP_{t-1}$) and the change in GNP growth ($DGNP_t$-$DGNP_{t-1}$), which is taken here as indicating the business cycle. The coefficients of interest are then identified from estimates of equation 5 as:

$$\delta_1 = (b_1 - b_2)/2, \text{ and} \tag{7}$$

$$\delta_2 = (b_1 + b_2)/2 \tag{8}$$

If employment is more cyclically sensitive in union than non-union plants, the estimate of δ_1 in the union subsample should exceed that in the non-union subsample.

The pooled time-series cross-section regression takes the first difference in the logarithm of size, and so differences out plant-specific time-invarient unobserved variables. One might also consider a distinct measure of the within plant variance of size (see Griliches and Hausman for general discussion):

$$(9) \qquad S'_{it} = S_{it} - \bar{S}_i$$

which takes each plant's annual size as a deviation from the plant's own mean. With T = 2, the two specifications are equivalent. For T > 2, d_{it} allows plant and year specific trend growth rates (or plant dummies with drift) while S'_{it} imposes a constant plant-specific trend across time. Because in this sense it is less restrictive, I shall present results for the first difference; although in some sense it may over-correct. The results on cyclical sensitivity do not differ substantially in the deviation from plant mean specification.

Suppose S_{it} were modelled with error V_{it}. Then since d_{it} is the first difference of S_{it}, $e_{it} = V_{it} - V_{i,t-1}$. This first differencing may be expected to remove most serial correlation that may have been present in the V_i's. Assuming, stationarity, $\bar{S}_i$ is then uncorrelated with the error term, because

$$(10) \qquad \mathrm{COR}(e_{i,2} - e_{i,1} \; , \; \frac{e_{i,2} + e_{i,1}}{5}) = 0$$

Note that because of regression to the mean (Leonard, 1984) we would expect an upward biased coefficient on size if initial year size were used as an independent variable. Here, such transient error is averaged out by grouping across time, and so does not bias our estimates of the effect of size on growth.

Prominent arguments about structural shocks are almost always presented in terms of differences across industries, with industries categorized at the 2 or 3 digit S.I.C. level. (Lilien, 1982). Because of data limitations, intra-industry variation is rarely discussed, but it seems reasonable to assume that as a matter of definition, an industry shock is one that is widely shared within the industry. Equation 5 tests for this by including a set of dichotomous variables indicating 2 or 3 digit SIC industry. If growth rates across plants within industry are highly correlated, or if inter-industry variation in growth is greater than intra-industry variation in growth, then the industry dummies should differ significantly from each other and should jointly significantly reduce the residual variance in growth rates. Otherwise, it follows that the structural change commonly measured across industry is just the tip of the iceberg, and that there is much greater within industry variation submerged beneath sight in most studies.

Additional controls such as these help to ensure that I am estimating a union rather than an industry or size effect, but are also of interest in their own right. By the law of large numbers, one would expect a lower variance at larger plants. If, for whatever reason, conglomerates cross-subsidize plants, or enjoy an economy of conglomeration (for example, company- as opposed to plant-specific skills), then plants that are part of multi-plant corporations would be expected to exhibit less variation than single plants. Given technological change, (or a presumed U.S. comparative advantage in skill intensive production) human capital intensive plants (with a high proportion of managerial and professional workers relative to their industry) may enjoy higher growth rates.

A major advantage of the first-difference specification is that it permits an inspection of individual year effects that are obscured in the moving average representation of Table 3. Equation 5 includes year-specific contemporaneous and lagged GNP growth rates to measure the impact of macroeconomic cycles at the microeconomic level of the plant.

There are then two hypotheses to be examined here concerning employment variation at union and non-union plants. The first is whether union plants are more cyclically sensitive. This is tested by comparing the response to changes in GNP growth rate across union and non-union subsamples. The second, and logically distinct, question is whether the residual variation from the equation as a whole is greater or less in the union compared to the non-union sector.

Results

The results in Table 6 indicate comparatively less cyclical sensitivity among union plants. The changes from year to year in annual growth rates, while not always significant, exhibit complex timing patterns (not shown here) that differ across sectors. The coefficient of cyclical sensitivity, δ_1, is .83 in the union sample. This is less than the estimated 1.13 among non-union plants, although the difference is not significant. The major finding of this paper is then that employment in union plants is no more cyclically sensitive than in non-union plants. This result is consistent with the strong-form efficient-contract model of unions.

Conditional on business-cycle effects and other plant characteristics, residual employment variation is less in union (S.E.E. = .290) than in non-union (.315) plants. Given the magnitude of these standard-errors and the correspondingly low R^2s, this could also be taken as showing roughly equal ignorance of the determinants of plant-size variation in both sectors.

There are no significant industry effects in either sector. There is no significant common component of plant growth rates across plants within an industry. Variation is far more substantial within than across industries. If one believes that structural change across industries are an important economic phenomenon, this finding indicates that the (typically hidden) changes within industry may be of even greater importance.

Contrary to the popular folk wisdom, there is no statistically significant relationship between mean size and growth. In particular, the plants with the smallest 5 year average size do not grow any faster than larger plants. This result differs from the folklore primarily because measurement error and other sources of transient variation have been averaged out of the size variable on the righthand side here. The implication is that previous claims that growth occurs disproportionately among the small depend on the statistical artifact of regression to the mean in a world with commonplace transient shocks (Leonard 1984b). It is also worth noting that growth rates do not differ significantly whether or not a plant is part of a larger corporation.

