PANEL ESTIMATES OF UNION EFFECTS ON WAGES AND WAGE GROWTH

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Abstract—Standard approaches for measuring the impact of unions on the age–earnings profile are unable to account for innumerable quality differences that exist between union and nonunion workers. Furthermore, the usual adjustments used to handle selectivity and heterogeneity are not robust. This paper circumvents these problems by using panel data of one-time union switchers. We find *no* evidence to support the assertion that unions flatten the age–earnings profile, and are able to reconcile these new findings with the opposite results reached by past cross-sectional analysis.

I. Background

The question of union effects on wages has evoked much attention. In the past, most analyses have been cross-sectional, with two results emerging with uniform regularity: first, that unions appear to raise wages, and second that unions tend to flatten age–earnings profiles.

Despite strong consistency across numerous studies these results have been questioned because of two innate selectivity-type biases inherent in cross-sectional estimates. One is akin to extrapolation errors in which observations with missing data, such as for those not in the labor force, are omitted from the analysis, yielding estimates of union effects that need not be global since they pertain to a potentially atypical sample. The second is a heterogeneity type bias caused by possible unmeasured differences between union and non-union members.

These selectivity bias problems have been treated within a cross-sectional simultaneous-equations context.[1] One equation dictates the probability of an observation being in the sample, while the other equation includes an inverse Mills ratio reflecting the previously omitted probability. Goldberger (1980) and Lewis (1983) have questioned the robustness of this technique.

An alternative that has recently gained wide attention is the use of panel data. Panel data enable one to run fixed-effect models with individual-specific intercepts comprising one dummy variable for each individual, so that in effect account is taken of unobserved person-specific heterogeneity. Such models (e.g., Brown (1980), Mellow (1981), Mincer (1981) and Freeman (1984)) are comparable to running regressions with variables expressed in deviation form.[2] Freeman (1984) criticizes this methodology on the grounds that errors of measurement, especially with regard to changes in union status, severely bias downward estimates of union effects. Chamberlain ((1980), p. 225) claims that the first difference technique "provides a consistent estimator . . . *provided that* there is sufficient variation in" ΔU. Thus, if too small a portion of the sample changes status, then inconsistent estimates of the union effects emerge.

II. The Approach

This paper develops a simple alternative to previous approaches by using a pooled cross-section time-series error-variance-component technique applied to panel data. Heterogeneity biases are avoided by comparing entire earnings profiles for given workers before and after their union switch. Potential measurement error biases inherent in first differencing are also avoided.

To be specific, in past analyses measurement errors usually take two forms. One constitutes reporting errors in union status, the other reporting errors in wages. With our approach a sufficient number of years is used so as to cross-check the data. More precisely, by dealing with one-time union status switchers, any change in union status is cross-checked by union status in each of the other years to insure against status being incorrectly reported. Further, any errors in wages are "smoothed" since they become only one data point in a wage regression.

Union joiners are most likely the able young workers, while union leavers are probably the less able involuntary leavers so that each group is potentially at opposite ends of the quality spectrum. By looking at the effects of unions separately for joiners and leavers one can test the similarity of estimated union effects. Given the diversity of the groups, similar parameters would imply that these estimates may be global. Nonetheless given that switchers represent a select sample, an inverse Mills ratio is also used to adjust for possible selectivity type biases. In addition, we reconcile our new panel results with past cross-sectional analyses.

III. The Data and Statistical Specification

We use the Panel Study of Income Dynamics (PSID) for this study. The PSID contains data on wages and union membership for each year from 1968 to 1981. Of

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[1] See Heckman (1976) who deals with globality, and Willis and Rosen (1978) who deal with unobserved heterogeneity.

[2] For data with two time periods this amounts to considering each variable as a first-difference.

