

The Trade-Off Between Union Wage Gains and Working Time Flexibility: How Gender Benefits Are Partially Offset

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Abstract

This paper examines whether unions' wage premia come at the cost of reduced working-time flexibility and whether this trade-off affects gender wage disparities in Australia. Using longitudinal data from the Household, Income and Labour Dynamics in Australia Survey from 2005 to 2022, the analysis uses Australia as a setting to investigate this comparatively less-studied question. An instrumental-variable approach is applied using lagged industry-level unionisation rates and flexibility prevalence as instruments. The results indicate a clear trade-off between wage gains and working-time flexibility. Union membership provides substantial wage premia, with larger benefits for women than for men. However, union members also experience systematically lower working-time flexibility. This reduction is particularly costly for women, who earn more from flexibility than men across most of the earnings distribution. When flexibility differences are held constant, the estimated union wage premium falls from 10.1% to 6.0% for men and from 15.1% to 10.4% for women. The remaining gender difference is no longer statistically significant. These findings suggest that constraints on working-time flexibility partially offset unions' gender-equalising potential.

1. Introduction

Unions are central arenas where workers negotiate wages and working conditions. Yet one dimension often overlooked in collective bargaining is working time flexibility, which increasingly shapes how workers balance paid work and caregiving. This issue is gendered because women continue to carry a larger share of unpaid care responsibilities, making schedule autonomy particularly important for their labour-market outcomes. While unions promote equity through collective wage agreements, their emphasis on standardised schedules may unintentionally restrict flexibility, creating a trade-off between higher pay and schedule autonomy. This tension raises a key question: does the very mechanism that promotes pay equity also limit the flexibility that many women need?

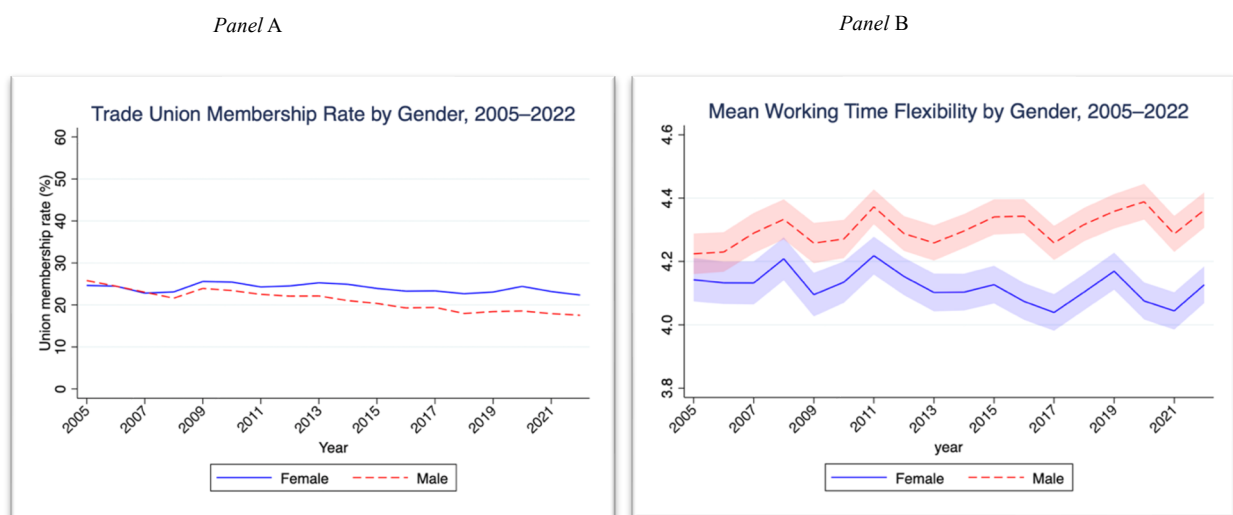
The Australian labour market provides a compelling setting to explore this question. Since 2007, Australian women have consistently shown higher union membership rates than men (see *Panel A* in Figure 1), yet the gender wage gap remains persistent—23% among unionised workers and 25% overall (HILDA Survey, 2022). At the same time, women report lower perceived working time flexibility (see *Panel B* in Figure 1), despite being more likely to use flexible arrangements. These patterns reveal a paradox: unions are expected to equalise pay, but the gender gap endures, suggesting that part of unions' wage benefits may be offset, especially for women who rely more on flexibility to balance work and caregiving. Existing research has examined unions' role in wage compression and gender wage inequality (DiNardo et al., 1996; Card et al., 2004; Blau and Kahn, 2017), as well as the labour-market effects of working time flexibility (Goldin and Katz, 2016; Chung, 2022), but these mechanisms are usually studied separately. Less is known about whether unions' wage gains could be offset by reduced working time flexibility, and whether this trade-off has gendered consequences.

