Union Effects on Employment Stability: A Comparison of Panel Versus Cross-Sectional Data*

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In analyzing the impact of unions on employment stability, we introduce two innovations to correct for previous overestimates of the union effect: (1) a different approach to measure employment stability (although we use traditional measures for comparisons) and (2) panel data to obtain estimates uncontaminated by unmeasurable person differences between union and nonunion workers. Using common cross-sectional techniques, the new measures of employment stability indicate that unions account for 15 to 20 percent of employment stability variance rather than the 40 to 60 percent indicated in previous findings. Using panel techniques further reduces these measured union effects to no more than 11 percent.

I. *Introduction*

The problem of how to measure the effect of labor unions on employee well-being has long intrigued labor economists and remains a subject of intense debate (see, for example, Freeman and Medoff, 1981). Until recently, well-being was generally measured in terms of wages, and the literature on this issue focused entirely on the comparison of union and nonunion wage gains (see Lewis, 1963, and 1982). Typically, in the early studies ordinary least squares regression analysis was applied to cross-sectional data. While specific estimates have varied widely, results have generally shown that union members receive a wage advantage ranging between 10 and 20 percent.

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Over the past several years, criticism of such studies has escalated (Lee, 1978; Schmidt, 1978; Schmidt and Strauss, 1976), particularly surrounding the argument that single equation regression methods do not take into account possible innate, and perhaps unmeasurable, differences between union and nonunion workers. To the extent that such differences exist — for example, because more productive workers are more strongly attracted to unions — single equation estimates of union effects are subject to an upward selectivity bias. Although attempts have been made to reduce this selectivity bias with the use of cross-sectional simultaneous equation techniques (see Heckman, 1976; Duncan and Leigh, 1980), the robustness of such simultaneous equation estimates has been questioned on the grounds that such techniques are extremely sensitive to errors in functional specification (Goldberger, 1980).

Proponents of the hedonic pricing model (Borjas, 1979; Duncan and Stafford, 1980) advocate another approach to the measurement of union effects on work force well-being. This approach treats wages as an element in the total package of union-supplied amenities that contribute to worker well-being. Viewed from this perspective, higher union wages may be traded off against relatively smaller nonwage amenities, leaving a smaller net gain in well-being for union members than an exclusive focus on wages would suggest. Because the hedonic pricing model permits a more comprehensive analysis of labor market dynamics than do the traditional wage-oriented models, it has been adopted in numerous studies measuring union effects on a wide range of nonwage amenities.

This study explores the effect of union membership on employment stability, an especially important nonwage amenity. Considerable work has been done in this area, most of it drawing on cross-sectional data and expressing employment stability in terms of either quit or layoff probabilities or job tenure. Most often, researchers have concluded that unions greatly increase employment stability, at least among union members (Becker, 1978; Bloch and Kuskin, 1978; Freeman, 1980a, 1980b; Kahn and Morimune, 1979; Medoff, 1979).

This study questions these results. First, the consideration of employment stability in terms of quit and layoff probabilities or job tenure misses the potential effects of strikes and temporary layoffs, and consequently, may create an upward bias on estimates of the union effect on stability. Second, the use of cross-sectional analysis may introduce upward selectivity bias if, for example, more "stable" workers are attracted to unions in the first place.

The results of this study suggest that past estimates of union effects on stability suffer from both problems. Considerable upward bias is detected in the use of tenure as a measure of employment stability. Also important, from the standpoint of method, are the differences that occur between cross-section and panel data estimates: Cross-section estimates persistently indicate a union effect while panel estimates show the effect to be small. This lends tentative support to the observation that part of the relatively strong effect of unions on employment stability found in cross-sectional studies may be attributable to the self-selection of more "stable" workers into unions.

II. *The Theory of Unions and Employment Stability*

The analytic framework most commonly used to study the effect of union membership on employment stability is drawn from the monopoly theory of unionization, primarily because of the wage effects that result from unionization. Unionization restricts the free flow of workers to the firm, reducing downward competitive pressure on wages and raising worker incentives to remain with the firm. In addition, higher wage rates force the firm to select higher quality workers in order to maintain optimal output levels in the product market, where "quality" may manifest itself either as improved productivity (output per hour) or as reduced turnover, which reduces the firm's long-run on-the-job training costs.