TABLE 1.—MGLS ESTIMATES OF UNION STATUS (COLUMNS 1–4), EMPLOYMENT CHANGE (COLUMNS 6–9) ON WAGES AND PSID 1981 CROSS-SECTIONAL OLS WAGE ESTIMATES (COLUMN 5)

	Union Status Change[a]						Employment Change[b]			
	Joiners		Leavers		X-section		Joiners	Leavers	Nonunion	Union
	(1)	(2)	(3)	(4)	(5)		(6)	(7)	(8)	(9)
Intercept	4.67	3.86	5.06	5.64	3.72		4.70	5.30	4.10	5.40
	(35.0)	(4.36)	(29.1)	(2.0)	(20.0)		(33.4)	(36.3)	(79.0)	(76.5)
s	0.065	0.105	0.045	0.022	0.075[g]		0.065	0.031	0.10	0.031
	(9.1)	(2.5)	(5.3)	(0.2)	(11.8)		(9.0)	(3.7)	(39.0)	(8.3)
j	0.021	0.020	0.016	0.017			0.021	0.020	0.04	0.008
	(3.1)	(2.8)	(3.5)	(3.4)			(3.0)	(3.9)	(22.3)	(2.6)
j^2	−0.0004	−0.0003	−0.0003	−0.0004			−0.0004	−0.0004	−0.0009	−0.00007
	(2.7)	(1.7)	(3.5)	(1.6)			(2.7)	(3.6)	(20.6)	(1.0)
e	0.034	0.035	0.0082	0.0093	0.074		0.056	0.017	0.031	0.037
	(1.0)	(1.1)	(0.3)	(0.3)	(11.5)		(2.3)	(0.9)	(3.6)	(3.5)
e^2	−0.00005	−0.00009	−0.0011	−0.0012	−0.001		−0.0026	−0.0025	−0.0018	−0.002
	(0.02)	(0.03)	(0.7)	(0.7)	(12.8)		(1.86)	(1.9)	(3.3)	(2.9)
U	−0.159	−0.423	0.0025	0.177	1.22	E	−0.214	−0.15	−0.11	−0.07
	(1.0)	(1.3)	(0.01)	(0.2)	(2.96)		(1.7)	(1.4)	(2.16)	(1.0)
$U*e$	0.072	0.071	0.045	0.045	−0.038	$E*e$	0.06	−0.01	−0.003	0.02
	(1.5)	(1.5)	(1.2)	(1.2)	(2.10)		(1.3)	(0.35)	(0.15)	(1.1)
$U*e^2$	−0.005	−0.006	−0.0014	−0.0014	0.00054	$E*e^2$	−0.004	0.0024	0.0012	−0.002
	(1.7)	(1.7)	(0.5)	(0.3)	(2.13)		(1.2)	(0.9)	(0.8)	(0.9)
Inverse Mills Ratio		−1.983		1.845						
		(0.94)		(0.2)						
DF	411	410	873	872	944		411	803	6725	1461
R^2	0.273	0.275	0.135	0.138	0.196		0.273	0.095	0.241	0.115
F-Ratio	19.33	17.25	13.20	15.40	33.0		19.35	10.6	267.8	23.7
Union Effect[c]	0.001	−0.213	0.217	0.38	0.20	Job[h] Change Effect	0.00024	−0.161	−0.021	−0.087
Net Union Effect[d]	0.091	−0.213	0.091	0.219						
Net Union Effect[e]	0.178	−0.126	0.196	0.359						
Slope Effect[f]	0.013	0.013	0.029	0.029	−0.004					

Note: t-values are in parentheses.
[a] $U = 1$ if union.
[b] $E = 1$ if employment change.
[c] See footnote 6.
[d] Union effect adjusted by employment change effect computed from Joiner and Leaver equations (columns 6 and 7).
[e] Union effect adjusted by employment change effect from Nonunion and Union equations (columns 8 and 9).
[f] See footnotes 10 and 11.
[g] $(U*S)$ coefficient $= -0.034$ ($t = 2.10$).
[h] See footnote 8.

the sample of 946 white male heads of household who could be followed for *each* year and who had consistent data,[3] 113 were always in a union and 518 were never in a union. The remainder of the sample made at least one switch. The analysis requires that we examine only one-time switchers since we need sufficient time-series data to estimate the age–earnings profile. There are 63 individuals who leave a union only once between 1968 and 1981; and 30 who join only once between these two dates.