From the perspective of macroeconomic policy, one might fear that the burden of a recession falls most heavily on small plants that are more likely to be credit constrained. This is tested here by splitting the sample by mean size. The results must be qualified by the fact that the sample is of plants, not firms, and by the fact that plants that go out of business are not included in the sample. The largest third of plants do show less total and residual variation (S.E.E. = .196) than the smallest third (S.E.E. = .233). Cyclical sensitivity bears a non-linear relation with size. It is signficantly greater in the mid-sized (365 to 811 employees) plants.

In both sectors, human capital intensive plants grow faster (within industry), and this effect is significant in the union sector. The larger the proportion of blue-collar or clerical workers in a plant's workforce, the slower its growth.

Unobservables: White-Collar Workers in Union Plants

All empirical work is subject to the qualification that unobserved or uncontrolled for variables may cause bias in the estimates. Here, one may reasonably wonder whether union and non-union plants really are being compared "all other things held fixed". Perhaps union plants face less cyclical product markets than non-union plants, even within the 2 or 3 digit SIC industries considered here. If so, some doubt would be raised considering the true nature of what I have called a union effect here. Without additional information, an unobservable must always remain an unobservable, a chronic but largely untreatable condition.

The white-collar workers in a unionized plant are typically not themselves unionized. This institutional regularity provides us with identifying information. If the differences previously observed between union and non-union plants were due to some unobserved factor, one might reasonably expect that factor to affect white-collar workers as well as blue-collar workers. A distinguishing test is then to ask whether the differences found between union and non-union plants for blue-collar workers are also found for white-collar workers. This is in the spirit of a "brothers" test that differences out a common within "family" unobservable. (Chamberlain and Griliches, 1975)

When the regressions of Table 6 are estimated for white-collar workers (managers, executives, professionals, technicals, and sales workers) residual variation is slightly larger for workers in union (S.E.E. = .217) than in non-union (.214) plants. More importantly, cyclical sensitivity is greater for white-collar workers in union (δ_1 = .645) than in non-union (δ_1 = .443) plants. Opposite results were found among blue-collar workers. Table 7 summarizes these results.

Table 8 presents a stronger variant of a "brothers" test that uses matched comparisons. The dependent variable here is the difference within an individual plant between the logarithm of the blue-collar employment growth rate and the logarithm of the white-collar employment growth rate. The white-collar logarithmic growth rate is used here as a control for plant specific demand shocks. If the previous results were an artifact arising because union plants were concentrated in less cyclical markets, then the blue-collar white-collar difference should have no cyclical relationship in non-union plants, and a negative cyclical relationship in union plants.

Table 8 shows the opposite. In union plants, δ_1 = .36 and δ_2 = -.58, while in non-union plants, δ_1 = 1.36 and δ_2 = 1.14. Union plants exhibit less cyclical sensitivity than do non-union plants in the difference between logarithmic growth rate of blue-collar workers and that of white-collar workers. In other words, conditional on changes in white-collar employment, blue-collar employment variation is, if anything, relatively dampened in union compared to non-union plants.

Whatever omitted variables there may be appear to work in the opposite direction to the union effect. It is difficult to think of an omitted plant characteristic that produces less cyclical sensitivity and less residual variation for blue-collar workers in union plants compared to non-union plants, but at the same time produces greater cyclical sensitivity and more residual variation for white-collar workers in the very same plants. The lower cyclical sensitivity of blue-collar workers in union compared to non-union plants is not due to the location of union plants in less cyclical markets because white-collar workers in union plants are more cyclically unstable than their counterparts in non-union plants. This should considerably strengthen confidence in the interpretation of the findings for blue-collar workers as revealing a true union effect.

Individual Union Effects

An important part of the efficient contracts literature has been concerned with examining the wage and employment decisions of a single union -- the typographers. These same models would lead one to expect that different decisions are made by different unions. This section takes advantage of the fact that one union typically bargains collectively in a number of different industries to attempt to identify union effects. The existence of such individual union effects is in doubt because of the considerable autonomy exercised by locals of some unions, and the heterogeneous preferences of various locals, as revealed for example in Pencavel (1984b).

Union A has collective bargaining agreements in 74 sample plants spread across 7 of the industry groups used here. It is known for its fiery national leadership and independent locals. Compared to other plants, whether unorganized or organized by other unions, the plants organized by union A are a model of stability. The standard error of the estimate of equation 5 (comparable to Table 6) is .235, and cyclical sensitivity (δ_1 = .76) is less than the union average. (See Table 7.) Industry effects should be captured by industry dummies in all estimates. If these are fine enough, this result may be interpreted as indicating differences between the employment goals union A pursues in bargaining and those pursued by other unions.

Union B has the reputation of an aggressive, independent, and autocratic union. It represents workers in 39 sample plants in 6 industries. It shows significantly greater cyclical sensitivity than the union average, and slightly less residual variance (S.E.E. = .259). Union C, on the other hand, exhibits greater residual variance (S.E.E. = .383) but no significant cyclical sensitivity. It bargains for workers in 39 sample plants across 7 industries.

