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Cross-sectional versus panel estimates of union wage effects

Evidence from PSID *

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The main objective of this paper is to compare and reconcile the cross-sectional and the panel estimates of union wage effects. It uses Panel Survey of Income Dynamics (PSID) data for a period of 6 years (1979–1984). Unlike what Freeman (1984) predicted (i.e., panel union wage estimates would be smaller than pure cross-sectional estimates due to the cumulative nature of the errors in reporting union status in the former), both the cross-sectional and panel estimates of union wage effects are very similar ranging from 11 to 13%. Unions seem to have flattening effect on the age–earnings profiles. Not surprisingly, the rate of return to schooling of non-union members is larger than that of the union members.

1. Introduction

Empirical evidence of the impact of unions on wages has grown tremendously in recent years. [1] Most of the previous studies which had addressed the impact of unions on wages have used cross-sectional data. In a life cycle earnings context, a pure cross-sectional analysis assumes that we are following an individual over his lifetime. This means that all individual differences (heterogeneity) can be accounted for by exogenous variables so that each person can be considered identical. [2] Some researchers [3] in recent years have suggested that unobserved characteristics of union and

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[1] See Lewis (1963) cross-sectional, (1986) cross-section/panel, Parsley (1980 cross-section), Freeman and Medoff (1984) cross-sectional/panel, and Hirsch and Addison (1986) cross-sectional/panel, for excellent surveys of recent literature.

[2] If the unmeasured productivity which is specific to the individual, is held constant over time, then we can look at changes over time in an individual's wage and union status to measure the 'union' effect. However, such an unmeasured productivity becomes part of the error term in a cross-sectional framework. If union workers are more productive than nonunion workers, then such an unmeasured productivity would be positively correlated with union status. Accordingly, the cross-sectional OLS estimates of the union effect could be biased *upward*. In such a case estimates will not even be consistent.

[3] See Freeman (1984), Mincer (1981), Mellow (1981), Polachek et al., (1987), and Wunnava (1988).

nonunion workers are very different, and such unmeasured characteristics should be adjusted before looking at the impact of unions on wages. Panel or longitudinal data enables the researcher to avoid unmeasurable sample heterogeneity by concentrating on *changes* in measured variables for given individuals, under the assumption that unmeasured variables remain constant over time. [4] Traditionally, the union–nonunion wage gap based on cross-sectional data [5] has been in the 25 to 30 + % range, while the longitudinal studies [6] yield much lower estimates, i.e., 10 to 20%. This paper provides a comparison and reconciliation between the cross-sectional and the panel estimates of union wage effects.

2. The methodology

Initial studies were aggregate cross-sectional analyses. Mean wages were compared for all union and non-union workers. No adjustments were made for worker attributes. In latter years, to adjust for individual attributes, earnings function type models have been used extensively in the literature. Essentially, OLS wage regressions were run with the following form:

$$\ln \textit{Earnings} = \beta_0 + X\beta + \alpha_0 U + \textit{Error}, \tag{1}$$

where ln = natural log, X = a vector of standard human capital variables, U = an union dummy variable.
However, estimating eq. (1) type models would *only* account for intercept differences. Some researchers are very critical of such an approach. In the words of Bloch and Kuskin (1978, p. 183):

> 'Wage equation is estimated for union and non-union workers with only a dummy variable indicating union membership explicitly constrain the labor market rewards to all other worker characteristics to be equal in the two sectors. Such a procedure not only fails to address the interesting question of differences in wage determination across sectors but also may yield poor estimates of union–non-union wage differentials for workers with otherwise similar characteristics.'

So they proposed fitting *two* separate equations for union an nonunion workers. An econometric alternative to estimating two separate equations for union and non-union members is to modify the single equation method by allowing interaction between U (i.e., union status) and all other independent variables in the model. So eq. (1) can be respecified as (2) by bringing in a vector of interaction terms $(U*X)$:

$$\ln \textit{Earnings} = \beta_0 + X\beta + \alpha_0 U + (U*X) + \textit{Error}, \tag{2}$$

by letting X = years of schooling (S), experience (E), and (E_2) [(E^2) to capture typical non-linearity inherent in earnings functions], unemployment rate (UE), and replacing 'Earnings' by 'W' (real

[4] See Hausman and Taylor (1981), Hsiao (1985) for an excellent discussion on the merits and demerits of using panel data.
[5] Cross-sectional union wage effects will be biased upward if the situation depicted in footnote 2 is in fact true.
[6] Freeman (1984) argues that *lower* union panel wage effects are mostly due to the measurement errors in reporting of union status.

