Unions and Employment Growth: Panel Data Evidence

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This study adds to the small empirical literature on the impact unions have on employment growth using data from Australia. Unlike previous studies, the data used are from a panel of firms surveyed at two points in time rather than a single cross section. The results indicate a negative union effect on employment in private-sector firms of about 2.5 percent per annum that, despite the very different institutional framework that prevails in Australia, is consistent with results obtained with North American data.

A NUMBER OF RECENT STUDIES confirm what economists have long suspected—that unions slow job growth. Blanchflower et al. (1992), for example, used workplace-level data collected as part of the British Workplace Industrial Relations Survey in 1984 to estimate an employment equation that controlled for the effects of union recognition (as well as a small number of other potentially confounding influences). They concluded that, over the period 1980 to 1984, employment in unionized establishments in the United Kingdom grew by 3 percent less per annum than employment in nonunion establishments.[1] Strikingly similar conclusions

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[1] Machin and Wadhwani (1991) estimated employment equations using the same data used by Blanchflower et al. (1991) but concluded that any negative impact of unions on employment was confined to workplaces that reformed working conditions during the period under consideration (1980–1984). These strikingly different conclusions are the subject of substantial debate within the articles themselves. We are persuaded by Blanchflower et al., who found, after replacing the simple union recognition dummy with a union density variable, that unions exert a negative effect on employment irrespective of whether organizational change occurred. Furthermore, as Blanchflower et al. (1991:829) noted, "the organisational change variable almost certainly captures part of the transmission mechanism by which unions lead to employment decline."

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have been reached in studies employing both U.S. and Canadian data. Using a sample of 1800 Californian manufacturing plants, Leonard (1992) reported that employment growth in union plants was between 2 and 4 percentage points less than in nonunion plants, whereas Long (1993), in a study of 510 Canadian firms, estimated a union employment growth differential of just under 4 percent per annum. Somewhat differently, Bronars et al. (1994) reported that firm-level unionization rates were significantly and inversely associated with employment growth over the period 1975 to 1982 in a small sample of publicly traded U.S. companies, with employment growth estimated to decline by between ½ and 1 percentage point in response to a 10 percent increase in unionization.

More recently, Blanchflower and Burgess (1996) reported the results of estimating employment growth equations using workplace-level data from both the United Kingdom and Australia. While they found that the presence of unions tended to have a restraining effect on employment growth in the U.K. sample, they found no evidence of a discernible impact of unions in the Australian data.

The present study reexamines the issue of employment growth and the role unions in Australia may have played in either fostering or retarding that process using data collected as part of the Australian Workplace Industrial Relations Surveys conducted in 1989–1990 and 1995. Consistent with the earlier North American and U.K. studies, the central feature of the analysis is the use of regression methods to isolate the union employment effect from other potentially confounding influences. Unlike the earlier studies, however, the data used are from a panel of firms surveyed in both 1989–1990 and 1995 rather than a single point-in-time cross section (where past employment is measured retrospectively). As a result, this study is able to control not just for the fixed characteristics of firms and workplaces but also for changes in the economic environment and in managerial strategy.

Unions and Employment: Theoretical Background

Following Long (1993), at least three routes through which unions reduce employment growth can be identified. First, through their ability to withhold labor supplies, unions are presumed to seek and capture a share of monopoly rents for their members. This is reflected in the presence of a substantial union wage markup that, in the United States, typically has been estimated as lying in the range of 10 to 15 percent (Hirsch and Addison, 1986:152–3). Given downward-sloping demand schedules, unionized firms therefore should tend to employ smaller quantities of labor than comparable nonunionized firms.

Second, unions may make it relatively difficult for firms to put into effect downward adjustments in the size of their workforce (through, for example, expensive redundancy packages and more costly and cumbersome dismissal procedures), which in turn will tend to discourage firms from expanding their workforce. In a static model, this may not affect long-run employment growth; a lower propensity to hire should be offset by a lower propensity to fire. However, if growth is endogenous, then long-run employment outcomes may be adversely affected.

Third, unions may adversely affect growth in sales, which in turn will inhibit employment growth. This might occur if unions reduce the incentive for firms to invest in new capital, reduce the scope for price cutting in an effort to maintain sales (especially when overcapacity exists), interrupt the reliability of supply (through frequent work stoppages), or impose restrictive work rules and practices.

In contrast, there are at least two counterarguments that suggest the possibility that unions actually could contribute to higher rates of employment growth. While unions raise wages, they also may raise worker productivity as a result of “voice” effects. Unions provide a mechanism for channeling grievances to management with a smaller risk of individual workers being victimized (Freeman and Medoff, 1984). This, it is claimed, can be beneficial for productivity by reducing the incidence of costly labor turnover, enhancing the incentives for both employers and employees to invest in training, improving communication flows, and raising overall worker morale. Unit labor costs, therefore, may not necessarily be any higher in the presence of unions.

