

Communications

Life Cycle Union Effects Based on a Pooled Regression Technique: Evidence from PSID*

I. Background

Many studies¹ have been done concerning the union-nonunion wage differential. Surprisingly, very few studies focused on the shape and relative slopes of age-earnings profiles of union and nonunion members. However, it has been inferred by those studies which focused on the slope of age-earnings profile that unions raise and flatten the age-earnings profile so as to decrease average returns to seniority.

The issue of interest is to ascertain whether earlier studies tell us the whole story about the union-nonunion wage differential, and the effect of unions on the slope of the age-earnings profile. This is because the individuals who join a union could be different from those who do not. How can we “adjust” for unobserved characteristics to avoid biases due to heterogeneity?

The available alternatives to the pure cross-sectional techniques are the panel data based standard “fixed effects” methods: firstly, the “mean deviation” approach [1], and secondly, the “first differencing” method [8; 10]. The former suffers from the fact that the effect of time invariant variables can not be estimated besides imposing a very strong set of restrictions on the regression coefficients. On the other hand, as pointed out by Freeman [4], the latter technique is polluted by measurement errors. Due to the shortcomings of present techniques we need a new technique to estimate union effects.

A viable alternative may be to use panel data focusing on one-time union switchers² (they could be either leavers, who switch from union to nonunion or joiners, who switch from nonunion to union). By this approach, given individuals can be followed over a period of time long enough to observe their age-earnings profiles before and after they switch from union to nonunion and vice versa. This approach will be taken in this paper.

II. The Model

It has been shown by Mincer and Polachek [9, S79] that:

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1. See Freeman and Medoff [5], and Lewis [7] for excellent surveys of recent literature.

2. Panel data for a sample of one-time union switchers was used by Polachek and McCutcheon [11] to analyse the impact of unions on employment stability. Polachek et al. [12; 13] and Wunnava [14] investigated the impact of unions on wages using a similar sample.

$$\ln Y_t = a_0 + rS + r \int_0^L k(\tau) d\tau \quad (1)$$

where

- Y_t = earnings in time period t ,
- r = the average rate of return to human capital,
- S = years of schooling,
- L = years of labor market experience after schooling,
- k_t = the ratio of \$ investment costs (C_t) to earnings capacity (E_t) or a measure of the fraction of time spent investing at time period t , and
- a_0 = innate earning potential.

Let us suppose that after schooling there are n segments in a given individual's working life (L), i.e., $L = \sum_{j=1}^n e_j$, where e_j is the duration of the j th segment. So (1) can be respecified as (2)

$$\ln Y_t = a_0 + rS + r \sum_{j=1}^n \int_0^{e_j} k_j(\tau) d\tau \quad (2)$$

by assuming that

$$k_j(\tau) = k_{0j} - \beta_j \cdot \tau. \quad (3)$$

Substituting (3) into (2) one can get (4)

$$\ln Y_t = a_0 + rS + r \sum_{j=1}^n \int_0^{e_j} (k_{0j} - \beta_j \cdot \tau) d\tau. \quad (4)$$

Since the focus is on one-time union switchers (all data in the PSID prior to 1968 are retrospective and data following 1968 are current), a natural division occurs in depicting the life cycle. By letting $n = 3$, three segments of post school investment emerge:

- (i). Experience prior to 1968 (= age in 1968 – Schooling in 1968 – 6); let us call it *EXP*.
- (ii). Experience since 1968 to prior to change in union status; let us call it *EB*.
- (iii). Experience after change in union status; let us call it *EA*.

Thus (4) can be simplified as (5):

$$\begin{aligned} \ln Y_t = & a_0 + \alpha_1 S + \alpha_2(EXP) + \alpha_3(EXP)^2 \\ & + \alpha_4(EB) + \alpha_5(EB)^2 + \alpha_6(EA) + \alpha_7(EA)^2 \end{aligned} \quad (5)$$

by adding a union status *CHANGE* dummy to capture the possible discontinuity in the earnings profiles due to a change in union status, and by bringing in a stochastic error term, equation (5) when adapted for panel data gives us the basic empirical specification of the segmented earnings function (equation (6)) used in this paper:

$$\begin{aligned} \ln W_{it} = & a_0 + \alpha_0(CHANGE)_{it} + \alpha_1 S_{it} \\ & + \alpha_2(EXP)_{it} + \alpha_3(EXP^2)_{it} + \alpha_4(EB)_{it} \\ & + \alpha_5(EB^2)_{it} + \alpha_6(EA)_{it} + \alpha_7(EA^2)_{it} + \varepsilon_{it} \end{aligned} \quad (6)$$

$i = 1, 2, \dots, m$ (cross-section), $t = 1, 2, \dots, n$ (time-series)

where W_{it} is the real hourly wages (in 1968 \$) of the i th individual in the t th year. Since only a smaller fraction of individuals are usually one-time union switchers, focusing on such a sub-sample may not yield “global” estimates. Even though “globality” is not the main focus of this paper, any similarity of parametric estimates concerning the effect of unions computed separately for leavers and joiners (who are two diverse groups) shall adequately address this issue. This issue can be further checked by bringing in an inverse Mills ratio term Heckman [6], and Duncan and Leigh [2; 3] as an additional regressor into equation (6). One could generate an appropriate inverse Mills ratio term (hereafter *SELECTIVITY*) by the following steps

(A). Defining $SWITCH_i$ which is an unmeasurable variable that takes the following form:

If $SWITCH_i > 0$, individual i is a one time switcher.