These differences across individual unions are at least as great along some dimensions as the difference between union and non-union plants. They may reflect differences across unions in distribution across detailed (3+ SIC) industries finer than those controlled for here. Otherwise, they indicate substantial differences in the weight different unions give to employment stability in bargaining. In the sense of there being on average no significant difference in cyclical employment variation compared to non-union plants, there is some evidence here supportive of the efficient contracts theory.

VI. DIFFERENTIAL EMPLOYMENT GROWTH WITHIN PLANTS AND THE DECLINE OF UNIONS

Employment grew by 17% in non-union plants compared to 1% in union plants over 5 year periods in the 1970s. Note that these are not the unweighted average of growth rates across plants (which are larger as shown in Table 4, because of faster growth at smaller plants in both sectors) but rather the growth rate of average employment within a fixed sample of plants in each sector. The differential growth rate in employment within plants is of fundamental importance in explaining the decline of private sector unions (Dickens and Leonard; Leonard 1984B). Union plants grow significantly and substantially slower than non-union plants, controlling for size, industry, region, period of observation, corporate structure and occupational structure. Table 9 presents a regression of the logarithm of the growth rate (the logarithm of the ratio of terminal to initial year employment) on a vector of detailed plant characteristics. Over five years, union plants grow 12.5 percentage points slower than their non-union counterparts.

In California manufacturing, the State Department of Industrial Relations reports that union members fell from 597,400 in 1973 to 555,500 in 1977, or from 35.8 percent of all manufacturing wage and salary workers to 31.9 percent. The implied number of non-union wage and salary workers increased from 1,071,315 to 1,185,879. By applying the employment growth rates of 17%

(non-union) and 4.5% (union = 17% - 12.5%) to the 1973 employment totals, the change in union density that could be accounted for by differential growth rates can be estimated. The imputed and actual (1977) employment levels are 624,238 (555,500) union and 1,253,439 (1,185,879) non-union. In other words, if union and non-union employment in California manufacturing had grown at the same rates we observe in the study sample, the proportion of union workers would have declined to .332 by 1977. Since unionization actually declined to .319, two-thirds of the decline in union density in California manufacturing between 1973 and 1977 could be accounted for by the slower growth of employment in union plants relative to non-union plants.

There are a number of qualifications surrounding this result. First, the growth rates are calculated in longitudinal samples and so do not include births or deaths. Second, the plants in the longitudinal sample are larger than the average plant and so may differ in growth rates. Third, union status is as of 1982, and fourth, not all the employees in plants with collective bargaining agreements are union members. Fourth, union plants may always have grown slower than non-union plants, in which case one must look for other explanations, such as the decline in organizing, for the difference between the growth in the proportion of the workforce unionized in the early 50's, and its subsequent decline.

While the remaining qualifications, and other such qualifications, could reduce the magnitude of the employment growth differential, they are unlikely to alter the qualitative conclusion: differences in the growth rate of employment across existing union and non-union plants are of fundamental importance in understanding why the proportion of the manufacturing workforce represented by unions has declined so markedly in recent years.

Discussion

Unions do not cause any greater employment variability, measured from year to year. Unions are associated with slower employment growth. Unions do, according to other studies, cause greater use of temporary, but not permanent, layoffs. These three results would seem to be essential elements to be explained by a model of union decision making. The strong-form efficient-contract model, as has been stressed here, is consistent with the first result. The second and third results raise interesting questions to which full answers have not been attempted here. One possibility is that part of what had appeared as a union effect in previous studies of temporary layoffs is an artifact of the correlation of (observed) union status with (unobserved in previous work) plant size. It is also important to note that the cost to a union of temporary layoffs is much less than that of permanent layoffs. The income loss is buffered by unemployment insurance, and risks must be considered in light of the high (70%) rate of recall (Lilien, 1980; Katz, 1985) -- an implicit contract that unions may help to enforce. But if unions find permanent layoffs so costly (and so avoid excess cyclical sensitivity), wouldn't they also avoid the slower employment growth also observed here? The crucial distinction here is between the costs of shrinking employment (borne in part by union members), and those of slower employment growth (diffused among the mass of unidentifiable potential union members). The unions in the study sample have avoided generating excess permanent layoffs over the business cycle, as well as those caused by shrinking workforce.

VII. NEW HIRES, TERMINATIONS, AND PROMOTIONS IN UNION AND NON-UNION PLANTS

Many observers would expect slower growth and greater variation about that growth in union plants. Does it then follow that new hire and promotion rates are lower, and termination rates higher in union plants?

Table 10 compares the rates of new hires, terminations and promotions across union and non-union plants in California manufacturing during the late 1970s. In each case, the rate is higher among the non-union plants. For example, the average rates of new hires during the preceding 12 months to last year's total employment is .32 among the non-union plants compared to .25 among the union plants, and the difference is significant. The termination rate is .32 among non-union plants, significantly greater than the .27 rate observed among union plants. The promotion rate is slightly higher in non-union plants, .14 compared to .13, but the difference is not significant.