Finally, if unions and employers bargain simultaneously over wages and employment rather than over wages alone, then efficient outcomes are unlikely to lie on the labor demand curve (see, for example, Hall and Lilien, 1979; McDonald and Solow, 1981). If this is so, the effects of unions on employment outcomes will be ambiguous. That said, the extent to which efficient settlements will diverge from the demand curve depends on how much weight unions place on employment.

The emerging consensus in North America and the United Kingdom is that negative union employment effects outweigh the positive effects. This reflects a combination of

1. The presence of a large union wage markup.
2. Growing evidence that any union productivity effect is either negative or very small (Kaufman and Kaufman, 1987; Lovell et

al., 1988; Machin, 1991; Wilson and Cable, 1991; Mitchell and Stone, 1992; Byrne et al., 1996).

3. Evidence that unions place a relatively low weight on employment compared with wages (MaCurdy and Pencavel, 1986; Wessels, 1991).

It cannot be presumed, however, that such conclusions necessarily will carry over to other economies where labor market institutions and practices may be very different. As noted earlier, Blanchflower and Burgess (1996) reported an insignificant union effect on employment in Australian data. While this result may reflect the short period over which employment change was observed—just 1 year—there are other good reasons why a weak union employment effect in Australia may have been expected.

Most important, it is not obvious that Australian unions have been able to exert a significant intraindustry effect on wages. As discussed by Wooden and Baker (1994:405), employment conditions for most Australian workers, at least until recently, have been determined largely by industry and occupational "awards" that provide for legally enforceable minimum rates of pay (as well as a range of minima with respect to other employment conditions) that apply to both union and nonunion workers employed in the industries and occupations covered by those awards. Moreover, the coverage of these awards is extensive, with 80 percent of all employees in Australia in 1990 estimated to have been covered. It thus follows that union-negotiated increases in award rates of pay will affect union and nonunion workers equally.

Within the awards system, however, scope traditionally has existed for unions to engage in informal "over-award" bargaining, and hence the presence of the awards system per se does not automatically mean that a union wage premium cannot exist. This said, this was not true of the period covered by the Blanchflower and Burgess (1996) study—1988 to 1989. The centerpiece of federal economic policy during this time was the Accord, a consensual-type incomes policy in which unions agreed not to pursue additional wage claims outside those provided by the centralized industrial tribunals through the awards mechanism (see Chapman and Gruen, 1990). In return, the union movement (and its membership) was to benefit from improvements in the "social wage" (as a result of the introduction of and improvement in a range of health and social welfare measures), a much enlarged role in policymaking, and more generally, the benefits of economic growth that were presumed to follow. In this climate, it follows that union members will not fare any better in terms of wages and other employment conditions than nonunion members who are also covered by the awards system. This, in turn, implies no difference in employment outcomes.

On the other hand, industrial relations structures in Australia have been undergoing enormous change during the 1990s, the period covered by the analysis reported in this study. As documented by Quinlan (1996), Callus (1997), and Hawke and Wooden (1998), the 1990s has seen both institutional and legislative change facilitating increased scope for more decentralized forms of bargaining and especially enterprise-level agreements. Hawke and Wooden (1998), for example, report that in excess of 1.7 million Australian employees were covered by enterprise agreements within the federal jurisdiction by September 1996. To this figure can be added another 800,000 workers covered by enterprise agreements under state jurisdictions (Joint Governments' Submission, 1997). In total, therefore, around 36 percent of Australian employees would appear to have been covered by enterprise agreements by late 1996. In contrast, such arrangements were rare in the 1980s. In this environment it seems plausible that wage differentials between union and nonunion workers might emerge and union employment effects become more evident.

The Data

The data used in this analysis were collected from Australian workplaces during 1989–1990 and 1995 as part of the Australian Workplace Industrial Relations Surveys (AWIRS). Described in more detail in Callus et al. (1991) and Morehead et al. (1997), the AWIRS involved a suite of structured questionnaires administered by a variety of methods to managers, union delegates, and (in 1995) employees at representative samples of Australian workplaces. The samples for both surveys were randomly selected from the *Australian Bureau of Statistics Business Register*, after stratification by location, size, and industry. The survey covered all industry sectors with the exception of agriculture, forestry, and fishing and defense. The scope of the survey also was restricted to workplaces with at least five employees, though researchers often were compelled to ignore workplaces with 5 to 19 employees because far less information was collected from the subsample of small workplaces.

The 1995 AWIRS also included a panel component. That is, a sample of workplaces from the 1989–1990 "main sample" (2004 workplaces with 20 or more employees) that participated in the 1989–1990 survey was selected to be resurveyed in 1995.[2] The 1989–1990 sample was

[2] The 1989–1990 sample of 2004 workplaces was achieved from 2300 contacts, giving a response rate of 89 percent.

screened for survivors, and 780 workplaces then were selected. Interviews were conducted successfully at 698 of these workplaces, giving a response rate of 89 percent. Compared with the cross-sectional data analyzed in previous studies, panel data provide a more accurate assessment of employment levels at different points in time. Moreover, they better enable the identification of those influences which are fixed and those which vary over time.