Pooling the data over both cross-section and time-series requires an estimation that accounts for the potential correlation among the disturbances. Standard error-components models attempt to account for such correlations by breaking the error term into an individual component (v_i) which represents the effect of omitted individual exogenous variables; a time component (p_t) since disturbances may be unique to specific time periods while affecting individuals more or less equally, and a component accounting for the possibility

[3] Data consistency implies positive earnings, and no missing data for union membership, education, and age. Union membership for 1973 was extrapolated since it was not reported in the PSID.

that disturbances may be peculiar to an individual at a specific point of time (z_{it}).

Denoting the stochastic error term to be E_{it} ($= v_i + p_t + z_{it}$), we employ a wage function of the form:[4]

$$\ln w_{it} = a_0 + a_1 S_{it} + a_2 j_{it} + a_3 j_{it}^2 + a_4 e_{it} + a_5 e_{it}^2 + a_6 U_{it} + a_7 (U_{it}^* e_{it}) + a_8 (U_{it}^* e_{it}^2) + E_{it}$$

$$i = 1, \ldots, N \qquad t = 1, \ldots, T \tag{1}$$

where $\ln w_{it}$ is the natural logarithm of real hourly wages (in 1968 dollars) of the i^{th} individual in the t^{th} year, S equals years of schooling, j equals experience prior to 1968, e equals experience since 1968, and where the squared terms denote typical nonlinearities inherent in earnings functions. The variants v_i, e_t, and z_{it} are assumed to be independent of each other as well as independent of the x_{itj}'s, and identically distributed with means of zero and variances σ_v^2, σ_p^2, and σ_z^2, respectively.

The coefficients of the union as well as the post-union 1968 experience interaction terms give a measure of union impact. The union effect is composed of (1) the direct or immediate impact on wages

$$\left[\frac{\partial \ln w}{\partial U} = a_6 + a_7 e + a_8 e^2\right]$$

and (2) the impact on the earnings-profile slope

$$\left[\frac{\partial^2 \ln w}{\partial U \partial e} = a_7 + 2a_8 e\right].$$

If unions flatten profiles, one should observe a deceleration of earnings growth for union joiners implying that $(a_7 + 2a_8 e)$ should be negative. Similarly for leavers, if unions flatten earnings profiles, one should observe an acceleration of earnings growth upon leaving the union, implying that $(a_7 + 2a_8 e)$ again be negative. Modified generalized least square estimates of equation (1) for joiners and leavers are presented in table 1.[5]

IV. The Union Effect on Wages

The direct wage effect is obtained by evaluating $\partial \ln w / \partial U$ at the year when the union status switch occurs. Joiners appear to gain 9.1% immediately upon joining, while leavers face a 21.7% loss immediately upon leaving.[6] Because most union status changes are accompanied by a change in employment these differences between the switcher groups are exaggerated. Theories of specific training and models of efficient-wage contracts imply that employment turnover usually results in an initial loss in earnings power, with involuntary turnover accentuating the loss. By accounting for job turnover effects, one can reduce the degree of measured differences in the status coefficients.[7]

Columns 6, 7, 8, and 9 (table 1) contain panel earnings regression estimates emphasizing job change for four strata of job changers. The effect of an employment change $[(\partial \ln w/\partial \text{ employment } \Delta)|_{e=\text{switch}}]$ being about zero for joiners, is a 13% loss for leavers, an 8.7% loss in earnings for the nonunion group, and a 2.1% loss for union members,[8] implying that the computed union effect is overestimated for leavers. Recomputation of the union effect after adjusting for job change yields a 9.1% union impact (using both the leaver and joiner employment-change regressions) and a 17.8% to 19.6% union impact (using the union and nonunion employment-change regressions).[9] These estimates fall between current panel and cross-sectional measures of union effects, which is perfectly reasonable given Freeman's (1984) claim that the true union effect lies between traditional panel and cross-sectional estimates.

[4] See Wunnava (1986) for empirical results with an alternative specification.

[5] Since the variance components are unknown, the modified (feasible) GLS estimate is used. The variance components of E are estimated by the "fitting-of-constant" method of Searle (1971).