This paper uses longitudinal data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey from 2005 to 2022 to estimate the relationship between union membership, working time flexibility, and real hourly wages. The analysis first estimates the direct association between union membership and wages while controlling for flexibility. It then compares this with the total effect of union membership estimated without flexibility

controls. The difference between these specifications is interpreted as evidence of an indirect flexibility channel through which unions may affect wages.

A major empirical concern is that both union membership and flexibility are potentially endogenous. Workers may select into unionised or flexible jobs based on unobserved productivity, preferences, caregiving constraints, or bargaining power. To address this concern, we use a joint instrumental-variable strategy based on lagged industry-level unionisation rates and lagged industry-level high-flexibility prevalence. These instruments capture institutional variation in union exposure and flexibility norms while reducing the influence of individual-level selection.

Figure 1 Motivating Puzzles



Note: Data source is Household, Income, and Labour Dynamics in Australia (HILDA) Survey (Waves 5–22)¹. Shaded areas represent 95% confidence intervals. *Panel A* shows the union membership rate by gender from 2005 to 2022. *Panel B* shows the average self-reported flexibility (Likert scale: 1-7) in working time arrangements.

The preferred IV estimates show union membership is associated with substantial wage gains for both men and women, with real hourly wages raised by around 10.1% for men and 15.1% for women when working time flexibility is held constant. Then, union membership is also associated with significantly lower perceived working time flexibility. The reduction in flexibility is similar for men and women, indicating that unions constrain schedule autonomy

¹ The mean working time flexibility is calculated for the overall sample. In subsequent empirical analyses, industry fixed effects will be controlled.

across both groups rather than targeting women specifically. Next, once flexibility is allowed to operate as part of the total union effect, the estimated union wage premium becomes smaller. The IV estimates fall from 10.1% to 6.0% for men and from 15.1% to 10.4% for women. This suggests that reduced flexibility partially offsets the direct wage gains from union membership. The gender difference in union wage premia also becomes statistically insignificant, indicating that flexibility constraints partially offset unions' gender-equalizing potential.

To understand why this trade-off matters for gender inequality, we conduct heterogeneity analysis across the wage distribution for four distinct family configurations. Working time flexibility generates higher wage returns for women than for men across much of the earnings distribution. Therefore, even if unions reduce flexibility by a similar amount for both genders, the wage cost of this reduction may be larger for women. This finding helps explain why unions' gender-equalising wage effects are partly weakened once the flexibility channel is considered. Family-structure analysis further shows that partnered mothers benefit most from both union membership and flexibility, while single mothers appear less able to benefit from either mechanism.

Our paper contributes primarily to three strands of literature. First, we contribute to the literature on union wage effects and gender wage inequality. Much of the literature shows that unions raise wages and compress wage dispersion, particularly by increasing wages for lower- and middle-paid workers and limiting managerial discretion in wage setting (Freeman and Medoff, 1984; DiNardo et al, 1996; Card, 2001; Card et al., 2004). Recent evidence from different institutional contexts, suggests that unions may continue to play a gender-equalising role by replacing individual wage bargaining with more transparent and standardised pay structures that raise women's wages relative to men's (Heining et al., 2026; Kostøl et al., 2025; Ma and Zhan, 2024). Our paper is closely related to this literature, but differs in one important respect: rather than estimating the union wage premium alone, we examine whether this wage premium could be indirectly affected by changes in working time flexibility. In doing so, we show that the unions could increase wages directly while also reshaping working conditions in ways that alter the overall value for different groups of workers.