A second analytic framework that is consistent with monopoly theory draws on "exit-voice" theory. According to this approach, which has been applied to unions by Freeman (1980a) and others (Blau and Kahn, 1980), employees dissatisfied with some aspect of their work have two options: voicing dissatisfaction verbally or showing dissatisfaction by leaving the firm. If no formal voicing mechanisms exist, dissatisfied workers are more likely to resort to exit behavior, thus raising turnover rates and reducing employment stability. Conversely, if formal voicing mechanisms do exist, such as labor unions, exit behavior should decline and employment stability should improve. Freeman points out that almost all union contracts contain explicit provisions for grievance and arbitration systems (U.S. Department of Labor, 1964; 1977). Many contracts also establish complex rules of procedure that management must follow in order to dismiss unsatisfactory workers, thereby raising severance costs and further stabilizing employment.

While both the monopoly and exit-voice theories imply greater employment stability among union than nonunion workers, counter-arguments can be raised. For example, high union wage requirements may encourage increased capital substitution or subcontracting to lower-cost producers. High wage requirements are also apt to raise the firm's sensitivity to fluctuations in the business cycle, so that temporary layoffs may increase dramatically in periods of recession.

Because there is no theoretical basis for ascertaining *a priori* which set of effects will predominate in actual practice — that is, whether union membership ultimately promotes or inhibits employment stability — the question must be pursued empirically. While past research strongly suggests a net positive effect, there is reason to suspect on both definitional and statistical grounds that this effect may have been substantially overstated.

III. *The Data*

The data set chosen for this study is the Panel Study of Income Dynamics (PSID), a cross-sectional *and* panel survey of about 5,000 "family units" over the period 1968 to 1976. The survey contains a wealth of economic and demographic information, including individual histories of union membership.

The PSID does contain some observations not pertinent to this study, raising the possibility of distortion by irrelevant exogenous factors. One such observation is discrimination. In order to eliminate any possible race and sex discrimination effects, nonwhite and female household heads are deleted from the sample. To minimize any effects of age, health, and disability discrimination, observations containing retired persons, handicapped and disabled persons, and persons who were student household heads when the survey was initiated are also deleted.

The effect of these deletions is to leave a sample of white males, employed or unemployed, but still in the labor force. (Persons not in the labor force in the first year of the survey are also deleted.) The remaining sample is divided into three groups:

1. *ALWAYS* — those persons with continuous union membership during the nine years of data collection;[1]
2. *NEVER* — those persons with no union membership during the nine years of data collection; and
3. *SOME* — those persons with discontinuous union membership during the nine years of data collection.

Two variables are used to measure employment stability: *TENURE* and *VAWH* (variation in annual work hours). *TENURE* measures years on current job (as of 1976) and is used to link this study to past research. *VAWH,* which is defined specifically to isolate the effect of union membership on variability in working hours, measures the standard deviation of annual hours worked both for those years that the respondent was in the union and for those years he was out of the union. Note that *VAWH* is sensitive to strikes and temporary layoffs while *TENURE* is not, since strikes and temporary layoffs tend to affect hours but not tenure. (An exception is where a worker in a higher cyclical industry quits to find more stable employment elsewhere.)

The data in Table 1 indicate substantial differences among the three union membership groups (*ALWAYS, NEVER,* and *SOME*) with respect to both stability measures (*TENURE* and *VAWH*). Respondents in the *ALWAYS* group have

[1]For ease of exposition, we refer to these people as being always in the union, although no information is available on their unionization prior to the sample. Thus, *ALWAYS* refers strictly to the nine years of sampling. This also applies to the *NEVER* group. One other caveat is that no union membership data are available for 1973. It is conceivable, therefore, that individuals within the *ALWAYS* and *NEVER* states changed status in 1973 but were undetected in the data. The probability of this is small, however, and at most very few individuals would be affected.