The first aspect of these numbers that deserve consideration is their sheer magnitude (see Parsons for a review of previous work). Terminations and new hires equivalent to roughly a third of all non-union workers and a quarter of all union workers occur annually. For comparison, the Bureau of Labor Statistics (1980, p. 16) reports an average monthly new hire rate of .031 and non-layoff separation rate of .030 in manufacturing during 1978. Multiplying by 12 gives rough annual rates of about 36 or 37 percent, larger than those observed here. The complete explanation of this large volume of turnover remains a challenge, but it may look larger than it really is. First, it is not surprising that the new hire and termination rates are of similar magnitude since it is plausible that newly hired workers are the most likely to be fired. Second, we could easily observe huge termination and new hire rates in a stable workforce. Consider the case of 365 new hires and terminations in a plant with 100 employees, 99 of whom never turnover, while the 100th position is newly hired each morning and fired each afternoon. Nevertheless, it is notable that turnover in the union sector is of the same order of magnitude as in the non-union sector even though the union sector is often thought of as having a far more stagnant and stable workforce.

The second notable aspect of the data presented in Table 10 is that the union sector appears, subject to the provisos noted above, to have more stable employment. While termination and new hire rates within each sector are of roughly equal magnitude, suggesting roughly stable total employment, these rates are both at least a fifth higher in non-union compared to union plants. While part of this difference may be due to the omission of temporary layoffs and recalls from the data, the initial reading would suggest that union plants have significantly less turnover.

These differences in turnover rates across union and non-union plants, however, are not explained by the effect of unions themselves so much as by the differences in other characteristics between union and non-union plants. As Table 4 showed, there are some major differences between union and non-union plants found in a single state. Union plants are significantly more likely than are non-union plants to be part of larger multi-establishment companies. They are also significantly larger and significantly slower growing than are non-union plants. The proportion of craftsmen, operators, laborers and service workers is higher in union plants. In both sectors, the observations on turnover are primarily made in the years 1975, 1976 and 1977, during the recovery from the 1974 recession.

These differences between union and non-union plants in size, corporate structure, industry and region are apparently of greater importance than union status itself in explaining the differences in turnover rates between union and non-union plants we observe in Table 10. Table 11 presents results from regressions of turnover rates within plants on a set of detailed plant characteristics. These include 17 dichotomous variables indicating two or three digit SIC industry, 5 dichotomous variables indicating geographic area within California, and 7 dichotomous variables indicating year of observation. The size, growth rate, and proportion of blue collar workers within each plant are controlled for. Indicators for the union status of the plant, and for whether or not the plant is part of a multi-establishment company are also included.

Union plants still have lower rates of hiring and termination, but the differences are not significant once other plant characteristics are controlled for. In particular, plant size, growth, and corporate structure have strong and significant effects on turnover that would have appeared as union effects in Table 10. Union plants appeared to have lower turnover rates in Table 10 in part because union plants are larger, slower growing, and more likely to be part of a multi-plant corporation -- all characteristics associated with lower turnover. Growth is related by an accounting identity to the difference between hires and terminations. Note also that differences in turnover between union and non-union plants are still insignificant when the regressions in Table 10 are repeated not controlling for establishment growth rates.

These results for hiring and termination rates are roughly consistent with Section V's results on annual employment variation. When employment variation is measured in proportionate terms so as to remove pure scale effects (as in, e.g., the variance of the logarithm of employment), employment is subject to no greater variation in union plants than in non-union plants. Here, once size and other plant characteristics are controlled for, hire and termination rates do not differ significantly across union and non-union plants. In the case of both annual total variation and of turnover, we find a non-linear relation with size.

The size effects are, however, of greater significance in explaining turnover. The importance of those size effects contrast with the insignificance of union effects. The "web of rules" governing employee relations in large formalized work places may be of no less importance than the collectively bargained rules in the union sector that have drawn so much more attention. To some degree, personnel practices in large non-union plants reflect a union threat effect, but the remainder represents a largely unexplored institution.

How do these plant level findings compare with earlier results from studies of individual workers? In maximum likelihood estimates of the probability of turnover (job change) among young men, Farber (1980, p. 45) finds that turnover rates are not significantly affected by the union status of the job. In contrast, Freeman (1980, p. 653, 658) finds in NLS and PSID samples that individual separation rates are significantly lower in union jobs, and that this arises not because of the unionization of innately more stable workers, but because of the changed behavior of workers under unionization. Caution must be exercised here because the individual and establishment level results are not directly comparable. In particular establishment level separation rates that are identical across union and non-union plants need not reveal sharply different separation rates among various subgroups of employees, such as the young for example. The exact relationship will depend on complicated sorting schemes that are not directly discernible from the data at hand. A full reconciliation of the establishment level results observed here with earlier individual level results has not been attempted, but it is apparent that other establishment characteristics in addition to unionization play an important role in determining turnover.

Table 11 also presents some interesting empirical support for Walter Oi's classic theory of labor as a quasi-fixed factor of production. The theory predicts that workers in whom greater amounts of firm specific human capital are invested will experience lower turnover rates over the business cycle. In particular, workers in occupations with relatively high hiring or training costs will be less likely than workers with less "fixity" to be laid-off in response to a decline in demand. Table 11 shows that within individual plants and controlling for year, industry, region, corporate structure, union status, size and growth rate; both termination and new hire rates are significantly and substantially greater in plants with a higher proportion of blue-collar workers. These workers are primarily semi-skilled and unskilled operatives and laborers for whom Oi's theory predicts the relatively higher turnover rates observed here.