[6] Mean value of e is 5.86 for joiners and 5.83 for leavers. Thus,

$$\left[\frac{\partial \ln w}{\partial U} = -0.159 + 0.072(5.86) - 0.005(5.86)^2 = 0.091(\text{gain})\right]$$

for joiners, while for leavers

$$\left[0.0025 + 0.045(5.83) - 0.0014(5.83)^2 = 0.217(\text{loss})\right].$$

Note that the U, $U*e$, and $U*e^2$ coefficients are jointly significant. Due to space limitations we are not discussing the results incorporating inverse Mills ratio (columns 2, 4 of table 1).

[7] Similarly, one can adjust for other changes that might accompany changes in union status. However, to the extent that such changes are not subsumed in a job change, their measurement is beyond the scope of this study. Too small a sample currently exists to measure these effects.

[8] These results computed as follows: $\partial \ln w/\partial E = a_6 + a_7(e) + a_8(e^2)$; evaluating at e = switch point yield

Joiners $= -0.214 + 0.06(5.86) - 0.004(5.86)^2 = 0.00024$

Leavers $= -0.15 - 0.01(5.83) + 0.0024(5.83)^2 = -0.127$

Always Union $= -0.07 + 0.02(5.8) - 0.002(5.8)^2 = -0.021$

Always Nonunion $= -0.11 - 0.003(5.8) + 0.0012(5.8)^2 = -0.087$

support the contention that union members accumulate less specific training than nonunion workers, but that union leavers are involuntary job changers.

[9] For more detail on the way in which these are computed with different specifications see Wunnava (1986) and Polachek et al. (forthcoming).

V. The Union Effect on the Age–Earnings Profile Slope

Regressions with and without selectivity corrections indicate that unions do *not* flatten the age–earnings profile. For union joiners there is a 1.3% increase[10] in the life cycle earnings profile from joining a union. We find a 2.9% decrease[11] for union leavers. Since *none* of these estimates is significantly negative, a conclusion that unions do *not* flatten age–earnings profiles is justified. This is consistent with unions organizing where the age–earnings profile is flat (Polachek (1979)), and with cross-sectional age–earnings profiles being flatter in unionized firms (Lazear (1983)), but inconsistent with the inference that unions flatten age–earnings profiles as is often inferred from cross-sectional and wage variance studies (Hyclak (1979) and Freeman (1980)).

VI. A Reconciliation of Cross-Sectional and Time-Series Parameters

Recall that past estimates of union effects computed from cross-sectional research are based on regression coefficients of unions and nonunion workers taken together. Similar regression coefficients for almost comparable strata[12] can be obtained by appropriately manipulating the results reported here. The j (and j^2) as well as e (and e^2) coefficients for joiners and leavers are analogous to past cross-sectional estimates, while the $(U*e)$ and $(U*e^2)$ coefficients are a measure of the time-series effects of unions, holding constant population heterogeneity. Computing slopes for each of the two groups separately before changes in union status yields cross-sectional profile slope estimates which can be used to reconcile our panel results with those of past cross-sectional estimates.

A. *Cross-sectional estimates*

Since joiners are non-union before joining, and leavers are union members before leaving we should find that prior to the status change the slope for joiners exceeds that of leavers. As can be seen joiners' experience coefficients [(j) and (e)] exceed those of leavers (0.021 vs. 0.016 and 0.034 vs. 0.008). Computing the slope at mean experience years in 1968 yields a similar result ($\partial \ln w/\partial j|_{\text{joiners}}$ at 17.94 years of experience) = 0.0066 which is greater than ($\partial \ln w/\partial j|_{\text{leavers}}$ at 20.2 years of experience) = 0.0038. Evaluating[13] $\partial \ln w/\partial e|_{\text{joiners}}$ at the switch point (5.8 years) yields 0.0345 which exceeds $\partial \ln w/\partial e|_{\text{leavers}}$ at 5.8 years (= 0.0242); also consistent with previous cross-sectional results that unions flatten the age–earnings profiles.