Table 1

Means for Employment Stability Variables by Strata

Variable	*ALWAYS*	*NEVER*	*SOME*	AGGREGATE DATA SET
TENURE	16.94	11.80	8.99	11.79
Standard Deviation of Annual Work Hours (*VAWH*)	305.04	360.90	417.04	369.19
Education	10.93	13.14	11.60	12.27
Age	39.00	37.71	36.92	37.68
Always	——	——	——	.172
Never	——	——	——	.509
N	175	518	324	1017

Source: PSID Data

Note: All variables are defined in the text. Precise definitions are as follows: *ALWAYS* refers to those individuals in a union in each year for which such information is available of the PSID tape (1968, 1969, 1970, 1971, 1972, 1974, 1975, and 1976). *NEVER* refers to those individuals who were not in a union in each of these years. *SOME* refers to all else; namely, those who were in a union in at least one but not all of the above years. *TENURE* refers to years on current job in 1976. *VAWH* is the standard deviation in annual hours worked over the period 1968-76. Age and education refer to values in 1968.

the highest stability ratings. Their *TENURE* stands at 16.9 years, compared with 11.8 years for the *NEVER* group and 9.0 years for the *SOME* group. Their *VAWH* stands at 305 hours, compared with 361 for the *NEVER* group and 417 for the *SOME* group. In other words, the tenure for the *ALWAYS* group is longer and the working hour variability is smaller than that of those with no union membership and those whose membership was discontinuous. Whether these union effects are attributable to differences in measurable worker characteristics is unclear.

IV. *Cross-Sectional Analysis*

In order to apply cross-sectional analysis to the study sample, the data for each of the three membership groups are pooled. Dummy variables *ALWAYS* and *NEVER* are defined to identify membership status for any individual observation, where the respondent's membership history over the nine years of the survey constitutes a single observation. Within the context of an ordinary least squares regression, the coefficients of these dummy variables can be taken as estimates of union effects on stability, with age and education included to adjust for individual characteristics. The regression takes the form:

$$\begin{bmatrix} TENURE \\ VAWH \end{bmatrix} = \alpha_0 + \alpha_1\, ALWAYS + \alpha_2\, NEVER + \alpha_3\, AGE + \alpha_4\, EDUC + \epsilon, \quad (1)$$

with coefficient estimates reported in columns (1) and (3) of Table 2.

Table 2

Cross-Sectional OLS Regressions of Employment Stability

	Dependent Variable			
Independent Variables	*TENURE*	*TENURE*	*VAWH*	*VAWH*
Intercept	−3.658 (−1.97)	1.928 (1.03)	560.841 (10.44)	530.333 (10.28)
ALWAYS	7.449 (9.17)	6.380 (8.11)	−107.332 (−4.56)	−45.211 (−1.93)
NEVER	2.248 (3.60)	1.765 (2.94)	−48.535 (−2.69)	−29.785 (−1.71)
Age	0.296 (10.30)	0.264 (9.49)	−3.283 (−3.94)	−.811 (−.98)
Education	0.241 (3.02)	0.205 (2.68)	−3.614 (−1.56)	−1.603 (−.72)
VAWH		−.010 (−9.56)		
TENURE				−8.340 (−9.56)
R^2	.18	.25	.04	.12
F Statistic	31.35	41.30	6.12	17.25

Note: PSID data for white male household heads in the labor force. Variables are as defined in Table 1. Regressions are also adjusted for extra training (in addition to schooling) and marital status. The number of observations equals 1,017. Student *t*-values are in parentheses. Both regressions are significant at better than the 0.01 level.

The estimates of α_1 measure differences in average employment stability between those individuals with continuous and those with discontinuous union membership. The estimates of α_2 measure stability differences between those individuals with discontinuous membership and those who held no membership at all. The arithmetic difference between the two estimates $(\alpha_1 - \alpha_2)$ measures the difference in stability between those always and those never in a union (See lines 1 and 3 of Table 3 for transformations of the union effect estimates into percentage terms.)