Note also that new hire, termination and promotion rates are all significantly greater in plants that are not part of a larger multi-establishment company, and that new hire and termination rates both fall and then rise with increasing plant size. Growth contributes to greater termination rates, which is surprising, and to greater new hire rates, which is not. The former effect may arise from a heavy concentration of terminations among the newly hired.

This section has found that, <u>ceteris paribus</u>, union plants do not differ significantly from their non-union counterpart in terms of new hire, promotion and termination rates.

VIII. CONCLUSIONS

This paper has empirically analyzed a number of aspects of employment variability, turnover, and growth comparing union and non-union plants in the California manufacturing sector between 1969 and 1981. The main findings here include the following:

(1) Employment is less cyclically sensitive in union than in non-union plants, although the difference is not significant. Differences in total or residual intertemporal variation are not substantial across these sectors. Different unions appear to give different weight to employment stability in bargaining.

(2) Termination and new-hire rates are greater in the non-union sector. However, this is not due to unionization itself, but rather because union plants are larger, slower growing, and more likely to be part of a multi-plant company. Once these other factors are controlled for, there is little difference between union and non-union plants in termination and new-hire

rates.

(3) Employment has grown at a slower rate within the union sector. This is consistent with models in which unions raise labor costs, and in which unionized plants are at a competitive disadvantage. This factor alone can explain much of the decline in the unionized proportion of the manufacturing workforce.

Keynesian theories of unemployment generally depend on rigid wages in the labor market. The union sector is typically taken as a prime example, and perhaps the dominant locus, of such wage rigidity. Employment may well be more variable in union plants over shorter time periods than those considered here. This paper has shown that neither total nor residual intertemporal employment variation is greater in union than non-union plants, and that union plants are no more cyclically sensitive from year to year. The finding that employment varies no more in union than in non-union plants is consistent with (but cannot prove) the efficient contract model of decision making under collective bargaining. This suggests a provocative concluding question: could it be that, despite unemployment and cyclical disturbances, the economy is closer than commonly realized to embodying part of the central element of a share economy, and precisely in the unionized sectors where it is least expected? Of course, if employment in all sectors were set according to spot-market wages, the aggregate supply curve would be vertical and there would be no unemployment. This is difficult to believe.

The main empirical finding directly observed here is that employment varies no more from year to year in union than in non-union plants. We can reject the null hypothesis that, in terms of annual employment variability, union and non-union plants differ. The interpretation of this result depends critically on how one assumes the non-union market operates. If employment in the non-union plants is determined as in a spot market, then the efficient contracts hypothesis may have some validity. Alternatively, the finding here of little difference between union and non-union plants may indicate that employment and wages are set in long-term contracts in the large manufacturing plants considered here. In other words, in terms of employment variation union plants may look like non-union plants not because union plants set employment at spot market levels like non-union plants, but rather because non-union plants set the terms of employment under long-term contracts like union plants. In either case, it is clear that year-to-year employment variation is not greater in union than in non-union plants.

Appendix: The Plant Level Data Set

This study examines a detailed data set on union status and plant level turnover and employment patterns. The union status of each plant in the sample was determined by examining the 1982 collective bargaining contract collection of the California State Department of Industrial Relations. The Department has more than 3,400 private-sector agreements on file, and makes intensive efforts to obtain all contracts covering 50 or more employees. In 1982 this file included 1,364 contracts in the manufacturing sector, covering 450,310 employees. Since unions never achieve contracts in many plants in which they are certified as exclusive bargaining agents (see Dickens and Leonard), only plants with collective bargaining agreements will be referred to as unionized in this paper.

The coverage of this file is extensive, especially for contracts covering more than 50 employees. According to the U.S. Department of Labor there were 2,001,000 employees in California manufacturing in 1980. (Employment and Training Report of the President, 1980, table d-2, p. 230). Applying the 1977 California average of 35 percent non-production workers in manufacturing yields 1,300,650 production workers. (U.S. Census of Manufactures, 1977, Vol. III, Geographic Area Series-California, Part I, Table 2b, pp. 5-8). In a pooled 1973-1975 CPS sample of 6022 private-sector production workers in California, Freeman, and Medoff estimate the proportion unionized at .35, close to the national average of .36 (Freeman and Medoff, 1979, p. 166, Table 4). Nationally, Freeman and Medoff report that 49% of production workers in manufacturing were union members. On this basis, we would expect to find 637,320 union members among production workers in California manufacturing.

Since the Freeman-Medoff comparison is in terms of union members, we must translate our data on contract coverage into union members. 88 percent of all employees covered by collective bargaining agreements covering at least 100 workers in California manufacturing are subject to union shop or modified union shop security clauses. (California Department of Industrial Relations, 1982, Table 1). If we pessimistically assume that none of the others are union members, we are left with (.88)(450,310) union members. This is 396,000 union members, or 62 percent of the number we would expect to find by applying the Freeman-Medoff estimates of percent unionized to BLS totals. Part of this discrepancy may be due to the striking decline in unionism in California. Union members as a proportion of all production workers in California manufacturing dropped from .56 in 1975 to .42 in 1979. (California Department of Industrial Relations, 1980, p. 2, Table 1). If we adjust Freeman and Medoff's 1973-1975 benchmark downwards by the same 25% to .37, then we would expect 481,240 union members in California manufacturing. The remainder are likely to be in establishments of less than 100 employees, which are excluded from the study sample. To the extent that some unionized establishments are still not identified as such, this measurement error will bias our results against finding any difference between the union and non-union sectors.