Second, we contribute to the literature on working time flexibility and gendered labour-market outcomes. Existing research shows that flexibility can play an important role in shaping women's employment continuity, hours choices, promotion prospects, and wage progression (Goldin and Katz, 2016; Chung, 2020, 2022; Chung and van der Horst, 2018; Berniell et al., 2023). A parallel literature on flexibility stigma shows that workers who use flexible arrangements may be perceived as less committed, with penalties that are often gendered and particularly relevant for mothers (Correll et al., 2014; Fernandez-Lozano et al., 2020). Building on Blau and Kahn's (2017) evidence that working hours and career interruptions remain important to the gender wage gap, our paper shows that working time flexibility is positively associated with wages with returns that differ across the wage distribution and are higher for women across much of it. This complements Maraziotis's (2024) evidence that flexible working-time arrangements support women's convergence towards men's working hours. This means that similar constraints on flexibility may impose unequal economic costs on men and women.

Third, we contribute to the smaller but growing literature on the interaction between collective bargaining institutions and flexible work arrangements. Earlier work suggests that unions may support family-friendly policies while also resisting forms of flexibility that threaten standardised hours, collective bargaining coverage, or equal treatment, while more recent evidence suggests that unions shape rather than simply oppose flexibility (Budd and Mumford, 2004). Olsson and Skans (2024) and Heining et al. (2026), for example, show that more individualised or discretionary employment arrangements may widen inequality when access to favourable working conditions is uneven. Our paper contributes to this literature by providing the first empirical evidence on the union–flexibility trade-off and its wage implications by gender. The main contribution is therefore to show that unions may reduce gender wage inequality through direct wage gains while also weakening this effect through reduced flexibility, a mechanism that is especially relevant for women's wage progression.

The remainder of this paper is structured as follows: Section 2 introduces the institutional background of unions, flexibility and gendered labour arrangements. Section 3 covers the data

and methodology. Section 4 presents the empirical findings, including the direct, indirect, and total effects, followed by heterogeneity analyses and robustness tests. Section 5 concludes with policy recommendations.

2. Institutional Background: Unions, Flexibility, and Gendered Labour Arrangements

Australia's industrial relations system combines strong collective bargaining traditions with legal mechanisms promoting fairness and workplace protection. The Fair Work Act 2009 formally institutionalises the role of trade unions in wage determination and workplace representation. It codifies many rights that originated from union campaigns, embedding union-negotiated conditions into national awards and enterprise agreements. Through the Act, unions are recognised as bargaining representatives in enterprise negotiations and are granted rights of workplace entry to represent members. These frameworks ensure minimum pay, predictable hours, and transparent wage progression across occupations—mechanisms that have historically narrowed wage inequality and provided stability, particularly for women in low-paid, feminised sectors such as health, education, and community services (Fair Work Ombudsman, 2009; Australian Council of Trade Unions (ACTU), 2024).

Unions negotiate a range of employment terms—not just base pay, but classification ladders, seniority increments, overtime/penalty rates, standard rosters, leave entitlements and job security protections. However, these provisions may matter differently across workers depending on their occupation, industry, job classification, tenure, and working time arrangement. For example, overtime and penalty-rate provisions may be more relevant for full-time or shift-based workers, while rostering rules and predictable hours may be particularly important for workers with caring responsibilities. On the one hand, this reduces the scope for bias, which is especially valuable for women. On the other hand, these same institutions can also make working time arrangements less individually adaptable, with employers sometimes relying on casual employment instead of navigating rigid part-time rostering provisions (Australian Government Attorney-General's Department, 2020).

From a gender perspective, Australian unions have been at the forefront of equal-pay

campaigns and women's labour-force advocacy. For example, union campaigns contributed to the 2012 equal-pay decision for the community-services sector, which raised wages for a predominantly female workforce by 23%–45% (Australian Trade Union Institute, 2021). These examples highlight the gender-equalising role of unions in Australia. The persistent gender pay gap, however, makes Australia a useful and necessary setting for examining whether unions' wage benefits for women could be indirectly offset by reduced schedule autonomy.

3. Data and Empirical Specification

3.1. The HILDA Survey

The data for this study is sourced from the Household, Income and Labour Dynamics in Australia (HILDA) Survey, General Release 22 (Waves 5 – 22). The HILDA dataset is a nationally representative longitudinal household panel survey that provides reliable and comprehensive labour market information. The dataset includes rich and detailed information on individual earnings, working hours per week, union membership status, working time flexibility indicators, gender, job characteristics of employed, household information and working experience. Additionally, industry classifications in the HILDA dataset are coded according to the International Standard Classification of Industries (ISIC).