The cross-sectional regression estimates suggest strong union effects on employment stability. Estimates from the *TENURE* regression indicate that respondents in the *ALWAYS* group maintain 45 percent longer tenure (5.2 years) than respondents in the *NEVER* group and 63.2 percent longer tenure (7.4 years) than those in the *SOME* group. These findings are consistent with results from the *VAWH* regression, which indicate that the standard deviation of annual hours worked was 107.3 hours lower for the *ALWAYS* group than for the *SOME* group and 57.9 hours lower than for the *NEVER* group. These constitute 29.1 and 15.9

Table 3

Union/Nonunion Percentage Differentials in Employment Stability

Union Effect	Comparison Groups		
	ALWAYS/NEVER	*ALWAYS/SOME*	*NEVER/SOME*
Increased Tenure (*TENURE*)			
Estimated with Eq(1)	45.0	63.2	19.1
Estimated with Eq(2)	39.2	54.5	15.0
Decreased Variance in Annual Hours Worked (VAWH)			
Estimated with Eq(1)	15.9	29.1	13.4
Estimated with Eq(2)	4.2	12.2	8.1

Note: Computed from Tables 1 and 2.

percent differences, compared to the 63.2 and 45 percent measures obtained within the "tenure" regressions.

Even these smaller differences in the *VAWH* regressions are overestimates of the union effect on hours variations. If higher union wages reduce turnover among union workers and turnover in turn reduces hours variability, then union members may have less hours variation solely because of diminished turnover. In short, hours variability may be higher among nonunion workers because job separations are not held constant. To avoid this bias, equation (1) is respecified to hold job duration constant:

$$\begin{bmatrix} TENURE \\ VAWH \end{bmatrix} = \beta_0 + \beta_1 ALWAYS + \beta_2 NEVER + \beta_3 AGE + \beta_4 EDUCATION + \beta_5 \begin{bmatrix} VAWH \\ TENURE \end{bmatrix} + \epsilon. \tag{2}$$

The inclusion of job tenure and hours variation serves to ascertain the union effect independent of known union/nonunion stability differences. As expected, union effects (though still statistically significant) have been reduced further to between 4 and 12 percent (column 4 of Table 2 and line 4 of Table 3). As a comparison, the regressions with tenure as a dependent variable are adjusted by *VAWH* (column 2 of Table 3 and line 3 of Table 3) and yield similar effects.

Of particular interest is the comparison between union effects on tenure and on hours-of-work variation, for it is in this comparison that much of the effect of such factors as strikes, temporary layoffs, and overtime can be isolated. Because each of these factors tends to increase significantly hours-of-work variation while having a minimal impact on tenure (in all but the most sensitive cyclical industries), union workers should report longer tenure and lower hours-of-work variation than nonunion workers. Indeed, this is the case. As already indicated, tenure was 45 percent higher for *ALWAYS* respondents than for *NEVER* respondents, while hours-of-work variability was 15.9 percent lower. These figures are 15.9 and 4.2 percent for the so-called adjusted regressions given in equation (2).

The results of cross-sectional analysis, then, support the hypothesis that unions increase employment stability. The results also suggest that the use of tenure as an operational measure of employment stability leads to an overestimate of the union effect because such factors as strikes, temporary layoffs, and overtime are not taken into account.

V. *Selectivity*

There is yet another problem in the interpretation of these cross-sectional estimates. On the one hand, employment stability is greater for the *ALWAYS* group than for the *SOME* and *NEVER* groups, a finding consistent with positive union effects on stability. On the other hand, as Tables 2 and 3 indicate, stability is greater for the *NEVER* group than for the *SOME* group. If unions do increase stability, how does one account for this apparent inconsistency?

The problem may be due in part to selectivity. If respondents in the different membership groups are in some way innately different from one another, then part of the effect on stability attributed to unions in the cross-sectional analysis may be attributable to these innate personal differences. It seems reasonable, for example, in light of the apparent inconsistency noted in the cross-sectional estimates, to suspect that those in the *ALWAYS* and *NEVER* groups may share certain innate characteristics more compatible with stability than those possessed by respondents in the *SOME* group. If the effect of such innate characteristics on stability were in fact more powerful than the effect of union membership, then the apparent inconsistency in the cross-sectional estimates would be resolved.

For reasons of selection, respondents in the *NEVER* group might innately be more prone to stability than respondents in the *SOME* group; and if this innate effect were greater than the union effect, then stability would necessarily be greater for the *NEVER* than for the *SOME* group.