hourly wages), and bringing in dummies for occupation, industry and regions we can obtain the empirical earnings function used in this paper as given in the eq. (3):

$$\ln W = \beta_0 + \beta_1 S + \beta_2 E + \beta_3 E^2 + \beta_4 UE + \alpha_0 U + \alpha_1 (U^*S) + \alpha_2 (U^*E) + \alpha_3 (U^*E^2)$$
$$+ \alpha_4 (U^*UE) + [\text{Vector of Controls for Occupation, Industry, and Regions}]$$
$$+ \mathit{Error}, \tag{3}$$

where

$$E = (Age\text{-}S - 6).$$

Unlike earlier studies we also bring in UE [because UE in Panel Survey of Income Dynamics (henceforth PSID) refers to the local area regional unemployment rate] and interaction between U and UE into the model to capture the influence of the local labor market conditions on the wage determination. One would expect a negative coefficient for UE and a positive coefficient for the interaction term because unions resist wage cuts in response to worsening labor market conditions and also union contracts are mostly multi year contracts. Accordingly, union wages are mostly unresponsive to short run fluctuations in the economy. [7]

To reiterate, eq. (3) will be the basic empirical specification of the wage function used for both the cross-sectional and the panel estimation. The union impact on the *level* of the age–earnings profile is equal to ($\partial \ln W/\partial U$), and on the *slope* of the age–earnings profile is equal to ($\partial^2 \ln W/\partial U\, \partial E$).

3. Data and results

The main purpose of this paper is to compare and reconcile the cross-sectional and the panel estimates of union wage effects. The longest currently available panel data containing information on unionism, wages, and other standard human capital variables is the Panel Survey of Income Dynamics (PSID). Data [8] spanning a complete business cycle from 1979 to 1984 are used for this study. By treating data in any given year of the panel as a cross-section, one can estimate eq. (3) by OLS. This estimation is implemented to estimate the cross-sectional wage regressions. Pooling the data over both cross-section and time-series [to estimate eq. (3)] requires non OLS estimation procedures because of potential correlation among the disturbances. The standard 'error components' [9] model attempts to account for these correlations, raising the asymptotic efficiency of the estimates. Comparison of the estimated mean square errors (MSE) of pooled OLS and error

[7] This argument is also consistent with the 'wage rigidity hypothesis' of Hendricks (1981) and Perry's (1980, 1986) concept of a 'wage norm,' a generally accepted rate of wage increases. According to Perry, such a norm is insensitive to short-term fluctuations. See also Mitchell (1985, 1986) for empirical evidence supporting the presence of wage norm relationship in mostly union sector. Interestingly, some recent research suggests an anomaly to this phenomenon during late seventies i.e., widening union nonunion wage gap even when inflation is relatively higher. This phenomenon can be explained by the widespread introduction of the COLA clauses in the union contracts in view of combating the effects of inflation (Hendricks and Kahn 1983), and also due to the spiralling stagflation (i.e., simultaneous increase in inflation and unemployment).

[8] To avoid possible biases due to racial/sexual discrimination only white male heads of household have been selected for the proposed research.

[9] See Fuller and Battese (1974), and Searle (1971). Pooled OLS and error components results are both displayed in table 2 for comparative purposes.

Table 1
Summary of cross-sectional union wage/slope effects PSID 1979–1984 (absolute t-ratio in parentheses). [a]

Year	$[\partial \ln W/\partial U]$	$[\partial^2 \ln W/\partial U\, \partial E]$
1979	0.125 (8.58)	−0.0053 (0.86)
1980	0.122 (3.10)	−0.0019 (0.38)
1981	0.124 (3.28)	−0.0012 (0.25)
1982	0.118 (2.44)	−0.0013 (0.30)
1983	0.126 (2.63)	−0.0011 (0.25)
1984	0.122 (3.36)	−0.0008 (0.23)

[a] Based on the OLS estimated of eq. (3) full regression results are available on request. Sample size for each year is 460.

Table 2
Pooled regression results: PSID 1979–984 (Dependent variable is in ln real hourly wages, absolute t-ratio in parentheses). [a]

Regressors [a]	OLS		Error components	
Constant	0.3511	(3.93)	0.4146	(3.43)
S	0.0558	(18.5)	0.0489	(12.57)
E	0.0168	(7.74)	0.0217	(8.18)
E^2	−0.0003	(6.87)	−0.0004	(8.10)
UE	−0.0186	(8.84)	−0.0022	(1.32)
U	0.6812	(8.73)	0.6217	(7.58)
$(U*S)$	−0.0348	(7.65)	−0.0375	(6.73)
$(U*E)$	−0.0121	(3.05)	−0.0061	(1.62)
$(U*E^2)$	0.00023	(2.81)	0.00014	(1.81)
$(U*UE)$	−0.0010	(0.30)	0.0009	(0.39)
$[\partial \ln W/\partial U]$	0.130	(5.76)	0.137	(4.28)
$[\partial^2 \ln W/\partial U\, \partial E]$	−0.002	(0.94)	−0.0003	(0.14)
Sample	2760		2760	
adj. R^2	0.423		0.478	
F-ratio	57.17		80.3	
MSE	0.0586		0.0149	
σ_i^2 (estimated cross-section error variance)			0.034	
σ_t^2 (estimated time-series error variance)			0.002	
σ_o^2 (estimated purely random error variance)			0.014	