The panel is not, however, a random sample of participants in the first survey for the obvious reason that it only includes firms that survived from 1989 to 1995 (representing 86 percent of the original workplace sample). Some analysis of survivors and “deaths” is provided in Morehead et al. (1997:48–51) that indicates that “deaths” were more likely among workplaces that were small, relatively young (less than 2 years old), part of a larger organization, part of a government business enterprise, had not been performing well in 1989–1990 (as indicated by low rates of capacity utilization and negative rates of return on assets), and had already been in the process of downsizing at the time of the 1989–1990 survey. There was, however, no difference in the “death” rate of workplaces with and without union members. Further analysis of these data also revealed that there were no significant differences in the mean level of union density (measured in 1989–1990) at surviving workplaces and at those which had “died.”

The employment data used in this analysis come from self-completion questionnaires that were mailed to participating workplaces prior to interview and relate to the total number of employees working at or from the workplace during the pay periods ended on or before September 30, 1989 and on or before August 18, 1995. Respondents were given instructions to include managers, all employees on paid leave, and all employees on workers’ compensation and to exclude contractors, agency workers, and home workers working on a contract-for-service basis.

Table 1 cross-classifies a selection of workplace characteristics by workplace employment levels in 1995 relative to employment in 1989. The elements in this table sum horizontally to 100 percent. Thus, for example, almost 10 percent of all workplaces declined in size by 50 percent or more between 1989 and 1995. It also can be seen over this period that the number of workplaces where employment fell by more than 10 percent exceeded those where it rose by more than 10 percent. These findings are not surprising and reflect the impact of the economic downturn on the Australian labor market during the early 1990s.

TABLE 1
EMPLOYMENT GROWTH OF WORKPLACES, 1989–1995, BY WORKPLACE CHARACTERISTICS (PERCENT)

	1995 Employment Level Relative to 1989 Level					
	<50%	50% to <90%	Within 10%	>10% to 50% Higher	>50% Higher	N^a
All workplaces	9.8	36.5	22.4	19.5	11.8	698
Workplace size, 1989 (no. of employees)						698
<50	7.4	37.3	21.1	20.9	13.2	
50–99	12.3	34.0	24.9	17.3	11.5	
100–299	10.9	35.7	23.4	21.1	8.8	
300 or more	14.9	43.0	18.5	12.7	10.9	
Firm size, 1989–1990 (no. of employees)						607
<100	10.3	43.4	9.7	23.6	13.0	
100–999	12.4	28.3	24.8	19.3	15.2	
1000 or more	9.7	39.2	21.5	18.4	11.1	
Organization status						698
Private sector	10.8	38.7	20.6	17.4	12.5	
Public commercial	30.9	31.1	20.3	6.5	11.3	
Public noncommercial	3.8	33.0	26.5	26.3	10.3	
Union density, 1989–1990						679
Zero	3.3	37.9	26.8	17.6	14.3	
>0% to <50%	12.0	30.6	17.8	26.2	13.4	
50% to <75%	11.0	34.8	25.5	18.5	10.2	
>75%	12.0	39.6	19.8	18.4	10.1	

[a]While unweighted sample sizes are reported here, the percentages presented in the body of the table are derived after applying sample weights to ensure a representative sample.
SOURCE: 1995 Australian Workplaces Industrial Relations Survey, panel component.

Table 1 suggests that workplace growth is associated with both workplace size and, more obviously, organization status. Only 7 percent of the smaller workplaces (less than 50 employees) experienced a large fall in employment, compared with 15 percent of large workplaces (300 employees or more), whereas at the other end of the spectrum, small workplaces were more likely to be growing rapidly, although the difference is not large. Relationships with overall firm size, as distinct from workplace size, were far less obvious, suggesting that part of the explanation for the weaker employment growth of large workplaces may be larger enterprises opting for smaller, more decentralized business units.

The growth experience of private and private-sector organizations and, more important, public commercial (i.e., government business enterprises and commercial statutory authorities) and public noncommercial organizations (e.g., government departments) also appears to have been very different over the period under examination. Public-sector organizations operating on a commercial basis were much more likely to have downsized and far less likely to have experienced rapid growth. Public noncommercial organizations, on the other hand, appear to have avoided large-scale employment reductions.

Finally, and of most relevance to this study, Table 1 provides little evidence of any direct relationship between employment change and unionization levels, especially among the large majority of workplaces where at least one employee is a union member. Such findings are consistent with the conclusions reached by Blanchflower and Burgess (1996). It now remains to be seen whether this conclusion holds once other influences on employment growth are taken into account.