If $SWITCH_i < 0$, individual i is from the omitted category.

(B). For the i th individual, the $SWITCH$ threshold equation is defined as

$$SWITCH_i = \mathbf{X}_i\boldsymbol{\beta} + \varepsilon_i.$$

\mathbf{X}_i includes all of the exogenous variables in the model.

This reduced form threshold function determining sample selection into $SWITCH$ and non $SWITCH$ sectors can be estimated using a Weighted Nonlinear Least Squares Probit (WNLSP) method.

(C). Then for the $SWITCH$ and non $SWITCH$ sectors, $SELECTIVITY$ variables are calculated as:

$$-f(SWITCH)/F(SWITCH) \quad \text{and} \quad f(SWITCH)/[1 - F(SWITCH)] \quad \text{respectively.}$$

$F(\cdot)$ is the cumulative distribution of a standard normal variable, and $f(\cdot)$ is its density function.

III. Empirical Results

Currently, the longest available longitudinal data set containing information on wages, unionism, and other standard human capital variables is the University of Michigan: Panel Study of Income Dynamics (PSID). Data from 1968–81 will be used for the empirical results of this paper. With the PSID, a sample of 946 white male heads of household (who had consistent union and wage data) could be followed. The break-down is as follows: 113 always union members, 518 always non-union members, 222 multiple switchers, and 93 one-time switchers (63 leavers who switched from union to nonunion, and 30 joiners who switched from nonunion to union). The methodology developed in this paper requires that we focus only on a sub-sample of one-time switchers.³ Pooling the data over cross-section and time-series requires non OLS estimation procedures because of potential correlation among disturbances. In this paper ‘error variance component’ technique will be used. Modified Generalized Least Squares (MGLS) estimates of the segmented earnings functions (as specified in equation (6)) for union joiners and leavers are presented in Table I.

3. Because given individuals need to be followed over a period of time long enough to observe the age-earnings profile before and after they shift from union to nonunion and vice versa.

Table I. MGLS Estimates: Segmented Earnings Regressions (Dependent variable: In real hourly wages, in 1968 dollars)

Independent variables	Joiners		Leavers	
	(1)	(2)	(3)	(4)
Intercept	4.9 (40.7)	4.7 (6.9)	5.13 (36.8)	5.74 (23.6)
<i>CHANGE</i> (Dummy)	-.206 (2.64)	-.157 (.1)	-.103 (1.5)	.066 (.75)
<i>S</i>	.069 (9.94)	.062 (.25)	.043 (4.9)	.07 (5.5)
<i>EXP</i>	.025 (3.80)	.026 (3.50)	.016 (3.3)	.015 (3.1)
<i>EXP</i> ²	-.0004 (3.30)	-.0005 (1.00)	-.00035 (3.4)	-.0003 (2.7)
<i>EB</i>	-.08 (3.80)	-.08 (3.20)	.045 (1.90)	.052 (2.10)
<i>EB</i> ²	.008 (4.00)	.008 (4.00)	-.002 (1.40)	-.003 (1.90)
<i>EA</i>	.06 (2.40)	.06 (2.30)	-.02 (1.10)	-.024 (0.90)
<i>EA</i> ²	-.002 (0.95)	-.002 (0.95)	.0015 (0.70)	.001 (0.50)
<i>SELECTIVITY</i>		-.157 (0.03)		.063 (2.90)
DF	411	410	873	872
<i>R</i> ²	.326	.337	.12	.13
<i>F</i> -Ratio	24.4	23.1	14.9	14.4

Note: *t*-statistics in parentheses.

Union Slope Effects

For joiners, there seems to be a 2.5% increase⁴ in the life cycle earnings profile. Leavers seem to have a decrease⁵ of 2.4% to 2.9% for leaving a union. In essence, the estimate of slope in the

4. Slope in the nonunion segment:

$$\partial \ln W / \partial EB = \alpha_4 + 2\alpha_5 EB = -.08 + 2(.008)5.8 = .012.$$

Slope in the union segment:

$$\partial \ln W / \partial EA = \alpha_6 + 2\alpha_7 EA = .06 + 2(-.002)5.8 = .037.$$

Thus, 3.7% - 1.2% = 2.5%. Interestingly, selectivity correction also yields identical results.

5. Slope in the union segment:

$$\partial \ln W / \partial EB = .045 - 2(.002)5.8 = .022$$

with selectivity correction, .052 - 2(.003) 5.8 = .017.

Slope in the nonunion segment is:

union segment for both groups is positive and larger than the respective slope estimates in the nonunion segment 3.7% versus 1/2% with or without selectivity for joiners; 2.2% (1.7% with selectivity) versus -.2% (1.2% with selectivity) for leavers. This enables us to conclude that, if anything, unions seem to steepen the age-earnings profiles for both joiners as well as leavers.