To pursue this possibility, it is necessary to demonstrate that union membership does not account for as much of the employment stability as the cross-sectional estimates suggest. This in turn requires that the effect of unions be isolated from other possible factors affecting stability, so that the "pure" union effect may be compared with the cross-sectional estimates of union effect. These steps are taken with the use of panel data.

VI. *Analysis of Panel Data*

Panel data provide observations on an array of characteristics for each individual in the sample collected at regular intervals over a period of years. The chief advantage of such data over cross-sectional data is that it permits the analyst to control for personal characteristics that might otherwise confound the analysis.

The panel data used in this study (and described in section III) include observations on all three membership groups. However, for the purpose of isolating the "pure" union effect on employment stability, it is the *SOME* group that is of immediate interest. Controlling for all other potentially relevant individual characteristics, one is able to compare employment stability before and after joining or leaving the union, recognizing that whatever differences are detected *must be attributed entirely to union membership.*[2]

The PSID data set contains nine time periods. Analysis of hours-of-work variation[3] requires that respondents in the sample have been both in and out of the union for at least several periods. Consequently, the study sample is restricted to those respondents in the *SOME* group who have switched only once (either into or out of the union) and who have been both in and out of the union for at least three of the nine years.[4] These criteria yield a total of six patterns, three for joiners and three for leavers (see Table 4).

The mean *VAWH* (standard deviation of annual hours worked) is computed for each of the six patterns, for all joiners and all leavers and for the periods before and after joining or leaving the union. Comparison of the means yields mixed results. Union membership improved stability for those with patterns a, d, and f (see Table 4) and reduced stability for those with patterns b, c, and e, so that a clearly positive union effect does not emerge. Analysis of individual observations sheds no further light on these results.

In such a direct comparison of means, however, adjustments are not made for factors other than union membership that may affect stability. As a next step, therefore, regression analysis is applied, so that possible confounding effects may be screened out. With regression, account can be taken of both potential sources of distortion: individual characteristics, both measurable, and economy-wide trends.

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[2]Some may argue that individuals change when they join a union (much like professors change when they get tenure). But we believe that such changes are union induced and, thus, should be considered as a union effect. Our approach incorporates these changes into the union impact.

[3]For the panel analysis, we concentrate on hours variations because analysis of tenure before and after a status switch would have little meaning. Often, union status changes entail a job change. Any recent job change implies low tenure, even if intentions of staying long on one's new job are high. Using variations in weeks worked would also be problematic. Definitional problems exist in terms of how to distinguish voluntary vacations from forced furloughs. Besides, our measure of annual hours encompasses *both* weeks worked per year and hours worked per week. For this reason we concentrate on *VAWH* for the panel analysis.

[4]The small sample size (23 union joiners and 28 union leavers) is not as problematic as some might suspect. First, two observations exist per observation, so that almost 100 degrees of freedom exist in the regressions performed in Table 5. Second, the robustness of the results is tested by comparing the union effects separately for joiners and leavers.

Table 4
Variation in Annual Work Hours for Union Status Switchers by Unionization Pattern

	Joiners			
	Pattern a	Pattern b	Pattern c	Overall Total
	000|11X111	0000|1X111	00000X|111	
VAWH (nonunion)	475.43	233.95	359.49	350.67
VAWH (union)	442.84	267.28	437.29	372.67
N	8	9	6	23
	Leavers			
	Pattern d	Pattern e	Pattern f	Overall Total
	111|00X000	1111|0X000	11111|X000	
VAWH (union)	243.98	479.24	173.94	288.47
VAWH (nonunion)	510.33	268.25	311.11	377.13
N	11	8	9	28

Note: PSID data 1968-76. *VAWH* is the standard deviation of annual hours worked for the years worked as indicated by the unionization pattern (0 ≡ nonunion; 1 ≡ union; and $X \equiv 1973$, for which union data are not available).