Modeling Employment Growth

The statistical analysis revolves around the estimation of a type of reduced-form equation where, following Blanchflower and Burgess (1996), the dependent variable is growth in employment rather than its level. *Growth g* is defined as the change in employment as a percentage of the average of employment in the two periods. That is,

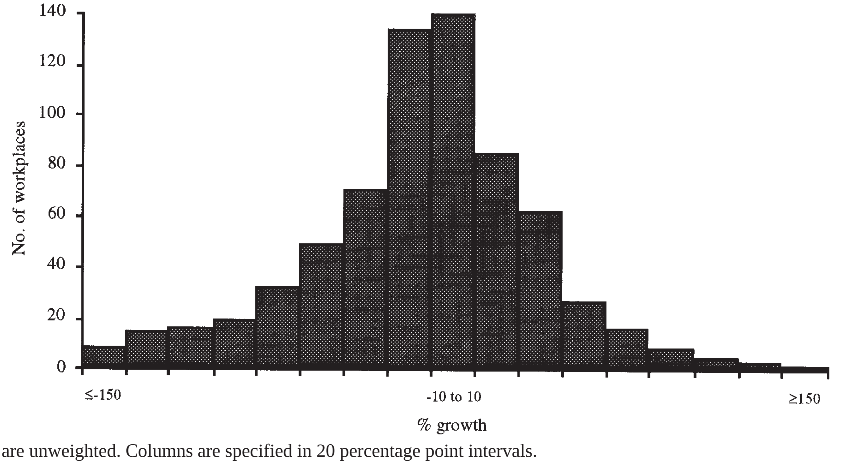
$$g = (E_{95} - E_{89}) / \left[(E_{89} + E_{95}) / 2\right] \times 100$$

where E denotes employment, and the subscripts refer to the year in which employment is observed.

Unlike the conventional growth measure (i.e., change in employment as a percentage of employment in the first period), this measure is approximately normally distributed (Davis and Haltiwanger, 1992). As indicated in Figure 1, the distribution in the values of g across the workplaces in the

FIGURE 1

DISTRIBUTION OF EMPLOYMENT GROWTH RATES[a]



[a] Data are unweighted. Columns are specified in 20 percentage point intervals.

sample (after weighting to account for sample stratification) appears to be close to symmetric. Further evidence that the distribution is close to normal is provided by an examination of skewness and kurtosis statistics. Descriptive statistics for both growth measures are reported in Table 2. An exact normal distribution exists where the skewness statistic has a value of 0 and the kurtosis statistic has a value of 3. This is clearly not the case for the conventional growth measure, which is both highly skewed and leptokurtic (relatively few cases concentrated in the tails of the distribution). In contrast, the preferred growth measure shows no signs of skewness, although a slight playtokurtic tendency (too many cases concentrated in the tails of the distribution) is apparent.

The effects of unions were modeled, at least initially, with two variables: (1) a simple dummy variable indicating the presence or absence of union members within the workforce in 1989 and (2) the density of union membership within the workplace in 1989. A range of controls for other influences on employment growth also was tested, with the control variables included in the specifications reported in this article being

1. Workplace employment in 1989
2. A series of dummy variables to control for market demand conditions in 1989 and 1995
3. Measures of the degree of autonomy workplace management has from higher levels in the organization and change in the level of that autonomy
4. A small number of variables to control for the effects of other firm characteristics
5. Dummy variables to control for changes in the workplace location, whether the workplace was part of a merger during 1989–1995, and whether the workplace was defined differently across the two surveys

TABLE 2

EMPLOYMENT GROWTH MEASURES: DESCRIPTIVE STATISTICS[a]

Descriptive Statistic	Conventional Growth Measure $(E_{95} - E_{89})/E_{89} \times 100$	Preferred Growth Measure $(E_{95} - E_{89})/[(E_{89} + E_{95})/2] \times 100$
Mean	4.86	–7.66
Standard deviation	70.01	46.01
Skewness	7.79[b]	–0.22
Kurtosis	120.90[b]	1.25[b]

[a]Unweighted $N = 698$.
[b]Significant at the 0.05 level in a two-tailed test.

As discussed further below, a control for workplaces where union membership declined to zero between 1989 and 1995 also was added to the final model specification. A description and statistical summary of the explanatory variables used are provided in Table 3. Note that exclusion of all cases with missing observations on any variable of interest resulted in a final sample size numbering 670 cases.

TABLE 3

EXPLANATORY VARIABLES: DESCRIPTIONS AND SUMMARY STATISTICS[a]

		Mean (SD)		
Variable Name	Description	Total Sample	Private Sector	Public Sector
Employment 89	Number of employees at the workplace in 1989	266.28 (512.04)	231.40 (463.78)	310.62 (565.20)
Union	A dummy variable for workplaces where in 1989–1990 any union members were present	0.881 (0.325)	0.813 (0.390)	0.966 (0.181)
Union density	Union members in 1989–1990 as a percentage of total employment in 1989	63.384 (32.564)	54.897 (35.807)	74.172 (23.944)
Deunionized	A dummy variable for workplaces where union membership fell to zero between 1989 and 1995	0.048 (0.213)	0.077 (0.267)	0.010 (0.101)
Demand1	A dummy variable for workplaces where demand for main product was expanding in 1989–1990 and in 1995	0.327 (0.469)	0.269 (0.444)	0.400 (0.491)
Demand2	A dummy variable for workplaces where demand for main product was expanding in 1989–1990 but stable in 1995	0.196 (0.397)	0.181 (0.386)	0.214 (0.411)
Demand3	A dummy variable for workplaces where demand for main product expanding in 1989–1990 but contracting in 1995	0.054 (0.226)	0.061 (0.240)	0.044 (0.206)
Demand4	A dummy variable for workplaces where demand for main product was contracting in 1989–1990 but expanding in 1995	0.036 (0.186)	0.037 (0.190)	0.034 (0.181)
Demand5	A dummy variable for workplaces where demand for main product was contracting in 1989–1990 but stable in 1995	0.061 (0.240)	0.085 (0.280)	0.031 (0.172)
Demand6	A dummy variable for workplaces where demand for main product was contracting in 1989–1990 and in 1995	0.031 (0.174)	0.037 (0.190)	0.024 (0.152)

