Union wage effects in Australia in a period of declining union power: The role of endowments and returns to endowments

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Abstract

This study explores the union-non-union wage differentials in Australia, using a quantile regression model and simulation-based counterfactual decomposition. We find that wages for unionists are higher, and more equally distributed, compared to non-unionists. The decomposition analysis reveals that the main reason for a positive union-non-union wage differential is the possession of better labour market endowments by unionists compared with non-unionists. We find that union wages are more equally distributed because endowments of key employment characteristics are more homogenously distributed among unionists. A corollary of this is that differences in the returns to endowments, the 'pure' union-non-union wage differentials, are estimated to be small, approximately 0 to 4 per cent for males and 0 to 2 per cent for females.

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Introduction

Despite a significant body of research on union-non-union wage differentials in Australia (see Cai and Waddoups 2011 and Nahm *et al*. 2017 for extensive recent literature reviews) two aspects of the union-non-union wage differential remain under-researched in Australia. The first relates to how the union-non-union wage differential varies along the wage distribution. The second relates to what factors drive union-non-union wage differentials. The only paper to tackle both aspects in the Australia context is Cai and Liu (2008). While not using a quantile regression approach, Cai and Waddoups (2011) do shed some light on the first aspect by estimating whether unions are better at raising wages for low skill, compared with high skill workers in Australia.

In terms of the first aspect, it has been well known since the work of Freeman and Medoff (1984) that unions tend to compress wages among union members. This could result from the pursuit by unions of standard rate policies that attach wages to jobs rather than to individuals. It could also result from union concerns to tackle wage discrimination (Bryson 2007). Wage bargaining models show that wage compression can also emerge from the preferences of risk averse workers who are motivated to seek insurance in the face of uncertainty. Cai and Liu (2008) find evidence to support wage compression among male union members, but not females.

Cai and Liu (2008) investigated the second aspect concerning what drives union-non-union wage differentials in Australia using the Blinder-Oaxaca technique to decompose the union-non-union wage differential into two parts. The first part is due to differences in endowments of personal and workplace characteristics, and the second part is due to differences in the returns to these endowments. They found that for males 70 per cent, and for females 30 per cent, of the observed differences between union and non-union wages could be explained by different returns to endowments, or to use the language of Freeman and Medoff, by the 'monopoly face' of unions.

The current study seeks to address both these underresearched aspects of the union-non-union wage differential. Specifically, the paper makes three contributions to the literature for Australia. First, by using panel data, fixed effects models and quantile regression techniques to analyse nine waves of HILDA data, we provide a more rigorous examination of union-non-union wage differentials along the wage distribution than those presented for Australia by Cai and Liu (2008) and Cai and Waddoups (2011). Second, the decomposition of the union-non-union wage differential in the current study controls for individual unobserved heterogeneity and also uses an econometrically superior estimation technique to Cai and Liu (2008) when simulating the marginal distribution of log wage. Third, the current study uses data covering the period 2009-17 and thus takes in a period when union density and associated bargaining power continued to decline. Specifically, Cai and Waddoups (2011) used data from 2001-2006. Cai and Liu (2008) used data up until 2004 when union density was 23 per cent whereas by 2017 it had fallen to 14 per cent, representing a decline of 40 per cent.

The paper has the following structure. The next section reviews the relevant literature and puts the study into its institutional context. Section 3 outlines the methodology employed. Section 4 discusses our data and Section 5 discusses our results. We find that unions raise the level of wages on average for union workers. They also tend to compress the spread of wages compared to non-union counterparts. We also find, in contrast to Cai and Lui (2008) that the differences in wage outcomes for union workers are overwhelmingly driven by differences in endowments, rather than by returns to endowments. In Section 6 we draw conclusions and outline some directions for future research.

Literature review and institutional context

The international literature identifies structural change, and increasing product market competition as a result of globalisation, as having reduced both the ability of unions to act on the behalf of members, as well as the incentive for workers to become union members (Gilfillan and McGann 2018, Schnabel 2013). Indeed two key factors determining the ability of unions to raise wages are the strength of union membership, and the legislative framework that either supports or hinders their ability to exercise their strength when negotiating with employers. Both these factors have moved against unions in Australia in recent decades. Union density is an important measure of union strength and in advanced economies it has been typically falling since the 1980s, for example average union density among OECD countries has fallen from 30 per cent in 1985 to 17 per cent in 2017 (OECD 2017). However, the decline in union density in Australia has been even more dramatic over the same period, falling from 46 per cent to 15 per cent (OECD 2020).

In addition, from the early 1990s in Australia there have been a number of important legislative changes that have reduced the bargaining power of trade unions and their capacity to be an effective form of collective voice. Throughout most of the twentieth century, industrial relations in Australia was regulated by the *Commonwealth Conciliation and Arbitration Act 1904*. Wage increases won by unionists under this centralised system were expressed in awards that would flow-on to all workers covered by the award, including non-unionists, thereby making it less likely that a union-non-union wage differential would be observed. Nevertheless research from this period found union-non-union wage differentials of 9-15 per cent (Miller and Mulvey, 1993). The shift towards enterprise-based bargaining occurred first under the *Prices and Incomes Accord 1991*. Under this system wage increases won by unionists at one workplace would flow-on to non-unionists at the same workplace, but they would not necessarily flow-on to non-unionists at other workplaces, thereby increasing the likelihood that a union-non-union wage differential might be observed (Wooden 2001).[1] While awards remained important in the decentralised system, they played a smaller 'safety net' role that reduced their ability to link wages and conditions between the union and non-union sectors. This change could be expected to increase the union-non-union wage differential (Cai and Waddoups, 2011). That said, the shift to enterprise bargaining was also accompanied by the creation of new methods for setting wages that did not involve unions. These included Australian Workplace Agreements (an individual contract), as well as non-union collective agreements. These would be expected to lower the bargaining power of unions and reduce the union-non-union wage differential.