In specifying the regression equation, it is again noted that for each individual two sets of hours-of-work variations (*VAWH*) were computed: one for the years before the individual joined or left the union and one for the years after. Along with the measured individual characteristics, then, two critical pieces of information must be taken into account in the regression: first, whether the individual was initially in the union and then left or was initially not in the union and then joined; second, the hours-of-work variation both before and after the switch. The first, included in the regression equation as dummy variable TD_t, makes it possible to relate measures of hours-of-work variation to membership status. The second, denoted $VAWH_{jt}$, makes it possible to account for unmeasured individual characteristics that might otherwise distort estimates of the union effect.

With TD_t, $VAWH_{jt}$, and the necessary personal characteristic variables incorporated into the regression equation, the coefficient of the union dummy variable VD_t can be estimated to measure the pure union effect on hours-of-work variability. The exact specification is

$$VAWH_{it} = a_0 + a_1A_t + a_2E_t + a_3UD_t + a_4TD_t + a_5VAWH_{jt} + \epsilon, \quad t = 1, 2, \ldots, T \quad i \neq j = 1, 2 \tag{3}$$

where

$VAWH_{it}$ ≡ work hours variation for the t^{th} worker in union status i;

A_t ≡ age of the t^{th} worker;

E_t ≡ education of the t^{th} worker;

UD_t ≡ union dummy variable (0 = nonunion and 1 = union) for the t^{th} worker;

TD_t ≡ time dummy variable ($TD_t \equiv 0$ when $VAWH_{it}$ refers to the initial work segment; $TD_t \equiv 1$ when $VAWH_{it}$ refers to the final work segment) for the t^{th} worker;

$VAWH_{jt}$ ≡ hours variation in alternative work segment (compared to $VAWH_{it}$) of the t^{th} worker;

T ≡ number of individuals; and

i, j ≡ time period of the work segment and a normal and independently distributed error term with a zero mean and constant variance for the sample of switchers.

The regression results are presented in Table 5. The first column of results, headed "joiners and leavers," reports estimates for all respondents in the restricted subsample; that is, those in the *SOME* group who either joined or left the union during the nine-year survey period, with at least three years both in and out of the survey. The second and third columns report disaggregated estimates for "joiners" and "leavers," respectively.

Table 5

Panel Estimates of the Impact of Unions on Employment Stability

	(1) Joiners and Leavers	(2) Joiners	(3) Leavers
Constant	205.65(1.2)	264.12(1.2)	89.97(0.3)
Age (A)	0.54(0.2)	0.17(0.1)	2.19(0.5)
Education (E)	−2.24(−.3)	−8.02(−.7)	11.00(0.7)
Union Dummy (UD)	−46.01(−1.0)	32.33(0.5)	−113.11(−1.7)
Time Period Dummy (TD)	76.45(1.6)		
Hours Variation in Alternative State ($VAWH_j$)	0.38(4.1)	0.47(3.4)	0.28(2.1)
R^2	.17	.25	.12
N (observations)	102	46	56
N (individuals)	51	23	28

Note: PSID data. Variables defined in text and previous tables. T-values in parentheses. Note there are two observations per individual, one before and one after union status change.

In the first column, only the coefficient estimate for $VAWH_j$, which represents unobserved individual characteristics, is statistically significant at the 95 percent level, followed by the estimates on the time period dummy, TD_t, and the constant. The low statistical significance of the remaining estimates in the regression implies an overall lack of understanding of employment variability in the market. Most of the explained variance is ascribed to unobservable individual characteristics and to overall market phenomena as embodied in the time trend dummy and the constant. Measured individual characteristics, while signed in keeping with theoretical expectations, have little explanatory power (although it is recognized that not all pertinent characteristics in this category, such as specific training, were included in the model) and neither, it appears, does the union effect. What clearly emerges is that *the positive effect of union membership on employment stability, so readily apparent in cross-sectional analysis, is of small importance when assessed with panel data.* Union effects that are blatant in Tables 2 and 3 are barely discernible in Table 5.

Even if the problem of statistical insignificance is ignored and the observed estimate of the union effect (-46.01) is taken at face value as the most likely point estimate, it still indicates a union effect of *at most* 11 percent (dividing the 46.01 hour reduction by the 417.04 sample mean), in contrast to the on average larger percentage effect estimated with cross-sectional analysis. The importance of this result is apparent: Cross-sectional analysis overestimates the union effect on employment stability.