A nationally representative sample of 19,914 people and 7,682 households participated in the first HILDA survey in 2001. Our empirical analysis uses Waves 5–22 of HILDA, covering the period from 2005 to 2022. The choice of this period is guided by the availability of the working time flexibility measure, which is consistently observed from Wave 5 onwards. This period also allows the analysis to combine a sufficiently long panel with detailed information on union membership, wages, hours worked, education, family characteristics, and industry affiliation. A merged unbalanced panel longitudinal data set of 213,965 observations was produced from the observations of individual workers with 13,435 men and 13,169 women.

A major advantage of the HILDA data set is that it provides variation across individuals, industries, and years, including changes in union status and reported flexibility, which supports the use of panel and IV methods.

3.2. Variables

As shown in Appendix Table A1, the dependent variable is the log of the respondent's real hourly earnings from their main job, defined as the job with the highest number of hours worked. Gross hourly wages are calculated by dividing current weekly gross income by the usual weekly hours worked in that job. Nominal hourly wages were converted to real terms using the OECD annual Consumer Price Index (CPI, base = 2015) to account for inflation over 2005–2022, and then logged. Taking the logarithm allows the estimated coefficients to be interpreted approximately as percentage differences in wages.

The first key explanatory variable is union membership. This is measured as a binary indicator equal to one if the respondent reports being a member of a trade union, and zero otherwise. Union membership captures the wage effect associated with collective bargaining, institutional wage-setting, and union representation. The second key explanatory variable is working time flexibility, which is measured on a seven-point Likert scale (strongly disagree to strongly agree of “My working time can be flexible”)². Gender is also included as a key variable because the study examines whether the relationship between union membership, flexibility, and wages differs between men and women. Gender is measured as a binary variable equal to one for female respondents and zero for male respondents. The coefficient of Gender can be interpreted as the extra wage effects of unions or working time flexibility on women compared to that on men.

Control variables in the wage regression model consist of work experience, education, marital status, number of children under age of 5, and part-time status, all of which are well-established determinants of earnings in the human capital literature (Pyatt & Graham, 1966; Waldfogel, 1998). Work experience is years since left full-time education. Education-equivalent years are imputed from respondents' highest educational qualification following

²Although the Fair Work Act 2009 provides employees with a statutory “right to request” flexible work, the Senate Standing Committee on Economics (2016) noted that employers can lawfully refuse on “reasonable business grounds” and that no appeal mechanism exists. As a result, perceived flexibility is used to capture actual schedule autonomy rather than formal entitlement. It also aligns with Gregory and Milner's (2009) distinction between employee-friendly and employer-friendly flexibility, as subjective assessments better capture whether flexibility is experienced as worker autonomy.

Zawadzki and Wojdyło (2018). Finally, the partnered status is captured by a binary variable indicating whether the respondent has a partner, including both legally married and de facto relationships.

Table 1 Descriptive statistics

	Men				Women			
	All	Union	Non-union	Difference	All	Union	Non-union	Difference
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Real hourly wage of main job	30.23 (29.42)	35.68 (29.59)	28.70 (28.14)	6.97*** (0.21)	28.46 (33.04)	33.29 (22.29)	26.93 (35.66)	6.37*** (0.24)
Flexibility scale	4.31 (1.95)	3.73 (2.00)	4.47 (1.90)	-0.74*** (0.02)	4.12 (2.07)	3.40 (2.07)	4.35 (2.01)	-0.95*** (0.02)
Education years	12.87 (1.93)	13.11 (2.04)	12.81 (1.90)	0.30*** (0.01)	13.19 (2.06)	13.97 (2.16)	12.94 (1.96)	1.04*** (0.02)
Number of children under age 5	0.22 (0.55)	0.22 (0.55)	0.22 (0.55)	0.00 (0.00)	0.16 (0.46)	0.16 (0.45)	0.17 (0.47)	-0.01*** (0.00)
Work experience	26.6 (12.97)	27.91 (11.72)	26.19 (13.31)	1.72*** (0.10)	26.34 (12.58)	28.12 (11.89)	25.68 (12.76)	2.44*** (0.1)
Partnered	0.69 (0.46)	0.75 (0.43)	0.67 (0.47)	0.09*** (0.00)	0.643 (0.48)	0.688 (0.46)	0.629 (0.48)	0.06*** (0.00)
Number of observation	110,527	24,111	86,416		103,438	25,018	78,420	

Note: Data source is HILDA Survey (Waves 5-22). The figures are averages and standard deviations are in parentheses. For the difference in means, standard errors are reported in parentheses. The number of men is 13,435 and women is 13,169. Hourly wages are expressed in real terms (2015 AUD) using the OECD annual CPI. Differences in columns (4) and (8) represent the disparities between union and non-union membership for men and women respectively. Education-equivalent years are imputed from respondents' highest educational qualification following Zawadzki and Wojdyło (2018). Statistical significance of the differences is based on Wald tests using robust standard errors. *** p<0.01, ** p<0.05, * p<0.10.