In addition to these changes to wage setting institutions, the reforms that began in the early 1990s introduced a range of other changes that directly reduced union power. The *Industrial Relations Reform Act 1993* introduced a provision permitting unions to undertake legal or protected forms of strike action during collective bargaining to further their claims. However, from the *Workplace Relations Act 1996* through to the Fair Work Act 2009 the ability to undertake such protected industrial action has become increasingly circumscribed by technical procedures, third party damage protections and onerous fines. The impact of these changes on union power is revealed by the collapse in working days lost per year per 1,000 employees due to industrial action that fell from 94.9 days under the *Industrial Relations Reform Act 1993* to 14.9 under the *Fair Work Act 2009* (Bornstein 2018, Isaac 2018). Isaac (2018) has catalogued other damaging changes for unions over this period including, the *Workplace Relations Act 1996* that outlawed closed shop unionism and other forms of union preference, as well as imposing significant restrictions on the right of entry of unions to workplaces.

These changes have prompted new research on union-non-union wage differentials in Australia. Wooden (2001) was the first to examine the idea that the absence of automatic flow-on under enterprise bargaining might facilitate the identification of a union-non-union wage differential. Using matched employer-employee data from the Australian Workplace Industrial Relations Survey 1995, he found a union-non-union wage differential of 10 per cent between individuals from different workplaces with different degrees of union density and activity (Wooden, 2001). Waddoups (2005) also investigated how the changes in union density and the industrial relations environment impacted the union-non-union wage differential in Australia, using cross-section data from the Survey of Education and Training for 1993, 1997 and 2001. Waddoups found a widening of the union-non-union wage differential over time, particularly for workers in industries with high union density. This widening was interpreted as being due to the reduced likelihood of flow-on under decentralised wage setting. It should be noted that the union-non-union wage differential was small at around 5 per cent for males in 2001. Waddoups (2008), using cross-section data from the Survey of Education and Training, and Cai and

1 A 'threat effect' implies that a flow-on could arise even under enterprise bargaining. This may occur if non-union employers pass on wage increases to non-union employees in order to prevent the unionisation of their workforce.

Liu (2008) who pooled four waves from HILDA (2001-4), found union-non-union wage differential of around 11 per cent for males and 5 per cent for females.

None of these papers mentioned above controlled for unobserved individual heterogeneity and other forms of endogeneity, something that has been shown to overestimate union-non-union wage differentials for males in Australia (Cai and Waddoups, 2011), the US (Mellow, 1981; Freeman, 1984; Hirsch and Schumacher, 1998) and the UK (Hildreth, 1999; Swaffield, 2001). In the Australian context, Cai and Waddoups (2011) used HILDA for the period 2001-6 and found that, once they controlled for unobserved individual heterogeneity, male union-non-union wage differentials fell from 9 per cent to 5 per cent and female effects fell from 6 per cent to 2 per cent. A more recent study by Nahm *et al*. (2017), using the HILDA (2001-13) estimated small negative union-non-union wage differentials for males (-6.2 per cent) and females (-6.8 per cent).[2]

For Australia Cai and Waddoups (2011) found statistically significant union-non-union wage differentials for low skill workers, with no effect found for workers at the high skill end. Cai and Liu (2008) investigated how union-non-union wage differentials vary along the wage distribution using a quantile regression model. They found that for males union-non-union wage differentials are greatest at the 10th percentile and decline along the wage distribution until the 90th percentile when they become negative. By contrast, for females union-non-union wage differentials exhibit more stability across the wage distribution.

Using a Blinder-Oxacca decomposition Cai and Liu (2008) found that for males, most of the observed difference in wages between union and non-union workers, 70 per cent or more, was due to differences in returns, while for females the difference, 70 per cent or more, were driven by differences in endowments. Therefore, for males the overwhelming majority and for females a significant minority of the union-non-union wage differential reflects the power of unions to influence the price of labour. These results should be regarded with caution, because they did not control for unobserved individual heterogeneity.

2 Note that this result is not directly comparable with the result reported in the present paper because of the difference in the treatment of individual heterogeneity. In the present paper, the difference in average heterogeneity between unionised and non-unionised workers is treated as an endowment effect, while it is decomposed into an endowment effect and a coefficient effect based on the coefficient estimates for the group-mean averages (i.e. the Mundlak devices, Mundlak 1978) in Nahm et al. (2017). This difference in the approaches is mandated by restrictions imposed by the methodology.

Methodology

Union effects on wages can vary considerably along the wage distribution. Existing studies typically analyse regression quantiles to examine varying union-non-union wage differentials (e.g. Cai and Liu, 2008, and O'Leary *et al.*, 2004). These studies, however, analyse cross-sectional or pooled data without properly controlling for heterogeneity of individual workers. It appears that union effect is typically inflated when heterogeneity is uncontrolled for; see for example Graham *et al.* (2018) and Nahm *et al.* (2017). The present paper estimates quantile regression coefficients, explicitly controlling for heterogeneity of individual workers. It then uses a counterfactual decomposition method to analyse the contribution of individual covariates towards union-non-union wage differentials, as well as their contributions as a group, at various points in the marginal (i.e., unconditional) wage distribution.

Quantile regression

Consider the model given by:

$$w_{it} = X_{it}'\beta(u_{it}) + \alpha_i \tag{1}$$

where w is log wage, X_{it} is the vector of observable covariates for individual i in time t, α_i represents unobservable individual heterogeneity for individual i, and $X_{it}'\beta(u_{it})$ is strictly increasing in u_{it} where $\beta(u_{it})$ is the coefficient vector that depends on $u_{it}\in(0,1)$. In general, α_i and u_{it} can be arbitrarily correlated, in which case identification of the model is problematic. To facilitate identification of the model, we draw on Canay (2011) and assume that α_i and u_{it} are independent from each other so that α_i is a location shifter, implying that the τ^{th} quantile, $\beta(\tau)$, is not affected by α_i:

$$P(w_{it} \leq X_{it}'\beta(\tau) + \alpha_i \,|\, X_{it}) = P(w_{it} - \alpha_i \leq X_{it}'\beta(\tau) \,|\, X_{it}) = \tau\in(0,1). \tag{2}$$

As Canay explains, this allows us to obtain consistent estimates of $\beta(\tau)$ using a simple two-step method under some regularity conditions. In the first step, the fixed-effects model is estimated and a new variable w^* is generated as $w_{it}^* = w_{it} - \hat{\alpha}_i$ where $\hat{\alpha}_i$ are the fixed-effects estimates. In the second step, the regression quantiles, $\hat{\beta}(\tau)$, are obtained by solving the following minimisation problem:

$$\min_{\beta}\sum_{i}\sum_{t}\rho_\tau(w_{it}^* - X_{it}'\beta) \tag{3}$$

where $\rho_\tau(v) = v\times\tau - v\times I(v<0)$ with $I(.)$ denoting the indicator function.[3]