TABLE 3–*continued*

Variable Name	Description	Mean (SD) Total Sample	Private Sector	Public Sector
Demand7	A dummy variable for workplaces where demand for main product was fairly predictable in 1989–1990 but largely unpredictable in 1995	0.139 (0.346)	0.152 (0.360)	0.122 (0.328)
Autonomy	An index of the level of autonomy of the workplace from higher levels in the organization (scored from 0 to 100), measured in 1989–1990[b]	78.271 (26.752)	87.501 (18.543)	66.537 (30.730)
Change in autonomy	Change in the level of autonomy between 1989–1990 and 1995	0.886 (22.217)	–2.624 (18.330)	5.348 (25.692)
Single-workplace organization	A dummy variable for single workplace organizations (measured in 1989–1990)	0.115 (0.319)	0.163 (0.370)	0.054 (0.227)
Head office	A dummy variable for workplaces that were also the organization's head office (measured in 1989–1990)	0.334 (0.472)	0.400 (0.491)	0.251 (0.434)
Exporter	A dummy variable for workplaces which produced primarily for the export market (measured in 1989–1990)	0.021 (0.143)	0.037 (0.190)	0.000 (0.000)
Importer	A dummy variable for workplaces which produced primarily for the domestic market and faced import competition (measured in 1989–1990)	0.187 (0.390)	0.309 (0.463)	0.031 (0.172)
Private sector	A dummy variable for workplaces that were part of a private sector enterprise	0.560 (0.497)	1.000 (0.000)	0.000 (0.000)
Public commercial	A dummy variable for workplaces that were part of a public sector commercial enterprise	0.140 (0.348)	0.000 (0.000)	0.319 (0.467)
Merger	A dummy variable for workplaces that had merged with other workplaces since 1989–1990	0.012 (0.109)	0.005 (0.073)	0.020 (0.141)
Changed location	A dummy variable for workplaces that had changed physical location since 1989–1990	0.128 (0.335)	0.107 (0.309)	0.156 (0.363)
Definitional change (A)	A dummy variable for workplaces that were defined more widely in 1995	0.004 (0.067)	0.008 (0.089)	0.000 (0.000)
Definitional change (B)	A dummy variable for workplaces that were defined more widely in 1989–1990	0.006 (0.077)	0.003 (0.052)	0.010 (0.101)

[a] Unweighted $N = 670$.
[b] For more details on the construction of this variable, see Morehead et al. (1997, fn 23, p. 102). The index has been reflected so that a score of 100 indicates complete autonomy.

Experimentation also was undertaken with variables not reported in the article. Included here are industry, occupational mix of the workforce, workplace age, firm size (as distinct from workplace size), foreign ownership, and measures of the degree of product market competition and changes in the level of that competition. None of these variables exhibited any significant explanatory power, and they were omitted subsequently from the analysis.

The equations are estimated using ordinary least squares (OLS).

Results

The main OLS results using the total sample are presented in columns 1 and 2 of Table 4. Since there is evidence of heteroskedasticity in all specifications, the *t* ratios have been adjusted accordingly.[3]

Looking first at the short specification in column 1, the union dummy is found to be positively signed but insignificant. This result, however, is not of large interest given that only 12 percent of workplaces in the sample do not have union members present. More interesting is the variable union density. Consistent with the North American and U.K. research, union density is negatively signed and highly significant. Moreover, the estimated coefficient implies that a workplace with the mean level of union density (63.4 percent) experienced a rate of employment growth during 1989–1995 that was 15 percentage points less than that of an otherwise comparable workplace with an almost zero level of union density, or about 2½ percent per annum. This result is thus consistent (although at the lower end of the range) with estimates obtained with North American and U.K. data.

Column 2 reports the results of estimating a longer specification in which a larger number of controls are added to the equation. In this specification, the coefficient on the union density variable is much smaller (although still sizable) and is no longer statistically significant, suggesting that the significant negative union effect on employment growth in specification (1) may have been due to correlation with relevant omitted variables. Note, however, that diagnostic testing suggests that the functional form of this specification is problematic.

The results on the other control variables are, for the most part, in line with expectations. Mergers between workplaces, of course, by definition, mean increased employment, while the negative sign on employment 89 indicates that small workplaces grow more rapidly than large workplaces.