Union-Nonunion Wage Differential

Regression results also indicate that union-nonunion wage gaps estimated in a life cycle context for joiners⁶ are in the magnitude of 20.5 to 25.5%. Leavers⁷ loss seem to be between 22 to 31%. Since most union switches are accompanied by a change in employment and also the theories of specific training indicate that there will be an initial loss in human capital and hence a decrease in earnings, Mincer [8], Polachek et al, [12; 13], and Wunnava [14] feel that above wage gap estimates need to be adjusted *upward* for joiners and *downward* for leavers by appropriate employment change effects.⁸

Selectivity Correction

Regression results with *SELECTIVITY* variables (presented in columns 2 and 4 of Table I) are very similar to the estimates obtained without selectivity (columns 1 and 3 of Table I). The only noticeable differences for joiners are a drop in the *t*-statistic for the intercept (from 40.7 to 6.9)

$$\partial \ln W / \partial EA = -.02 + 2(.0015)5.8 = -.002,$$

with selectivity, $-.024 + 2(.001)5.8 = .012$.

Thus, $2.2\% + .2\% = 2.4\%$, with selectivity: $1.7\% + 1.2\% = 2.9\%$.

6. The effect of unions on the levels of age-earnings profiles is estimated by the difference between union and nonunion segments evaluated at the mean years of switch. After cancelling the common coefficients in both segments, it amounts to:

$$\begin{aligned} & [\alpha_0 + \alpha_6(t) + \alpha_7(t^2)] - [\alpha_4(t) + \alpha_5(t^2)]; \\ & [-.21 + .06(5.8) - .002(38.6)] - [-.08(5.8) + (.008)(38.6)] = .205; \\ & [-.16 + .06(5.8) - .002(38.6)] - [-.08(5.8) + (.008)(38.6)] = .255 \end{aligned}$$

(with selectivity).

7. Using the same approach as joiners, we get

$$[.066 - .024(5.8) + .001(38.6)] - [.052(5.8) - .003(38.6)] = -.22$$

(with selectivity);

$$[-.1 - .02(5.8) + .0015(38.6)] - [.045(5.8) - .002(38.6)] = -.31.$$

Even though the partial derivatives (involving more than one regression coefficient) reported in this paper are not individually significant, a standard Chow test indicated that they are indeed collectively significant at 10%.

8. It may be true that other variables change as union status changes but with a limited sample of one-time union switchers, it would be difficult to control for these effects since all probably occur simultaneously. Even though the dependent variable ($\ln W$) is measured in real terms (1968 dollars) there may still remain the counter cyclical influence of unions on the level and slope of age-earnings profile, see Mellow [10] and Freeman [4]. I wish to thank the referee for making this point.

and for the S or schooling coefficient (from 9.94 to .25). For leavers t value dropped from 36.8 to 23.6 for the intercept. However, neither values nor significance level of the other coefficients have changed.⁹

Inverse Mills ratio estimates of coefficients on the selectivity variables (labeled as *SELECTIVITY*) are presented at the bottom of columns 2 and 4 of Table I. By defining *SELECTIVITY* as negative $[-f(\cdot)/F(\cdot)]$ for one-time switchers (i.e., for joiners and leavers) means that a negative coefficient estimate is required to provide evidence of positive selectivity. For joiners (*SELECTIVITY* coefficient = $-.157$ with $t = .03$) does provide evidence that wage distribution associated with these workers is not significantly different from the wage distribution that would be observed for an individual selected at random from a sample of (always union, always non-union, and switchers) workers with the same personal and job related characteristics. However, for leavers there seem to be a sort of negative selectivity (*SELECTIVITY* coefficient = $.63$ with $t = 2.9$). Unlike the apprehensions raised by Freeman and Medoff [5], and Lewis [7] concerning the robustness of this technique, it seems to be effective in the present context (as results without selectivity correction are not different with selectivity correction).

IV. Conclusion

This paper circumvents some of the basic problems of earlier cross-sectional and panel estimates (i.e., heterogeneity bias, selectivity bias, bias due to measurement errors) of life cycle union effects by using longitudinal data (Panel Survey of Income Dynamics) focusing on one-time switchers (who change their union status only once between 1968–81). These one-time switchers are divided into two groups: leavers (who switch from union to nonunion) and joiners (who switch from nonunion to union). The model (a segmented earnings specification) then compares the age-earnings profile of a representative worker while he is in the union(nonunion) with his profile in nonunion(union). The Heckman's selectivity correction adjustment is also used so as to avoid the problem of a truncated/restricted sample of one-time switchers. Estimates of segmented earnings regressions computed separately for leavers and joiners indicate that:

- (A). Lifetime union effects on the level of age-earnings profiles is about 25%.
- (B). There is no evidence to support the claim that unions flatten the age-earnings profiles.

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9. The only other exception is the value of dummy intercept for leavers. It switched gears from a value of $-.1[t = 1.5]$ to $.066[t = .75]$.

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