3 Mahuteu *et al.* (2017) recently used the same method to examine public-private sector wage differentials.

The set of the covariates for our model consists of conventional determinants of wage, including: work experience and its square, occupation tenure and its square, job tenure and its square, a public-sector dummy, a dummy for marital status, a dummy for having one or more children aged between 0 and 4, a dummy for children aged between 5 and 9, two dummies for country of birth, four dummies for education level, three dummies for employment contract type, seven dummies for occupation, nine dummies for industry, five dummies for organisation size, twelve dummies for regions, eight dummies for waves, and a constant term. These are standard explanatory variables that are widely used for the analysis of wage determination; see for example Cai and Liu (2008), Cai and Waddoups (2011), Dobbie *et al*. (2014), Nahm *et al*. (2017) and Wooden (2001). The models for unionised workers and non-unionised workers are simultaneously estimated by adding the interactions of these variables with a union dummy variable. The issue of self-selection into union membership is not explicitly considered in our analysis, and it is assumed that selection into union membership is exogenous to the wage equations.[4] However, even if the selection process is endogenous, controlling for individual heterogeneity must have mitigated potential problems of this omission to the extent that the self-selection is a function of time-invariant characteristics of individual workers.

Counterfactual decomposition

Regression quantiles provide valuable information about how the covariates affect log wage across its conditional distribution and how these effects are different between union and non-union. However, the regression quantiles themselves do not show how the wage differential between the two groups can be decomposed into the part that is due to the difference in the characteristics (*endowment effect* hereafter) and the part due to the difference in the returns to these characteristics (*coefficient effect* hereafter). Furthermore, regression quantiles represent the effects of covariates along the *conditional* distribution of log wage, and not along the *marginal* distribution. For instance, if the 90th regression

4 This is a practical decision we made in order to focus on the issue of varying wage differentials over different wage levels without making the analysis too complicated. The issue of self-selection into union membership is explicitly considered in Nahm *et al*. (2017) where they analyse mean wage differentials, as opposed to quantiles, by estimating a simultaneous-equations model that consists of two wage equations and an equation representing self-selection into union membership. In the context of this model, the present decision not to explicitly model the self-selection process could be justified if the random error term of the selection model were uncorrelated with the random errors of the wage equations, in which case the selection process is exogenous. If the selection process were exogenous, inclusion of an explanatory variable that is endogenously correlated with self-selection in the wage equations would not cause inconsistency in the estimates. For example, inclusion of occupation dummies in the present study would not be problematic even if union membership were a necessary condition for some occupations.

quantile for education is 0.05 (assuming that the variable is measured in years), it is implied that an additional year of education leads to a 5 per cent increase in wage for a worker whose wage is at the 90th percentile in the distribution of wages for workers with the same levels of the characteristics, including education. One would be more interested in these effects along a wage distribution that is not conditional on a specific set of covariates so that the effects can be analysed for workers with all levels of the characteristics. This would be achieved if the conditional distribution, which is conditional on a specific set of covariates, could be integrated over the whole domains of the covariates. Machado and Mata (2005) introduce a method of empirical integration through resampling. We draw on their method to decompose the overall wage differentials between union and non-union and to measure the endowment effects of individual covariates. We then extend their idea and attempt to measure the coefficient effects of individual covariates as well as their endowment effects. Machado and Mata's (2005) method and our extension to the methodology is explained in the appendix.

Data

The data used for this study are extracted from the Household, Income and Labour Dynamics in Australia (HILDA) database. The HILDA is a household-based longitudinal study that has been collecting data annually from year 2001. The present paper analyses waves 9 (2009) through 17 (2017). Each wave includes a panel of six to eight thousand employees who were directly interviewed and aged between 15 and 65 in the year of interview, with a total of 67,572 wave-person observations. The number is evenly divided between males (33,744) and females (33,828), out of which 21.6 per cent and 24.0 per cent are union members, respectively; see Table 1.

It is worth noting that since the first wave of HILDA data in 2001 the estimates of trade union membership are typically 4-6 percentage points higher than those obtained from the Australian Bureau of Statistics (ABS). Wooden (2009) has argued that there are two reasons for this (i) HILDA includes membership of professional associations in its trade union membership question and (ii) HILDA only includes responses from people interviewed and excludes proxy responses from interviewees about other members of the household which the ABS includes. In terms of (i) from 2009 onwards a separate question has been included in the survey that asks only about trade union membership and that is the question used in this study. Furthermore, in terms of (ii) Wooden (2009) pointed out that this is expected to reduce the ABS estimate by at most 0.5 percentage points. Despite the change in question HILDA estimates continue to be 6 percentage points above the ABS estimates for any given year. Importantly, however, the HILDA estimate of trade union density has undergone the same dramatic decline since the period covered by Cai and Liu (2008). Specifically, the HILDA estimate fell from 30.5 per cent in 2004 to 21.5 per cent in 2017 a decline of 30 per cent.

Table 1 indicates that on average, union members are paid higher wages than those who are not union members ($33.23 vs. $30.08 per hour), where the gap is slightly higher for female workers ($31.41 vs. $27.68) than male workers ($35.26 vs. $32.40). The standard deviations and the Gini indices in Table 1 show that wages are more equally distributed among union members than non-unionised workers. Wages are more equally distributed among female workers than male workers, and the gap in inequality between union and non-union is larger for men than for women. The ratios of 25th percentile to 10th percentile imply that, for those who are paid relatively low wages, there is little difference in the degree of inequality between the gender-based groups or the union membership-based groups. For higher wage earners, however, the wage compression effect of unions is evident as the ratio of 90th percentile to 75th percentile for unionised workers remains largely unchanged while the ratio increases significantly for non-unionised workers.