Table 1 presents the descriptive statistics for the main variables, disaggregated by gender and union membership status. The difference of real hourly wage of main job in Column (4)

and (8) reveals a clear gender wage gap, with men earning higher average hourly wages than women across both union and non-union sectors. On average, male union members earn 6.97 AUD more per hour than their non-union counterparts, while the gap for women is 6.37 AUD. This preliminary evidence suggests that unions may play a role in promoting wages, though the magnitude and statistical significance of the gendered effect requires formal econometric analysis.

Regarding working time flexibility, men report higher perceived flexibility scores than women, despite women's greater utilization of flexible work arrangements documented by Diversity Council Australia (2024). This distinction reflects the difference between formal policy usage and subjective schedule autonomy. Most importantly, unionised workers report substantially lower flexibility scores than non-unionised workers, with the union-non-union flexibility gap being larger for women (-0.948) than for men (-0.743). Among these, unionised women experience the lowest flexibility levels across all groups.

Besides, as a validation of the subjective flexibility measure, Appendix Figure A 3 shows that greater scale of working time flexibility is associated with a wider dispersion of working hours across workers. This suggests that the flexibility scale captures meaningful differences in working time arrangements rather than only subjective preferences.

What's more, Table A 2 reveals several important patterns in union membership dynamics. Most workers never join a union, while around 28% switch union status during the observation period. Gender differences in union membership dynamics are minimal. This switching pattern provides within-individual variation that supports the use of fixed-effects estimation to control for time-invariant individual heterogeneity later.

3.3. Model Specification

Following the reduced-form approach in applied microeconometrics, we interpret differences between specifications with and without flexibility as capturing an indirect channel through which unions affect wages (Angrist & Pischke, 2009).

We first show the direct effect of union memberships on wages using Equation (1):

$$\ln(wage)_{it} = \theta_0 + \theta_1 Un_{it} + \theta_2 Fx_{it} + \theta_3 X_{it} + \lambda_t + \delta_i + \varepsilon_{it} \quad (1)$$

Where $\ln(Wage_{it})$ represents the natural logarithm of the real hourly wage (inflation-adjusted using OECD CPI, base = 2015) of main job for individual i at time t , facilitating the interpretation of coefficients in terms of percentage changes; Un_{it} is the binary indicator for union membership; Fx_{it} represents working time flexibility (strongly disagree to strongly agree: 1-7); X_{it} represents individual controls; λ_t represents year-specific fixed; δ_i captures individual-specific effects; ϵ_{it} is the error term.

Consistent with a reduced-form interpretation, we examine whether working time flexibility operates as an indirect channel in the relationship between union membership and wages using a two-step method:

$$Fx_{it} = \gamma_0 + \gamma_1 Un_{it} + \gamma_2 X_{it} + \lambda_t + \delta_i + \epsilon_{it} \quad (2)$$

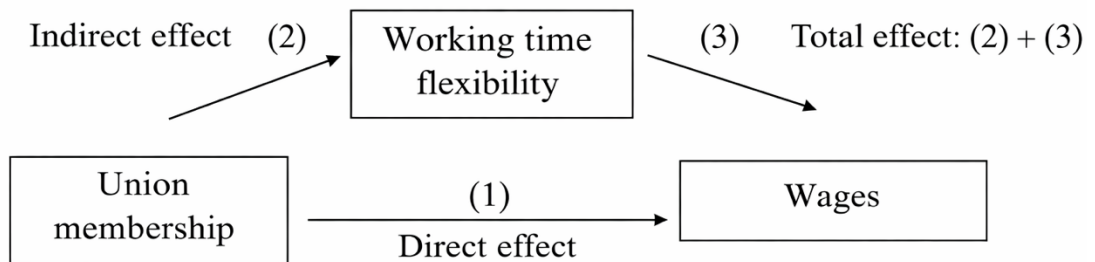
Then, we conduct a total effect of union memberships on wages using Equation (3) as follows:

$$\ln(wage)_{it} = b_0 + b_1 Un_{it} + b_2 X_{it} + \lambda_t + \delta_i + \epsilon_{it} \quad (3)$$

Comparing the total and direct effects reveals the role of flexibility as an indirect channel. Specifically, if the total effect b_1 is smaller than the direct effect θ_1 , this indicates that working time flexibility constrained the wage premium to some extent, serving as an important indirect mechanism.