Information on new hires, terminations and promotions as well as data on total (first quarter) employment for a 5 year period is primarily available in the durable goods manufacturing industries, and in the larger plants, so the average size of a plant studied here is significantly larger than the average manufacturing plant. Of course, these large plants account for a disproportionate share of all manufacturing employment. Note also that

California, in industrial relations as in other spheres, may well not be representative of the nation as a whole. In particular, California is unique among the states in paying unemployment insurance to workers working fewer hours per week than the standard hours. We would expect to see more work-sharing through hours reduction and less layoffs than in other states. This data set is discussed at greater length in Leonard (1983b).

Notes

1. The efficient contract model used in this paper assumes that unions can redistribute income to members not currently employed, as is the case with supplemental unemployment benefits, for example. Fifty-one percent of all major contracts have formal income maintenance provisions specifying work guarantees, severance pay or supplemental unemployment benefits (BNA, 1983). Where this is not possible, the union will bargain directly over both wages and employment. (McDonald and Solow) The existence of efficient contracts has been analyzed in the past by testing the hypothesis that employment varies in response to opportunity wage, but not in response to own wage. (Ashenfelter and Brown, 1983; Card, 1984; Pencavel, 1984). From this perspective, the tests to be presented here use the business cycle and contemporaneous changes in non-union employment as proxies for changes in the opportunity wage. Models in which the union cannot redistribute income are more complex. Their implications for employment variation depend on the movement of the opportunity wage relative to the labor demand schedule, as well as on the union's preferences and risk aversion. In this weaker, but perhaps more realistic sense, contracts may be pareto-efficient even if greater employment variation is observed under unionism. The approach taken here is to first test for the strong form of first-best efficient contracts.

2. See Leonard (1984a) for further analysis of the union employment effect. See Leonard (1984b) for a detailed study of the dynamics of establishment size.

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Table 1: The Variability of Employment in Union and Non-Union Plants. Correlation Matrices of the logarithm of Plant Size Over Time.

Union Plants
N=245

Mean	Standard Deviation		Year 1	Year 2	Year 3	Year 4	Year 5
6.62	1.09	Year 1	1.000	.979	.956	.934	.922
6.64	1.06	Year 2		1.000	.972	.947	.929
6.66	1.06	Year 3			1.000	.979	.959
6.65	1.07	Year 4				1.000	.981
6.66	1.07	Year 5					1.000

Non-Union Plants
N=342

Mean	Standard Deviation		Year 1	Year 2	Year 3	Year 4	Year 5
6.23	1.06	Year 1	1.000	.971	.951	.940	.934
6.27	1.03	Year 2		1.000	.978	.955	.948
6.34	1.02	Year 3			1.000	.971	.960
6.35	1.04	Year 4				1.000	.986
6.36	1.07	Year 5					1.000

Note: The median and mode period of observation is from 1972 (Year 1) to 1976 (Year 5) in both subsamples.

Table 2: The Variability of Residual Employment in Union and Non-Union Plants. Correlation Matrices Over Time of Plant Size Within Industry, Region and Year.

Union Plants
N=245

	Year 1	Year 2	Year 3	Year 4	Year 5
Year 1	1.000	.960	.922	.882	.868
Year 2		1.000	.947	.904	.878
Year 3			1.000	.959	.924
Year 4				1.000	.964
Year 5					1.000

Non-Union Plants
N=342

	Year 1	Year 2	Year 3	Year 4	Year 5
Year 1	1.000	.964	.942	.928	.918
Year 2		1.000	.976	.950	.940
Year 3			1.000	.964	.949
Year 4				1.000	.982
Year 5					1.000

These are correlation matrices for the residuals from regressions of the logarithm of plant size on a vector of dichotomous variables for industry (13), region (5) and calendar year (8).

Table 3: The Transience of Change. Correlation Matrices of the Change in the Logarithm of Plant Size.

Union Plants
N=245

Mean	Standard Deviation		di2	di3	di4	di5
.022	.225	di2	1.000	-.105 (.10)	-.028 (.67)	-.137 (.03)
.013	.253	di3		1.000	.078 (.23)	-.058 (.37)
-.003	.216	di4			1.000	.047 (.46)
.000	.208	di5				1.000

Non-Union Plants
N=342

Mean	Standard Deviation		di2	di3	di4	di5
.039	.252	di2	1.000	-.025 (.65)	-.203 (.00)	-.019 (.73)
.066	.218	di3		1.000	-.106 (.05)	-.128 (.02)
.012	.248	di4			1.000	-.050 (.36)
.006	.180	di5				1.000

Note: Plus ca change, plus c'est la meme chose.
Figures in parentheses are probability of observing greater correlations under null hypothesis of zero correlation.
$d_{it} = S_{it} - S_{it-1}$ where S_{it} = logarithm of size of plant i in year t.