Table 1: Summary statistics of wage rate by gender and union membership

	Male			Female			All		
	Union	Non-union	All Male	Union	Non-union	All Female	Union	Non-union	All
Average Wage	35.26	32.40	33.02	31.41	27.68	28.58	33.23	30.08	30.80
S.D.	16.42	18.99	18.51	13.14	14.70	14.43	14.91	17.18	16.74
Gini Index	0.223	0.278	0.268	0.207	0.240	0.235	0.217	0.264	0.255
P25/P10	1.224	1.237	1.238	1.225	1.212	1.204	1.241	1.221	1.218
P90/P75	1.240	1.412	1.360	1.209	1.348	1.306	1.224	1.390	1.334
Number of Obs.	7,299 (21.6%)	26,445 (78.4%)	33,744	8,126 (24%)	25,702 (76%)	33,828	15,425 (22.8%)	52,147 (77.2%)	67,572

Notes: Wages are real hourly wages in 2011/12 Australian dollars.

Table 2: Sample means of the covariates by union membership and gender

	Male		Female			Male		Female	
Variable	Union	Non-union	Union	Non-union	Variable	Union	Non-union	Union	Non-union
Work experience	24.8**	18.9	22.1**	17.2	*Industry*				
Occupation tenure	13.3**	8.5	13.1**	7.5	Primary	0.00	0.03**	0.00	0.01**
Job tenure	11.6**	5.7	10.9**	5.3	Blue collar	0.34**	0.30	0.03	0.08**
Public sector	0.45**	0.15	0.61**	0.23	Wholesale/trans.	0.14**	0.12	0.02	0.05**
Married/de facto	0.77**	0.70	0.72**	0.67	Retail	0.05	0.08**	0.08	0.11**
Children 0–4	0.15	0.19**	0.11	0.14**	Hospitality	0.01	0.04**	0.01	0.07**
Children 5–9	0.15	0.15	0.14	0.15	Business services	0.05	0.10**	0.05	0.13**
Country of birth					Government	0.16**	0.08	0.08**	0.07
Australia	0.82**	0.80	0.83**	0.80	Education	0.12**	0.04	0.33**	0.12
Other English	0.08	0.10**	0.08	0.08	Health/community	0.08**	0.05	0.37**	0.23
Non-English	0.09	0.10**	0.09	0.12**	Recreation/other	0.05	0.15**	0.03	0.13**
Education					*Firm size*				
Tertiary	0.26	0.28**	0.50**	0.33	< 20	0.05	0.25**	0.03	0.23**
Adv. Dip. & Dip.	0.10**	0.09	0.11	0.11	20–99	0.08	0.18**	0.08	0.17**
Cert. III or IV	0.36**	0.28	0.16	0.21**	100–499	0.16	0.17**	0.14	0.17**
Year 12	0.13	0.17**	0.10	0.17**	500–999	0.09**	0.07	0.07	0.08*
Year 11 or below	0.15	0.17**	0.12	0.18**	1,000–4,999	0.19**	0.12	0.11	0.12
Employment contract					≥ 5,000	0.43**	0.20	0.57**	0.23
Fixed-term	0.07	0.10**	0.11	0.11	*Region*				
Casual	0.06	0.16**	0.07	0.22**	Sydney	0.14	0.19**	0.16	0.17*
Permanent/ongoing	0.87**	0.74	0.82**	0.67	Other NSW	0.16**	0.10	0.15**	0.11
Other	0.00	0.00*	0.00	0.00	Melbourne	0.19	0.18	0.18	0.20**
Occupation					Other VIC	0.08**	0.06	0.08**	0.07
Manager	0.07	0.18**	0.06	0.10**	Brisbane	0.09	0.10**	0.11*	0.10
Professional	0.22*	0.21	0.50**	0.24	Other QLD	0.12	0.11	0.10	0.11*
Technician/Trades	0.22	0.21	0.02	0.05**	Adelaide	0.07**	0.06	0.06	0.07
Com. & pers. Service	0.11**	0.06	0.17	0.16	Other SA	0.02	0.03**	0.02	0.02**
Clerical & admin.	0.07	0.08**	0.13	0.27**	Perth	0.05	0.07**	0.06	0.07**
Sales	0.02	0.06**	0.07	0.11**	Other WA	0.03**	0.02	0.01	0.02**
Machinery operator	0.17**	0.10	0.01	0.01**	TAS	0.04**	0.03	0.05**	0.03
Labourer	0.11*	0.10	0.05	0.06*	NT	0.01	0.01**	0.01	0.01**
					ACT	0.02	0.03	0.02	0.02**

Notes: Work experience, occupational tenure and job tenure measured in years. * and ** implies significantly larger than the other at 5 % and 1 %, respectively, in the comparison between union and non-union.

Table 2 reports the sample means of the variables used in the model. As normally observed in the literature, unionised workers in the sample are more experienced and have longer occupation and job tenures. Further, compared with non-unionised workers, unionised workers are more likely to be born in Australia; have post-school education; be a permanent employee; work for a large company; and be employed in a blue-collar industry, or government, education, or health and community sector. These observations reveal very little beyond what is already well known in the literature. What is less known, however, is detailed information about the drivers of the wage differentials, especially at different points in the wage distribution. The empirical analysis in the next section addresses this issue.

Empirical Results

This analysis has generated a large amount of output. For brevity we report and discuss only the results of the decomposition analysis. In addition, when we further decompose the overall results reported in Table 3 into the individual cofactors we report only a small subset of these that are especially relevant to our core analysis. The results from our quantile regression analysis and decomposition are generally sensible and consistent with expectation and previous literature. A full set of results, with a written report, are available on request.

Results from overall decomposition

In Table 3 the combined difference (as defined by equation A2 in the appendix) from the overall decomposition for males indicates that wages for unionised workers differ from those for non-unionised at the 10th, 25th, 50th, 75th and 90th percentiles by 17.5 per cent, 19.6 per cent, 18.3 per cent, 9.7 per cent and -5.4 per cent. The negative differential at the 90th percentile is not statistically significant.[5] The corresponding numbers for females are 14.5 per cent, 15.1 per cent, 18.8 per cent, 15.7 per cent and 6.4 per cent.[6] These findings are consistent with the idea that unions tend to raise wages, and to do so more for workers at the lower end of the wage distribution than at the top, and by so doing help to compress union wages.

5 Fransden (2012) also finds evidence for a significant increase in wage effect at the low end, and a negative wage effect at the very top, of the distribution for workers in the USA estimating quantile treatment effects within a regression discontinuity design.