[a] Other controls include dummies for occupations, industry, and regions. Full regression results can be obtained upon request.

component regressions in table 2 shows that error components model estimates are relatively more efficient. The summary [10] of union effects based on the OLS cross-sectional regressions [of equation (3)] are reported in table 1. Cross-sectional OLS estimates from table 1 indicates that $(\partial \ln W/\partial U)$ yield a result that unions increase wages significantly, and in the range 11 to 13%. Further, a negative estimate of $(\partial^2 \ln W/\partial U\, \partial E)$ reinforces the conclusion [11] of earlier cross-sectional studies that union members have flatter profiles.

[10] To conserve space, these regression results are not reported in their entirety. They are available upon request.

[11] This in the past literature is based on some of the following observations:

(a) Smaller experience coefficients for union members relative to non-union members [Bloch and Kuskin (1978)].

(b) Union effect on the level of wages is smaller for older workers than for younger workers [Mellow (1981), Mincer (1981)].

(c) Smaller dispersion of wages for union members than for non-union members [Freeman (1980), Hyclak (1979), Hirsh (1982)].

For the same cross-sections used in table 1, pooling data from 1979 to 1984 yields very similar results, which challenges the finding of Freeman (1984) that panel estimates yield a lower estimate of union wage effects. [12] These estimates are presented in table 2. Union wage effect [13] for both pooled OLS and the error components model is in the neighborhood of 13%. Further, a negative estimate [14] of (∂^2 in $W/\partial U\ \partial E$) for both pooled regressions is also in accordance with the cross-sectional hypothesis that unions flatten the profiles. [15] Since both the cross-sectional and panel estimates yield similar union wage effects, we could conclude that union wage effects obtained in this study are indeed robust.

One can also notice that the rate of return to schooling for union members is much smaller [16] than nonunion members. This supports the assertion that once workers become union members, they have a lesser incentive to invest in further schooling. [17] This result also explains why occupations with less specific training are most likely unionized (i.e., unions are predominant in mostly blue collar as opposed to white collar occupations).

4. Conclusion

This paper makes a cross comparison between the cross-sectional and panel union wage effects. Data (PSID) for a period of 6 years (1979 to 1984) is used. Cross-sectional union wage level estimates are in the range of 11 to 13% and interestingly the pooled estimates also yield similar results. Both cross-sectional and pooled estimates seem to support the conventional wisdom that unions flatten the age-earnings profiles. The rate of return to schooling for nonunion members is much larger than for union members, which enables one to assert that once workers become union members they have a lesser incentive to invest in further schooling.

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[12] See also footnote 6.

[13] ($\partial \ln W/\partial U) = 0.6812 - 0.0348(S = 11.92) - 0.01216(E = 22.84) + 0.000224(E^2 = 664.67) - 0.001(UE = 7.96) = 0.13$ (for pooled OLS), and $0.6217 - 0.0375(S = 11.92) - 0.00605(E = 22.84) + 0.000141(E^2 = 664.67) + 0.0009(UE = 7.96) = 0.137$ (for error components model).

[14] ($\partial^2 \ln W/\partial U\ \partial E) = -0.01216 + 2(0.000224)(E = 22.84) = -0.002$ (for pooled OLS), and $-0.00605 + 2(0.000141)(E = 22.84) = -0.0003$ (for error components model).

[15] However, the evidence presented by Polachek et al. (1987), and Wunnava (1988) based on a sub-sample of one time union switchers for PSID 1968–1981 data support an opposing view.

[16] Note that the $U*S$ coefficient is negative and statistically significant for the both pooled regressions. Even for all the 6 cross-sectional OLS regressions (of Table 1) the same phenomenon is observed.

[17] This needs to be qualified. A low rate of return to schooling in the *union* sector may be partially due to union solidarity, standardization of wages, flat or across the board increases in wages, and higher fringe benefits etc. See also Johnson and Youmans (1971), and Lewis (1986) for empirical evidence suggesting that rate of return to education in union sector is lower than nonunion sector. Interestingly, White (1982) offered an alternative explanation that unions tend to lobby mostly around the median rank and file workers (who are relatively less educated) for their own political survival.

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