Table 4: Characteristics of Union and Non-Union Plants

Variable		Non-Union	Union
SIZE, initial year of observation		964	1535
$(SIZE)^2$		3,322,572	8,384,594
LOGARITHM OF SIZE, initial year of observation		6.23	6.62
MEAN-LOGARITHM OF SIZE, over 5 years		6.31	6.65
GROWTH RATE, annual average		1.056	1.037
% BLUE-COLLAR		.490	.570
% CLERICAL		.073	.106
SINGLE (Not Part of Multi-Plant Co)		.088	.049
Industry			
SIC20	Food	.003	.065
SIC26	Paper	.003	.004
SIC27	Printing	.012	.016
SIC28&29	Chemicals, Petroleum & Coal	.018	.008
SIC30	Rubber & Plastic	.012	.004
SIC33	Primary Metal	.044	.029
SIC34	Fabricated Metal	.181	.118
SIC35	Machinery	.132	.184
SIC36	Electrical Equipment (except 366 & 367)	.056	.106
SIC366	Communications	.050	.082
SIC367	Electronics	.275	.073
SIC37	Transportation Equipment	.137	.273
SIC38	Instruments	.050	.008
SIC39X	Miscellaneous Other Manufacturing: SICS 23 (Apparel), 24 (Lumber), 32 (Stone, Clay & Glass), 39 (Miscellaneous)	.029	.028
Geographical Area			
Los Angeles SMSA		.509	.604
San Diego, Anaheim, Riverside SMSA's		.269	.229
San Francisco SMSA		.038	.061
San Jose SMSA		.088	.024
Other Northern California		.018	.012
Other Southern California		.078	.070
Years of Observation			
1969-1973		.038	.029
1970-1974		.073	.102
1971-1975		.173	.208
1972-1976		.257	.212
1973-1977		.193	.180
1974-1978		.105	.122
1975-1979		.096	.094
1976-1980		.053	.049
1977-1981		.012	.044

Table 4 (Continued)

Beta Coefficient on Trend		
of Size	43.0	7.16
of Logarithm of Size	.033	.0075
Intertemporal Variance, Raw		
of Size	101,181	510,017
of Logarithm of Size	.048	.051
Intertemporal Variance, Detrended		
of Size	30,091	167,557
of Logarithm of Size	.020	.019
Number of Plants	342	245

Table 5: Regressions of the Intertemporal Variance of the Logarithm of Employment Within Plants on Plant Characteristics. N = 587

Equation	Raw			Detrended		
Variable	1	2	3	4	5	6
UNION	.0033 (.0096)	-.0027 (.001)	.0038 (.011)	-.0016 (.0034)	-.0036 (.0040)	-.0006 (.004)
SIZE		-5.99×10^{-6} (7.83×10^{-6})	-2.47×10^{-6} (7.69×10^{-6})		-4.24×10^{-6} (2.72×10^{-6})	2.64×10^{-6} (2.63×10^{-6})
$SIZE^2$		5.90×10^{-10} (6.84×10^{-10})	4.86×10^{-10} (6.70×10^{-10})		2.65×10^{-10} (2.37×10^{-10})	2.18×10^{-10} (2.29×10^{-10})
GROWTH		----	.205 (.041)		---	.093 (.014)
SINGLE		-.0028 (.020)	-.0027 (.020)		-.0050 (.007)	-.0050 (.007)
%BLUE-COLLAR		.044 (.036)	.054 (.035)		.025 (.012)	.029 (.012)
% CLERICAL		-.042 (.122)	-.024 (.119)		.013 (.042)	.022 (.041)
PERIOD 1970-1974		-.038 (.032)	-.065 (.031)		-.020 (.011)	-.023 (.011)
PERIOD 1971-1975		-.060 (.030)	-.071 (.029)		-.026 (.010)	-.031 (.010)
PERIOD 1972-1976		-.062 (.029)	-.068 (.028)		-.024 (.010)	-.027 (.010)
PERIOD 1973-1977		-.053 (.029)	-.052 (.029)		-.027 (.010)	-.026 (.010)
PERIOD 1974-1978		-.046 (.031)	-.056 (.030)		-.020 (.011)	-.024 (.010)
PERIOD 1975-1979		-.066 (.032)	-.075 (.032)		-.033 (.011)	-.037 (.011)
PERIOD 1976-1980		-.065 (.035)	-.076 (.035)		-.034 (.012)	-.039 (.012)
PERIOD 1977-1981		-.058 (.059)	-.067 (.058)		-.013 (.021)	-.018 (.020)
S.E.E.	.115	.116	.113	.041	.040	.039
R^2	.00	.04	.09	.00	.10	.16

Note: Equations 2, 3, 5, and 6 also include dichotomous variables indicating 2 or 3 digit SIC industry (17), and region within California (5). The omitted period is 1969-1973. Standard errors in parentheses.