6 These results are numerically similar those in Schmitt (2008) who finds (for pooled males and females) 20.6 per cent at the 10th percentile and 6.1 per cent at the 90th percentile, using USA data for 2003-7.

Table 3: Decomposition of log-wage difference between union and non-union

	P10	P25	P50	P75	P90	P25/P10	P90/P75	SD	Gini
Males									
Combined difference	**0.175**	**0.196**	**0.183**	**0.097**	-0.054	0.002	**-0.043**	**-0.093**	**-0.018**
Endowment effect	**0.179**	**0.165**	**0.150**	**0.062**	**-0.102**	**-0.010**	**-0.045**	**-0.105**	**-0.019**
Coefficient effect	-0.004	**0.031**	**0.033**	**0.035**	**0.048**	0.013	0.003	**0.012**	**0.001**
Females									
Combined difference	**0.145**	**0.151**	**0.188**	**0.157**	**0.064**	-0.002	**-0.029**	**-0.039**	**-0.009**
Endowment effect	**0.126**	**0.136**	**0.167**	**0.140**	**0.071**	0.000	**-0.022**	**-0.031**	**-0.007**
Coefficient effect	**0.019**	**0.014**	**0.021**	0.016	-0.007	-0.002	-0.007	**-0.008**	**-0.002**

Notes:
For each characteristic function, g(.), the differences are g(union) minus g(non-union).
Bold figures are significant at 5 %, implying that 95 % confidence interval does not include zero. The confidence intervals are estimated using the bootstrap method with 200 repeated samples.

The closest study to the current paper for Australia is Cai and Liu (2008). The estimates reported as the "total wage gap in simulated data" and the "gap due to difference in returns", in Table 2 of their paper are comparable with our estimates for the combined difference and the coefficient effect, respectively. They found that for males the total wage gap in simulated data between union and non-union wages decrease from 20 per cent, 16.7 per cent, 13.9 per cent, 7.1 per cent to -2 per cent across the 10th, 25th, 50th, 75th, and 90th percentiles. For females, the corresponding union-non-union wage differential was estimated to be 17 per cent, 14.2 per cent, 18.5 per cent, 17.7 per cent and 10 per cent. Despite differences in methodology and data period employed, these results are numerically similar to those in this paper, and the qualitative picture they paint about the effect of unions on inequality across the distribution is the same. The fact that Cai and Liu's results are systematically larger than ours is possibly explained by their not controlling for unobserved heterogeneity. It might also reflect the fact that their sample data came from a period where union power was greater for the reasons discussed previously.

Cai and Waddoups (2011) also explored whether unions are more effective at raising wages at the top or bottom of the wage distribution using HILDA from the years 2001-6. While they did not estimate quantile regressions, part of their analysis involved estimating separate fixed effects regressions for high and low skill workers, where skill was defined by education and predicted wages. They found statistically significant union-non-union wage differentials only for low skill workers. For low skill males, wages were 8 per cent higher using predicted wages, and 5 per cent higher using education levels, than for non-unionists. For females they found a statistically significant union effect of 4 per cent at the second quartile using predicted wages. No statistically significant union effects were found at the higher skill end for male or female unionists.

Table 3 contains other evidence for wage compression. The numbers in the columns headed P25/P10 (ratio of 25th to 10th percentile), P90/P75 (ratio of 90th to 75th percentile), SD (Standard Deviation) and Gini (Gini Index) are the differences between the union and non-union values of these statistics. As such they are measures of 'relative inequality' between the two groups, with a negative sign suggesting less inequality among union compared to non-union workers. The negative sign in front of the P90/P75 ratios in Table 3, in the combined difference rows for both males and females, indicates lower inequality at the top of the wage distribution for unionists compared to non-unionists. The same comments apply to the numbers recorded under the headings, SD and Gini, in the same row for both genders.

Comparing union effects on wages across countries is challenging as the nature of unions and the institutional environments in which they operate are quite different (Bryson, 2014). Moreover, most of the results with which ours might be compared are from regressions on the conditional distribution which often do not control for unobserved heterogeneity. That said, a significant international literature finds support for the idea that unions both raise and compress wages. See for example: Schultz and Mwabu (1998) using 1993 data for South African workers; Fritzenberger, Kohn and Lembcke (2013) for German workers using matched employer-employee data for 2001; Blunch and Verner (2001) for manufacturing workers in Ghana; O'Leary, Murphy and Blackaby (2004), Marquilef- Bachler *et al*. (2009) and Hildreth (1999) for the UK (Marquilef-Bachelor *et al*. only find evidence for this for public sector males, and only once the decision to join a union is endogenised); Falaris (2004) for Panama; Fransden (2012) and Schmitt (2008) for the USA.

Very few of these studies attempt to decompose the total differences between union and non-union workers into a part that is due to differences in endowments, and a part due to different returns to those endowments. As noted in the introduction to this paper, Cai and Liu (2008) conducted such a decomposition and found that 70 per cent of the observable difference in wages between male union and non-union workers could be explained by different returns, with the rest due to differences in endowments. The corresponding figures for females were 30 per cent for returns and 70 per cent due to different endowments. A clear point of difference in our results, reported in Table 3 is that for both genders the differences in earnings between union and non-union workers is driven overwhelmingly by endowment effects, as can be seen from the endowment effect rows for males and females. These effects are statistically significant at all the reported quantiles. Table 3 indicates that the coefficient effects are also mostly statistically significant and positive, implying that unionised workers receive higher returns than non-unionised workers for their characteristics. However, the size of these effects is much smaller than the size of the endowment effects, indeed the endowment effects tend to be about six times the size of the coefficient effects on average. This finding is consistent with a previous Australian study by Nahm *et al*. (2017). In that study the decomposition of the estimates, at the mean values of the covariates, suggested that around 80 per cent of the difference between the wages of union and non-union workers was due to different endowments.

The SD and Gini measures of relative inequality in Table 3 for the combined difference for both genders are negative. This implies that wages among unionised workers are more equally distributed than those among non-unionised workers. An examination of the endowment effect and coefficient effect for both genders in Table 3 indicates that the reason for this is that union workers have endowments that are more homogeneously distributed. For male workers, the coefficient effects on these measures are positive, meaning that the reward system for unionised workers increases wage inequality relative to non-unionised workers. However, this effect is dominated by the inequality-reducing effect of more homogeneous distribution of characteristics among unionised workers than non-unionised workers, resulting in the negative values of relative inequality measures. For female workers, both endowment effect and coefficient effect are negative, but the size of the endowment effect is many multiples that of the coefficient effect.