[3] All equations have been estimated using LIMDEP (version 7.0). The estimated covariance matrix to correct for heteroskedasticity is based on the estimator devised by White (1978).

TABLE 4

EMPLOYMENT GROWTH EQUATIONS: OLS RESULTS

	Total Sample		Private Sector		Public Sector	
	(1)	(2)	(3)	(4)	(5)	(6)
Employment 89	−0.82**	−1.30***	−1.60***	−1.60***	−1.12**	−1.06**
(coefficient × 100)	(2.40)	(3.49)	(3.00)	(2.97)	(2.18)	(2.08)
Union	11.55	4.95	6.83	15.21*	3.05	0.28
	(1.41)	(0.63)	(0.77)	(1.69)	(0.20)	(0.02)
Union density	−0.24***	−0.14	−0.18*	−0.27**	−0.008	0.007
	(2.70)	(1.47)	(1.72)	(2.50)	(0.05)	(0.04)
Deunionized				−25.47**		45.47
				(2.12)		(1.38)
Merger	63.79***	60.06***	31.14	30.20	65.10***	65.34***
	(5.15)	(3.69)	(1.04)	(0.95)	(3.42)	(3.43)
Changed location	−9.35	−13.17*	−20.23*	−18.99*	−3.99	−3.40
	(1.22)	(1.76)	(1.86)	(1.73)	(0.40)	(0.34)
Definitional change (A)	−14.50	−18.28	−13.38	−12.28		
	(0.23)	(0.28)	(0.20)	(0.18)		
Definitional change (B)	−1.15***	−1.09***	−1.73***	−1.51***	−0.86**	−0.86**
(coefficient/100)	(5.00)	(3.93)	(20.69)	(10.17)	(2.57)	(2.57)
Demand1		7.18	3.69	2.26	13.22	13.98*
		(1.50)	(0.64)	(0.40)	(1.65)	(1.74)
Demand2		7.08	2.92	1.73	16.78*	17.24*
		(1.19)	(0.40)	(0.24)	(1.71)	(1.77)
Demand3		−6.77	−13.29	−14.22*	−0.60	0.32
		(0.83)	(1.57)	(1.73)	(0.04)	(0.02)
Demand4		−25.40**	−36.18**	−35.83**	−15.88	−15.36
		(2.57)	(2.36)	(2.46)	(1.47)	(1.43)
Demand5		−9.35	−18.76*	−19.73*	14.36	15.01
		(0.89)	(1.79)	(1.94)	(0.49)	(0.51)
Demand6		−19.50	4.37	6.90	−65.33***	−64.84***
		(1.43)	(0.31)	(0.52)	(3.73)	(3.71)
Demand7		−10.88*	−20.72***	−17.95***	2.16	3.11
		(1.90)	(3.08)	(2.83)	(0.22)	(0.32)
Autonomy		0.27**	0.35**	0.35**	0.27*	0.26
		(2.40)	(2.12)	(2.17)	(1.69)	(1.64)
Change in autonomy		0.45***	0.44***	0.42***	0.51***	0.51***
		(4.07)	(2.93)	(2.85)	(3.12)	(3.13)
Single-workplace		−0.82	−6.15	−4.96	1.33	−4.72
organization		(0.12)	(0.79)	(0.63)	(0.16)	(0.36)
Head office		−6.11	−5.13	−5.03	−10.74	−10.24
		(1.05)	(0.74)	(0.74)	(1.06)	(1.02)
Exporter		2.03	1.70	4.56		
		(0.23)	(0.20)	(0.51)		
Importer		−9.99**	−11.76**	−12.66**		
		(2.00)	(2.30)	(2.52)		
Private sector		−6.99				
		(1.37)				
Public commercial		−30.52***			−31.97***	−31.83***
		(3.92)			(4.11)	(4.15)
Constant	−4.83	−13.18	−19.39	−19.96	−28.44*	−27.55*
	(1.00)	(1.25)	(1.31)	(1.38)	(1.71)	(1.68)
Adjusted R^2	0.056	0.129	0.125	0.139	0.150	0.153
Diagnostics:						
Model fit (F)	6.65***	5.50***	3.67***	3.87***	3.89***	3.79***
Heteroskedasticity (χ^2)	51.89***	121.22**	105.46***	121.29***	45.60***	43.00***
Functional form (F)	1.35	4.36***	0.55	1.42	3.80**	4.06***
N	670	670	375	375	295	295

NOTE: Heteroskedastic consistent t ratios are in parentheses.
*Significant at the 0.10 level in a two-tailed test.
**Significant at the 0.05 level in a two-tailed test.
***Significant at the 0.01 level in a two-tailed test.