These results can be corroborated by using traditional cross-sectional techniques. By treating data in any given year of the panel as cross-sectional data, one can perform the standard ordinary least squares (OLS) regression. We have used 1981 data for this purpose. A specification comparable to our equation (1) (column (5) of table 1) yields a result typical of standard cross-sectional analysis that unions increase wages in the neighborhood of 20%, and also unions flatten the profile.

B. *Cross-sectional versus panel estimates*

If cross-sectional results are true one should find that after a status change the slope for leavers (who are now nonunion) should exceed that of joiners (who are *now* union members). Our panel estimates do not support this phenomenon. Our results[14] show that the slope for joiners after joining (0.048) exceeds the slope for leavers after leaving (−0.0046), that the slope for joiners after joining (0.0485) exceeds their slope in the nonunion segment (0.0345), and that the slope for leavers after leaving a union (−0.0046) is less than their earnings profile slope in the union segment (0.0242). Thus, contrary to the cross-sectional studies unions do not necessarily flatten the age–earnings profile. If anything, unions might even steepen them.

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[10] $$\left.\frac{\partial^2 \ln w}{\partial U \partial e}\right|_{\text{at Switch}} = 0.072 - 2(0.005)(5.86) = 0.013.$$

[11] $$\left.\frac{\partial^2 \ln w}{\partial U \partial e}\right|_{\text{at Switch}} = 0.045 - 2(0.0014)(5.83) = 0.029.$$

[12] The strata are not entirely comparable because we omit those individuals who are either union members or nonunion affiliates over the entire panel.

[13] From equation (1),

$$\frac{\partial \ln w}{\partial e} = a_4 + 2a_5 e + a_7 U + 2a_8 e$$

where U is 1 in the union segment and 0 in the nonunion segment. The following are computed at the switch point: Year = 5.8:

$$\begin{aligned}
\text{Slope for joiners after joining} &= 0.034 + 2(0.00005)5.8 \\
&\quad +0.072 - 2(0.005)5.8 \\
&= 0.0485 \\
\text{Slope for leavers after leaving} &= 0.0082 - 2(0.0011)5.8 \\
&= -0.00456 \\
\text{Slope for joiners before joining} &= 0.034 + 2(0.00005)5.8 \\
&= 0.0345 \\
\text{Slope for leavers before leaving} &= 0.0082 - 2(0.0011)5.8 \\
&\quad +0.045 - 2(0.0014)5.8 \\
&= 0.0242.
\end{aligned}$$

[14] See footnote 13.

VII. Conclusion

Standard cross-sectional approaches to measuring union effects on age–earnings profiles suffer from selectivity. They are not able to adjust for quality differences that exist among union and nonunion workers. The usual adjustments to handle selectivity are not robust. On the other hand, the current panel studies are plagued by measurement errors (Freeman, 1984). This paper circumvents these past problems by using a pooled cross-section time-series error variance component-model applied to panel data (PSID, 1968–81) for a set of one-time union switchers (so that age–earnings profiles of given workers can be compared before and after switching union status). We find that

(1) the union wage effect is about 15%, which falls between past cross-section and panel analyses much as predicted by Freeman (1984),
(2) there is no evidence to support the assertion that unions flatten the age–earnings profile, and
(3) a reconciliation exists between past cross-sectional results and our current panel analysis.

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WAGE ADJUSTMENT IN CONTRACTS CONTAINING COST-OF-LIVING ALLOWANCE CLAUSES

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Abstract—The emerging literature on the determinants of wage change in contracts containing cost-of-living adjustments is reviewed. Models currently in use are shown to belong to the same class. Strong arguments can be adduced in support of an alternative class of models. It is shown that both classes are nested within a more general framework which contains the models of wage adjustment used to analyse COLA and non-COLA contracts in both Canada and the United States. Empirical evidence, drawn from a sample of Canadian wage agreements, favours the new class of models. Canadian and U.S. results are compared.

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I. Introduction

Research on wage adjustment increasingly relies on contract data where the regressand is the annualised rate of change in the base wage rate over the life of each agreement. There appears to exist some agreement on