The 'pure' union wage effects

The coefficient effects in Table 3 indicate that the top three quarters of unionised male workers enjoy an overall wage premium of 3.1 per cent to 4.8 per cent, with the size steadily increasing as one moves up through the wage distribution. This is consistent with the positive signs of the coefficient effects for SD and Gini index for male workers. In contrast the coefficient effects in Table 3 for females indicates that only the bottom half of these workers enjoy a significant wage premium, which ranges between 1.9 per cent and 2.1 per cent. These are the 'pure' union-non-union wage differentials in the sense that they do not result from differences in endowments. They result from differences in the way the market rewards those endowments between union and non-union workers. Note that they are small by comparison to the role played by different endowments in explaining the combined wage differences that we observe for both males and females in Table 3. It is again relevant to compare our results to those of Cai and Liu (2008) which for males ranged from 16 per cent at the 10th percentile to 1 per cent at the 90th percentile. For females the figure was a relatively constant 5 per cent for the bottom 75 per cent of the wage distribution, and a statistically insignificant negative 0.6 per cent at the 90th percentile. There are a number of reasons for the differences in magnitude between Cai and Liu and our paper. First, we control for unobserved heterogeneity. Second, we use a somewhat different methodology to do the decomposition. Third, we use a different, arguably more appropriate measure of individual union membership available from wave 9 onwards. Fourth, they may reflect reduced union power in our sample period.

Differences in endowments and what do unions do?

The analysis of pure union-non-union wage differentials suggests that the ability of unions to raise wages and thus earnings is quite limited in Australia. However, some of the endowment differences between unionists and non-unionists may also be the result of union activity. Specifically, Freeman and Medoff (1984) discuss how the 'collective voice' face of unions often involves unions pushing for procedurally fair arrangements for workers around such things as retrenchment, redeployment and redundancy, as well as grievance and performance matters. These arrangements are typically designed to limit the power of employers to dismiss employees at will and function to increase the employment security of unionists which generates better employment stability. Consequently, unionists compared to non-unionists will have superior amounts of three important forms of human capital: general labour market experience, occupational tenure and job tenure that have been shown to be drivers of higher lifetime earnings for workers in Australia (Dobbie *et al*. 2014), the US (Kambourov and Manovskii 2009) and the UK (Zangelidis 2008). In the current study unionists do in fact have more of these three forms of human capital compared to non-unionists. This is evident in Table 4, which reports a decomposition analysis in which the overall differentials are further decomposed into the effects of three individual experience cofactors (as defined by equation A3 in the appendix).

Table 4: Decomposition of log-wage difference between union and non-union, labour market experience cofactors

	P10	P25	P50	P75	P90	P25/P10	P90/P75	SD	Gini
Males									
Work experience									
Combined difference	**0.079**	**0.090**	**0.080**	**0.062**	**0.042**	0.001	-0.006	**-0.020**	**-0.005**
Endowment effect	**0.074**	**0.079**	**0.071**	**0.044**	**0.029**	-0.000	-0.005	**-0.016**	**-0.004**
Coefficient effect	**0.006**	**0.012**	**0.009**	**0.018**	0.013	0.002	-0.002	**-0.004**	**-0.001**
Occupation Tenure									
Combined difference	**0.033**	**0.031**	**0.029**	**0.028**	-0.001	-0.002	-0.008	-0.019	**-0.003**
Endowment effect	**0.038**	**0.044**	**0.034**	**0.038**	**0.010**	0.001	-0.008	-0.016	**-0.003**
Coefficient effect	-0.005	**-0.013**	**-0.005**	**-0.011**	**-0.011**	-0.002	0.000	**-0.003**	**-0.000**
Job Tenure									
Combined difference	**0.039**	**0.063**	**0.054**	**0.059**	0.016	0.007	-0.013	**-0.012**	**-0.003**
Endowment effect	**0.038**	**0.066**	**0.055**	**0.059**	0.019	0.008	-0.012	**-0.012**	**-0.003**
Coefficient effect	0.000	-0.003	-0.001	0.000	-0.003	-0.001	-0.001	0.000	0.000
Females									
Work experience									
Combined difference	-0.006	-0.003	-0.018	-0.012	**-0.028**	0.001	-0.004	**-0.012**	**-0.001**
Endowment effect	**0.025**	**0.034**	**0.028**	**0.045**	**0.043**	0.002	-0.001	0.003	0.000
Coefficient effect	**-0.032**	**-0.037**	**-0.046**	**-0.057**	**-0.070**	-0.001	-0.003	**-0.015**	**-0.001**
Occupation Tenure									
Combined difference	0.009	**0.025**	**0.034**	**0.047**	**0.037**	0.005	-0.003	0.014	0.002
Endowment effect	**0.023**	**0.039**	**0.041**	**0.051**	**0.041**	0.005	-0.004	0.007	0.000
Coefficient effect	**-0.014**	**-0.013**	**-0.007**	**-0.004**	-0.004	0.000	0.000	**0.008**	**0.001**
Job Tenure									
Combined difference	**0.028**	**0.052**	**0.048**	**0.047**	**0.030**	0.008	-0.005	-0.002	-0.001
Endowment effect	**0.032**	**0.060**	**0.056**	**0.053**	**0.035**	0.009	-0.006	-0.002	-0.001
Coefficient effect	**-0.005**	**-0.008**	**-0.008**	**-0.006**	**-0.004**	-0.001	0.001	0.000	**0.000**

Notes:
1. For each characteristic function, g(.), the differences are g(union) minus g(non-union).
2. Bold figures are significant at 5 %, implying that 95 % confidence interval does not include zero. The confidence intervals are estimated using the bootstrap method with 200 repeated samples.