The positive sign on autonomy indicates that workplaces that are tightly controlled by higher levels within the organization are far less likely to have experienced employment growth compared with workplaces that are relatively autonomous. This finding thus suggests that centralized, bureaucratic organizational structures work against employment growth. Further, the results suggest that employment growth has been affected not just by the level of managerial autonomy within the workplaces but also by changes in that level over the survey period. Finally, large differences were found between the employment experiences of workplaces that were part of private-sector organizations, those which were part of public commercial enterprises, and those which were part of public noncommercial enterprises. As noted earlier in the discussion of Table 1, employment growth appears to have been strongest among the latter (the control group) and weakest (and typically declining) at workplaces that were part of government-operated commercial enterprises.

Further analysis revealed that the union variable was particularly sensitive to the inclusion of the private-sector dummy variable, and hence the equations were reestimated after splitting the sample into its private- and public-sector components. These results are reported in columns 3 and 5, respectively, and reveal that the magnitude of the impact of union density on employment is very different in the two sectors. In the public-sector subsample, there is no evidence of any negative union impact on employment. In contrast, in the private sector the effect is sizable (10 percentage points when calculated at the mean), although still only weakly significant.

Finally, in columns 4 and 6, our specification is augmented with a variable that isolates workplaces that completely deunionized over the period under examination. As noted in Blanchflower et al. (1991), use of OLS ignores the possibility of simultaneity bias—causation running not only from unions to employment growth but also in the reverse direction. Like Blanchflower et al., we believe that this is unlikely to be a serious problem given that the decline in union density in Australia clearly predates declines in employment. Australian Bureau of Statistics data, for example, indicate that the number of employees who were trade union members rose by less than 4 percent between 1982 and 1990, whereas the number of employees over this period rose by 27 percent.[4] Indeed, research undertaken in parallel to that reported here could find no significant role for either workplace size or changes in workplace size in explaining changes in union density in these data (Wooden, 1999). Nevertheless, there are good reasons to expect that in the small number

[4] For an overview of the decline in union density in Australia and a review of relevant literature, see Griffin and Svensen (1996).

of cases (n = 33) where the workforce had, subsequent to 1989, totally deunionized (i.e., there were no longer any union members at the workplace), the change in unionization levels may well have been a response to falling levels of employment. While this is not necessarily a problem for our analysis, given that the key explanatory variable is measured as a level rather than as a rate of change, the fact that these deunionized workplaces typically had relatively low rates of union density (mean unweighted density was 41.9 percent) suggests either omitting these cases or holding their effects constant through inclusion of a dummy variable.

As it transpires, the treatment of these deunionized workplaces has a major bearing on the results, at least within the private-sector workplaces. Once deunionized workplaces were controlled for, the size (in absolute terms) and significance of the coefficient on union density in the private sector increased markedly, with the estimated coefficient again suggesting an adverse union employment effect of almost 2½ percent per annum. Comparable results for the public sector are reported in column 6, and again, there is no evidence for any adverse union employment effect in this sector.

A number of different variants of the column 4 specification were estimated, and in the main, it proved difficult to improve on the specification reported. First, union density was replaced with a unionization variable measured at the industry, rather than workplace, level. The inclusion of this variable would be justified if the main vehicle for unions obtaining higher wages for their members was through the system of industry-based awards, and hence any union wage effects would be passed on to all workers in the industry. This variable did not exhibit any statistically significant association with employment growth.

Second, a number of more direct measures of union activity within the workplace (such as the number of union delegates per employee, the presence of union delegates who spend the majority of their work time on union business, the presence of joint consultative committee arrangements, and whether workers would rely on union representations to management when concerned about work practices) were experimented with. Rarely did the measures examined achieve any statistical significance, let alone outperform union density.

Third, interaction effects between union density and the main control variables were tested for. Weakly significant associations were reported for interactions with both employment 89 and change in autonomy, with the positive coefficients suggesting that the negative effects of union membership on employment are somewhat less severe in both large workplaces and workplaces where the shift toward greater managerial autonomy has been most pronounced. In addition, a significant negative interaction between union density and one of the demand variables—demand6 (identifying workplaces where demand was contracting in both 1989 and 1995)—also was found. This interaction is not unexpected and presumably reflects the ability of unions to restrict layoffs [a result that is consistent with other findings reported by Miller and Mulvey (1991)]. The estimated size of all these effects, however, was extremely small and hence can be ignored safely.

Fourth, the issue of whether or not the presence or absence of arrangements supporting closed union shops exerted any additional effect on employment growth was examined. Unfortunately, the presence of closed shops is not measured directly in either the 1989–1990 AWIRS or in the panel survey. Nevertheless, a dummy variable indicating 100 percent unionization among nonmanagerial employees failed to achieve statistically significant levels.

Fifth, the robustness of the results to outliers was tested for. Specifically, all cases where the value of the dependent variable was more than 2 standard deviations from the sample mean were omitted. This had very little effect on the results. Most important, the size and significance of the coefficient on the union density variable in the private-sector equation was little changed ($\beta = -0.290$; SE = 0.083).

Finally, the sensitivity of the results to the specification of the dependent variable in terms of rates of change rather than levels was examined. Specifically, the dependent variable used here was replaced by log employment. The overall pattern of results was little changed, although the size of the estimated effect of union density was slightly smaller—2.2 percent per annum within the private sector. A specification in terms of the level of employment, however, is less attractive because it restricts workplaces of different sizes to growing at the same rate.