Table 6: Cyclical, Sectoral, and Residual Employment Variation. Change in the Logarithm of Blue-Collar Employment by Union Status of Plant in Pooled Time-Series Cross-Section Samples

Sector:	Non-Union		Union	
Variable	Coefficient	(Standard-Error)	Coefficient	(Standard-Error)
$DGNP_t$	2.327	(.30)	.817	(.32)
$DGNP_{t-1}$	.061	(.31)	-.845	(.34)
Mean of Logarithm of Blue-Collar Employment	.016	(.01)	.015	(.01)
Single	.013	(.03)	.029	(.05)
SIC20	.060	(.17)	.016	(.07)
SIC26	-.138	(.17)	.023	(.16)
SIC27	.015	(.14)	.052	(.10)
SIC28 & 29	.033	(.08)	-.066	(.12)
SIC30	.003	(.09)	-.031	(.16)
SIC33	.064	(.07)	.061	(.08)
SIC34	.034	(.05)	.029	(.07)
SIC35	-.009	(.06)	-.022	(.06)
SIC361-365,369	.046	(.06)	.011	(.07)
SIC366	.018	(.07)	-.108	(.07)
SIC367	.033	(.06)	-.180	(.07)
SIC37	.011	(.06)	-.115	(.07)
SIC38	.067	(.07)	-.021	(.13)
LOS ANGELES	-.052	(.04)	-.062	(.04)
SAN DIEGO	-.022	(.04)	-.084	(.04)
SAN FRANCISCO	-.045	(.06)	-.005	(.06)
SAN JOSE	-.047	(.04)	.017	(.07)
NORTHERN CALIFORNIA	.030	(.08)	.065	(.09)

Table 6 (Continued)

Sector:	Non-Union		Union	
Variable	Coefficient	(Standard-Error)	Coefficient	(Standard-Error)
PROPORTION CLERICAL	-.154	(.25)	-.392	(.39)
PROPORTION BLUE-COLLAR	-.129	(.07)	-.252	(.08)
INTERCEPT	-.035	(.10)	.210	(.12)
R^2	.06		.06	
S.E.E.	.315		.290	
N	1360		976	
δ_1 (cycle)	1.13		.83	
δ_2 (trend)	1.19		-.01	

Note: Omitted dichotomous variables are the region other Southern California, and miscellaneous and other manufacturing, including SICs 23, 24, 32 and 39. Blue-collar workers are craftworkers, operators, laborers and service workers.

Table 7: Summary of Results on Total, Residual, and Cyclical Employment Variation

	N	Total Variation (= Standard Deviation)	Residual Variation (= Standard Error of Estimate)	Cyclical Sensitivity (= δ_1)
Blue-Collar Workers				
NON-UNION	1360	.322	.315	1.13
UNION	976	.295	.290	.83
UNION "A"	296	.245	.235	.76
UNION "B"	156	.267	.259	1.95
UNION "C"	156	.387	.383	.26
White-Collar Workers				
NON-UNION	1368	.217	.214	.44
UNION	980	.220	.217	.65
All Workers				
SMALL PLANTS (<365)	780	.236	.233	.78
MEDIUM PLANTS	692	.250	.241	1.37
LARGE PLANTS (≥ 812)	876	.198	.196	.62

Table 9: The Union Effect on Employment Growth. N = 587

UNION	-.125
	(.039)
SIZE	-3.76×10^{-5}
	(2.62×10^{-5})
$SIZE^2$	2.46×10^{-10}
	(2.29×10^{-9})
SINGLE	.008
	(.068)
% BLUE-COLLAR	-.264
	(.119)
% CLERICAL	-.349
	(.408)
S.E.E.	.388
R^2	.14

Note: The dependent variable is the logarithm of the ratio of employment in year t+4 to employment in year t.

This equation also includes dichotomous variables, indicating 2 or 3 digit SIC industry (17), region within California (5), and period of observation (8). Standard errors in parentheses.

Table 10: New Hire, Termination and Promotion Rates in Union and Non-Union Plants

	Non-Union Plants	Union Plants
1. NEW HIRE RATE	.323	.250
2. TERMINATION RATE	.322	.265
3. PROMOTION RATE	.145	.131
4. NUMBER OF PLANTS	411	287

Table 11: Regressions for New Hire, Termination, and Promotion Rates for all Employees in Union and Non-Union Plants N=558 Plants

	New Hire Rate	Termination Rate	Promotion Rate
UNION	-.016	-.0059	.0065
	(.020)	(.018)	(.0099)
SINGLE	.189	.176	.068
	(.039)	(.036)	(.019)
SIZE	-.000072	-.000061	-.0000061
	(.000016)	(.000014)	(.0000077)
$(SIZE)^2$	5.0×10^{-9}	4.3×10^{-9}	-1.6×10^{-10}
	(1.5×10^{-9})	(1.4×10^{-9})	(7.6×10^{-10})
GROWTH	.594	.110	.125
	(.045)	(.041)	(.022)
PBLUE	.240	.228	.032
	(.044)	(.041)	(.022)
R^2	.44	.33	.26

Note: Standard Errors in Parentheses
All equations include 17 industry dummies, 5 region dummies, and 7 year dummies.
UNION = 1 if plant covered by collective bargaining argreement
SINGLE = 1 if plant not part of multi-establishment company
SIZE = Total employment in year previous to year of turnover observation
GROWTH = Growth rate of total employment during year of turnover observation
PBLUE = Proportion of Blue-Collar (Craft, Operatives, Labor, Service) workers in previous years