In order to test for consistency between joiners and leavers, the regression was applied to both groups separately. Separation into two groups, however, leads to perfect correlation between *TD* and *UD*, so that the time trend cannot be isolated and estimates of the *UD* coefficient will consistently be biased. The problem can be circumvented by subtracting (adding) the time trend estimate reported in the first column of Table 5 from (to) the a_3 coefficients estimate on *UD* for joiners (leavers), so that

$$a_{3j} \equiv \text{union effect for joiners} = 32.33 - 76.45 = -44.12$$

$$a_{3l} \equiv \text{union effect for leavers} = -113.11 + 76.45 = -36.66.$$

This arithmetic manipulation reveals a union effect of almost identical magnitude for the two groups (8.8 percent for leavers and 10.6 percent for joiners).

That nearly equal union effects are obtained for such seemingly diverse groups suggests that the efforts to screen out the effect of individual characteristics have been largely successful and that the study estimates, based on a limited sample of "switchers," may be generalized to the population at large.

VII. *Summary*

In this study, the effect of union membership on employment stability is investigated. We allege that past studies overestimate the union effect for two reasons. First, traditional quit, layoff, and tenure measures fail to take into account the variations in hours of work caused by strikes, temporary layoffs, and overtime. Second, the cross-sectional methods adopted in these studies are not sensitive to possible selectivity bias.

In response to these problems, we introduce two innovations. First, employment stability is measured by the standard deviation of annual hours worked rather than by traditional tenure, quit, and layoff indices. Second, panel data are used, making possible uncontaminated comparisons of employment stability before and after changes in union membership status.

The results are strong. Differences between union and nonunion hours-of-work variation are smaller than the differences between union and nonunion tenure, suggesting that use of tenure data to measure employment stability leads to overestimates of the union effect on stability. In fact, when stability is measured in terms of hours-of-work variation rather than tenure, cross-sectional estimates of the union effect decline from the 40 to 63 percent range to the 4 to 30 percent range. Panel analysis, free from selectivity bias, further reduces these union effect estimates to not more than 11 percent. Thus, although unions may have some affect on employment stability, past estimates appear to suffer from substantial upward bias. The introduction of a new measure of employment stability and the use of panel data have reduced estimates of the union effect by approximately five-fold.

APPENDIX

Individual Data on Annual Work Hours Variations by Unionization Pattern

Union Joiners				Union Leavers			
Pattern	Observation Number	Out of Union	In Union	Pattern	Observation Number	In Union	Out of Union
	1.	709.66	470.58		1.	102.74	280.71
	2.	1151.80	493.58		2.	47.00	549.69
	3.	315.44	420.40		3.	86.75	257.71
000|11X111	4.	196.17	144.65	111|00X000	4.	52.73	295.28
	5.	237.08	848.79		5.	140.00	240.19
	6.	311.26	149.71		6.	502.22	1137.38
	7.	826.94	860.98		7.	602.94	996.03
	8.	55.11	154.05		8.	31.07	285.69
					9.	737.42	847.49
					10.	155.47	422.16
					11.	219.43	291.25
	1.	445.47	438.73		1.	689.83	186.73
	2.	69.59	339.39		2.	242.60	145.96
	3.	90.80	369.57		3.	938.11	101.38
0000|1X1111	4.	150.79	183.14	1111|0X0000	4.	41.83	347.45
	5.	22.44	23.89		5.	426.18	311.73
	6.	306.50	257.27		6.	433.30	278.34
	7.	40.28	44.40		7.	822.71	381.94
	8.	747.27	505.21		8.	239.38	392.46
	9.	232.31	243.44				
	1.	320.08	826.80		1.	253.47	402.95
	2.	456.47	543.51		2.	285.17	361.93
	3.	344.71	359.01		3.	92.58	289.04
00000X|111	4.	149.06	407.52	11111|X000	4.	81.29	275.69
	5.	577.77	286.02		5.	424.75	449.76
	6.	308.83	200.86		6.	158.52	432.79
					7.	36.49	340.66
					8.	134.45	91.94
					9.	98.79	155.24
		14 Better	9 Better			21 Better	7 Better

Note: Variables and sample are defined in Table 4.

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