Conclusions

The results presented in this article confirm conclusions reached in North American and U.K. research—unions slow job growth. This adverse effect of trade unions on employment growth in Australia, however, appears to be confined to the private sector. No evidence of any impact of unions on employment growth in the public sector could be found. Within the private sector, the union employment effect is quite large, with workplaces with average levels of union density in 1989–1990 estimated to have experienced rates of employment growth that are close to 2.5 percentage points per annum less than lowly unionized workplaces. This estimate is consistent with North American and U.K. research, where the union impact on employment has been estimated to lie in the range of 2 to 4 percent per annum.

Less clear is the mechanism through which unions lower private-sector employment growth in Australia. Three possible explanations lie in the effect unions have on (1) wages, (2) the cost of adjustments to workforce size, and (3) sales and output growth. It is argued here that it is the latter route that is the most likely explanation of the negative union employment effect identified in this article. This said, it is admitted that this conclusion is reached not on the basis of any strong supporting evidence in its favor, but rather, because of evidence suggesting that the alternative explanations are unlikely to be of large importance.

Turning first to the role of union wage effects, it was hypothesized earlier that while the awards system historically has acted to constrain the emergence of a union wage differential in Australia, such effects may have diminished during the period covered by this study as a result of the spread of enterprise-level collective bargaining. Recent empirical evidence, however, is not consistent with this hypothesis. Miller and Mulvey (1996), for example, in what is arguably the most well controlled study of union wage effects undertaken in Australia to date, reported that after controlling for interindustry differences, the union wage differential in 1993 data was insignificant for women and very small for men (around 2.6 percent).[5] Such conclusions receive further confirmation from an analysis of individual wage data collected from employees as part of the same data collection employed in this study—the 1995 AWIRS (Wooden, 1998).[6] These findings thus suggest that it is difficult to attribute much of the lower rate of employment growth in unionized firms over the period examined to differential rates of wage growth.

The second hypothesized route—that unions reduce long-run employment growth by reducing the ability of firms to lay off workers—is also contentious, as evidenced by the recent debate in Australia following the introduction of review processes for unfair dismissals in the federal jurisdiction as part of the Industrial Relations Reform Act 1993.[7] While it is clear that layoffs are less common where union density is high (see Miller

[5] In a more recent paper, Miller et al. (1997) report evidence of the existence of a union wage premium in Australia, but only in lowly unionized industries.

[6] Freeman (1994) also reported insignificant union wage effects for Australian workers in his cross-country study of union-nonunion wage differentials.

[7] Many of the provisions introduced in this act were either removed or significantly watered down as part of the Workplace Relations Act of 1996, which became law on December 31, 1996 (see MacDermott, 1997).

and Mulvey, 1991), it does not necessarily follow that this will translate into lower rates of employment growth, since reduction in hires should be, at least in part, compensated for by reductions in firings. The evidence presented in this analysis, while hardly conclusive, provides little support for the hypothesis that the dampening effect on hiring outweighs its effect on firing. For example, if union-imposed restrictions on the ability to lay off workers have an impact on the employment decisions of firms, then we might expect the impact of demand growth to be weaker in highly unionized workplaces than in lowly unionized workplaces. However, as noted earlier, interaction terms between union density and the high growth demand variables (e.g., demand1) were statistically insignificant when added to our preferred specification.

This thus leaves the impact of unions on sales and output as the only remaining candidate. Unfortunately, unlike in the United States, the impact of unions on output measures has not been the subject of extensive research in Australia. Moreover, what research evidence is available is mixed. Both Crockett et al. (1992) and Drago and Wooden (1992), for example, have reported evidence of a (weak) negative union impact on productivity in Australia using workplace-level data. Phipps and Sheen (1994), on the other hand, analyzed aggregate time-series data pooled across industries and found evidence that high levels of unionization have been associated with relatively rapid rates of total factor productivity growth.

Overall, it is clear that far more research is required before any firm conclusions can be reached about the relative importance of the various channels through which unions impede employment growth.

Another puzzle is the apparent discontinuity in the relationship between union density and employment growth at zero density rates. Specifically, rates of employment growth were found to be 15 percentage points less in nonunion workplaces than in workplaces with positive but very low rates of unionization. This result is very different from overseas studies. One explanation lies in the unique processes of formalized industrial relations in Australia. The presence of trade unions at a workplace provides scope for these workplaces to be "roped into" more general industrial relations objectives of the trade union movement. This "union threat" is likely to increase with union density. As a result, managers of workplaces with low levels of unionization may be encouraged to adopt improved human relations practices that dampen the incentive for other employees to join a trade union. Workplaces without a trade union presence, on the other hand, may perceive the chance of their workforce becoming unionized as more remote and hence fail to adopt strategies that might have enhanced employment growth at the workplace. Such explanations, however, are highly conjectural.

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