Table 4 shows that both male and female unionists have better endowments of work experience, occupational and job tenure, than their non-union counterparts, across the entire wage distribution. The coefficient effects of work experience are positive for men, but they are negative for women, at all wage levels. By contrast, Cai and Liu (2008) report negative coefficient effects for experience for unionised males, with the opposite being the case for unionised females. O'Leary *et al*. (2004) report negative union effects for experience for both males and females in the UK. Our results imply that male unionised workers are rewarded better for longer work experience than non-unionised workers, but that the opposite is true for female unionists. The penalty experienced by female unionists increases steadily as wage rate increases.

In addition, our analysis shows that work experience is a factor that reduces inequality in the distribution of wages among unionised workers in comparison with non-unionised workers for both men and women. This is evident from the fact that the SD and Gini in the combined difference row in Table 4 are significant and negatively signed. However, our analysis shows that the main drivers for this are different between men and women. For men, work experience lowers inequality among unionised workers mainly because they have a more homogenous distribution of work experience, the SD and Gini on endowment effects are -0.016 and -0.004 respectively. For women, however, inequality in wages is lower among unionised workers because they are penalised more for being unionised at higher wages, with the coefficient effects going from -0.032 to -0.07 as we move up the female wage distribution.

For the other two experience-related characteristics, the endowment effects are significantly positive throughout the wage distribution for both men and women, except for the 90th percentile of job tenure for men where the effect is positive but insignificant. This is consistent with what was reported earlier in the paper, that unionised workers have significantly longer occupation and job tenures on average than non-unionised workers. The coefficient effects are significantly negative except for the case of job tenure for men. Both tenure variables contribute toward lowering inequality in the wage distribution among unionised males in comparison with non-unionised males, but their effects on inequality are insignificant for female workers. As was the case for work experience, the main reason for this, in the case of males, is the more homogeneous distribution of these characteristics among male unionised workers. This analysis suggests that unions can increase the earnings of their members directly by raising wages and indirectly by increasing experience related forms of human capital. It also adds additional insight into how union associated wage compression occurs.[7]

7 One objection to this line of argument is that unionists have greater amounts of experience related human capital, not because of the activity of unions, but because unionists are on average older than non-unionists. To test this we split the sample into a younger workers, aged 18-40 years and older workers aged 41-65 years. Even among younger workers unionists have significantly greater amounts of the three forms of human capital than do non-unionists. The reader may be concerned about potentially bi-directional causality between union membership and human capital. However,

One final piece of evidence to support this comes from Dobbie *et al*. (2017). That paper found that unions in Australia trade-off wages in return for better workplace training options for union members. These better workplace training options should correlate with better work experience, job and occupational tenure outcomes.

Conclusion

Our analysis indicates that unionists earn significantly more than equivalent non-unionists across the wage distribution. Our results suggest that the union-non-union wage differential tends to diminish as we move up the wage distribution and for males at the very top there may be no effect at all from union membership. These results are consistent with various approaches which suggest unions tend to compress wages. We decompose this differential into a part that relates to different endowments and a part that relates to different returns to those endowments. The latter are the pure union effects. We find them to be small and indeed smaller than typically found in the previous Australian literature. This may be partly due to our methodology which allows us to control for factors which have been shown to otherwise inflate union effects. It may also reflect the more challenging environment in which unions operate in Australia these days. The bulk of the observed union-non-union wage differential, what we call the combined difference, is due to the possession of superior endowments by unionists. Moreover, our analysis reveals that these endowments also tend to be more homogeneously distributed among unionists and this is the main reason that union wages are more compressed.

While not strictly part of the analytical core of this paper we speculate that these better endowments may result, at least in part, from the actions that unions undertake in respect to the 'collective voice'. These actions aim to create greater employment stability for unionists and hence result in them possessing more of the various forms of labour market experience that have been shown to add to lifetime earnings. Future research on union-non-union wage differentials may be better focused on exploring in more detail the role played by collective voice, as opposed to the monopoly face of unions.

under the same assumption of exogenous selection into union membership, such reverse causality in the selection model would not be problematic to obtaining consistent estimates of the coefficients of the wage equations.

Appendix: Counterfactual decomposition

This appendix describes the method by Machado and Mata (2005) and our extension to the methodology. For convenience of the description of the methodology, let X^1 be the $(n_1 \times K)$ matrix of n_1 observations on the K covariates for union members in the sample, and X^0 be the $(n_0 \times K)$ matrix of the covariates for non-unionised workers. The quantile regression coefficient vectors corresponding to these design matrices are denoted by $\beta^1(\tau)$ and $\beta^0(\tau)$, respectively, where τ has a uniform distribution between 0 and 1. The marginal distributions of log wage that are consistent with the estimated quantile coefficients are simulated as follows.

(a1) Randomly draw a large number of observations (say, m observations) from the uniform distribution between 0 and 1. Denote these m observations as $\{\tau_i\}_{i=1}^{m}$.

(a2) Estimate the quantile regression model (2)-(3) for each τ_i using the n_1 observations of unionised workers and then using the n_0 observations of non-unionised workers. Denote these 2m sets of quantile coefficient estimates as $\{\hat{\beta}^1(\tau_i)\}_{i=1}^{m}$ for union and $\{\hat{\beta}^0(\tau_i)\}_{i=1}^{m}$ for non-union.

(a3) Generate a random sample of m observations on the covariates by drawing with replacement from the rows of each of X^1 and X^0. Denote these two random samples by $\{X_i^1\}_{i=1}^{m}$ for union and $\{X_i^0\}_{i=1}^{m}$ for non-union.

(a4) The marginal distributions of w^* are then obtained by $\left\{w_i^*\left(X^j;\beta^h\right) \equiv X_i^j\hat{\beta}^h(\tau_i)\right\}_{i=1}^{m}$ where j = 0, 1 and h = 0, 1.

(a5) Individual heterogeneity ($\hat{\alpha}_i$) is a significant part of predicted log wages. Hence, it is sampled together with the other cofactors in step (a3), and it is added to w_i^* to construct the marginal distributions of log wage:

$$\left\{w_i\left(X^j,\hat{\alpha}^j;\hat{\beta}^h\right) = w_i^*\left(X^j;\hat{\beta}^h\right) + \hat{\alpha}_i^j \equiv X_i^j\hat{\beta}^h(\tau_i) + \hat{\alpha}_i^j\right\}_{i=1}^{m}. \quad \text{(A1)}$$

This allows us to simulate not only the factual union and non-union marginal distributions that are consistent with the regression quantiles, but also counterfactual distributions. For instance, the marginal distribution of log wage that would prevail in the counterfactual situation where the covariates are distributed as for non-unionised workers while the returns to those covariates are the same as those applied to unionised workers is given by $\left\{w_i\left(X^0,\hat{\alpha}^0;\hat{\beta}^1\right)\right\}_{i=1}^{m}$.

Let $g(w)$ be a characteristic function, such as a quantile or the standard deviation, of the marginal distribution of w. Further, let w^1 and w^0 denote log wages for union and non-union, respectively. Then, the differential in $g(w)$ between union and non-union is obtained as follows:

$$\begin{aligned}
&\mathrm{g}(w^{1}) - \mathrm{g}(w^{0}) \\
&= \mathrm{g}\left[w\left(X^{1}, \hat{\alpha}^{1}; \hat{\beta}^{1}\right)\right] - \mathrm{g}\left[w\left(X^{0}, \hat{\alpha}^{0}; \hat{\beta}^{0}\right)\right] + \text{residual} \\
&= \left\{\mathrm{g}\left[w\left(X^{1}, \hat{\alpha}^{1}; \hat{\beta}^{1}\right)\right] - \mathrm{g}\left[w\left(X^{0}, \hat{\alpha}^{0}; \hat{\beta}^{1}\right)\right]\right\} \\
&+ \left\{\mathrm{g}\left[w\left(X^{0}, \hat{\alpha}^{0}; \hat{\beta}^{1}\right)\right] - \mathrm{g}\left[w\left(X^{0}, \hat{\alpha}^{0}; \hat{\beta}^{0}\right)\right]\right\} + \text{residual}
\end{aligned} \tag{A2}$$

The first row represents the difference in the characteristic function between the distributions of w^{1} and w^{0} in the sample, which is referred to as the *total difference*. The difference between the first two terms in the second row represents the difference in the characteristic function between the simulated distributions of w^{1} and w^{0}, which we refer to as the *combined difference*. The third row is the difference between the simulated distribution of wage for union workers and the counterfactually simulated distribution of wage for non-unionised workers that would prevail if the returns to their characteristics were the same as those for unionised workers. Thus, this represents the *endowment effect*. Finally, the difference between the first two terms in the fourth row represents the *coefficient effect* because it is the difference between the two distributions with the same covariate distributions but different coefficients. Note that the coefficient for heterogeneity is unity, and hence it is treated as an endowment. As such, for the ease of exposition, heterogeneity is no longer denoted separately, and it is treated like any other cofactor hereafter.

In addition to the above decomposition of overall difference, the difference pertaining to each cofactor is also decomposed into the endowment effect and the coefficient effect. Let z be the cofactor of interest and x be the vector of the other cofactors, where the coefficient vector is accordingly divided into $\delta(\tau)$ for z and $\theta(\tau)$ for x. The combined difference in the characteristic function pertaining to covariate z is decomposed as follows:

$$\begin{aligned}
&\mathrm{g}\left[w\left(x^{1}, z^{1}; \hat{\theta}^{1}, \hat{\delta}^{1}\right)\right] - \mathrm{g}\left[w\left(x^{1}, z^{0}; \hat{\theta}^{1}, \hat{\delta}^{0}\right)\right] \\
&= \left\{\mathrm{g}\left[w\left(x^{1}, z^{1}; \hat{\theta}^{1}, \hat{\delta}^{1}\right)\right] - \mathrm{g}\left[w\left(x^{1}, z^{0}; \hat{\theta}^{1}, \hat{\delta}^{1}\right)\right]\right\} \\
&+ \left\{\mathrm{g}\left[w\left(x^{1}, z^{0}; \hat{\theta}^{1}, \hat{\delta}^{1}\right)\right] - \mathrm{g}\left[w\left(x^{1}, z^{0}; \hat{\theta}^{1}, \hat{\delta}^{0}\right)\right]\right\}
\end{aligned} \tag{A3}$$

The rows above represent the combined difference, endowment effect, and coefficient effect of z, respectively. The distribution of $w\left(x^{1}, z^{1}; \hat{\theta}^{1}, \hat{\delta}^{1}\right)$ is simulated by

(A1). The counterfactual distribution of $w\left(x^1, z^0; \hat{\theta}^1, \hat{\delta}^1\right)$, which represents the distribution of wage that would prevail if only the distribution of z were like non-unionised workers while all the other cofactors and all the coefficients, including those for z, are the same as unionised workers. We draw on Machado and Mata (2005) to simulate this distribution as follows.

(b1) Classify the n_0 observations for non-unionised workers in the sample into subgroups according to the size of z, and calculate the relative frequency (f_k) for each subgroup. (For discrete variables, the subsets are naturally given and hence classification is not necessary.)

(b2) Consider the distribution $\left\{w_i\left(x^1, z^1; \hat{\theta}^1, \hat{\delta}^1\right)\right\}_{i=1}^{m}$ that was obtained for the decomposition of overall difference. Randomly draw $m \times f_k$ observations with replacement from the subgroup of this distribution for which z belongs to subgroup k. Repeating this for all subgroups provides the desired distribution of $w\left(x^1, z^0; \hat{\theta}^1, \hat{\delta}^1\right)$.

Machado and Mata (2005) only consider the endowment effects and hence they did not construct the distribution $\left\{w_i\left(x^1, z^0; \hat{\theta}^1, \hat{\delta}^0\right)\right\}_{i=1}^{m}$. We obtain this distribution as follows.

(b3) Identify the coefficients, $\hat{\beta}^0(\tau_i)$, that correspond to X_i^1 in constructing $X_i^1\hat{\beta}^0(\tau_i)$ for each τ_i in steps (a4)-(a5).

(b4) For the m observations generated in step (b2), $w\left(x^1, z^0; \hat{\theta}^1, \hat{\delta}^1\right)$, replace $\left\{\hat{\delta}^1\left(\tau_i\right)\right\}_{i=1}^{m}$ with $\left\{\hat{\delta}^0\left(\tau_i\right)\right\}_{i=1}^{m}$ to produce $\left\{w_i\left(x^1, z^0; \hat{\theta}^1, \hat{\delta}^0\right)\right\}_{i=1}^{m}$.

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