The conceptual mechanism of framework is illustrated in Figure 2.

Figure 2 Conceptual Framework for the Union–Flexibility–Wage Mechanism



The direct effect (Path 1) captures the wage premium associated with union membership. The indirect path operates through working time flexibility: union membership may influence perceived flexibility (Path 2), which in turn affects wages (Path 3). The total effect of union

membership on wages therefore combines both the direct and indirect components (Path 2 + Path 3) allowing us to assess whether union wage gains are partly offset by reduced schedule autonomy.

We estimate the models using three approaches. Pooled OLS provides a baseline estimate, fixed effects control for time-invariant individual heterogeneity, and joint IV estimation addresses the potential endogeneity of both union membership and working time flexibility. To examine gender differences, we estimate models separately for men and women and test coefficient differences using pooled regressions with female interaction terms. All specifications include year and industry effects.

After that, heterogeneity analysis based on wage distribution effects (Foresi & Peracchi, 1995) and family structure will be conducted to show the trade-offs faced by different segments of the workforce. Finally, a robustness test will be conducted to ensure our findings are robust to alternative instrument specifications.

3.4. Endogeneity concerns and IV strategy

Union membership is likely endogenous in wage equations due to several theoretical concerns identified in the literature. Following Booth and Bryan (2001), workers may sort into union membership based on unobservable characteristics such as productivity, motivation, or firm-specific investment preferences. Additionally, reverse causality may operate if higher-wage workers have different propensities to join unions, either seeking protection (positive selection) or avoiding unions due to superior individual bargaining power (negative selection).

Working time flexibility presents similar endogeneity concerns. Workers may self-select into flexible jobs based on unobserved preferences for work-life balance or caregiving responsibilities, while employers may offer flexibility strategically to attract or retain certain types of workers. Without controlling for these selection processes, estimates of flexibility's wage effects may be biased.

To address these concerns, we employ a joint instrumental variables approach that simultaneously instruments for both union membership and working time flexibility. For union membership, we use one-year-lagged industry-year-region unionization rate as an instrumental

variable, following Booth and Bryan (2001)⁴. This instrument captures the institutional environment affecting union participation—workers in highly unionised sectors face greater social pressure and opportunity to join unions. The exclusion restriction is plausible because industry-level unionization rates reflect historical organizing patterns and institutional legacies that are largely predetermined relative to current individual wage determination. The one-year lag further ensures that the instrument is predetermined and not influenced by contemporaneous labour market shocks. These rates vary across industry-year-region cells, capturing local institutional environments while our specification controls for broad industry characteristics through industry fixed effects, making spillover effects unlikely to violate the exclusion restriction.

For working time flexibility, we use the one-year-lagged industry-year-region prevalence of high perceived flexibility (proportion reporting flexibility scores ≥ 5) as an instrument. High-flexibility prevalence captures institutional differences between high-flexibility and low-flexibility industries, which are exogenous to individual wages after controlling for industry effects. Following Angrist & Pischke (2009), the binary IV allows us to apply the Wald estimator, which identifies the Local Average Treatment Effect (LATE) — the effect for workers whose flexibility status is actually shifted by industry-level norms.⁵ Our estimates therefore represent the policy-relevant effect for workers on the margin of flexibility change. The instruments exhibit substantial variation across industries and over time, as shown in Appendix Figure A 1-Figure A 2 and Appendix Table A3, supporting their relevance for identifying institutional exposure to unionisation and flexibility norms.

⁴ In robustness check later, we employ an alternative IV constructed as the rolling average of the industry union rate over the past three years capture short-term fluctuations in unionization.

⁵ In contrast to binary indicator, continuous IVs such as mean value of flexibility are often noisier, more sensitive to extreme values, and typically yield a weighted average of local effects, making their causal interpretation less straightforward. Besides, we compare several candidate instruments based on the proportion of workers within each industry-year reporting flexibility above different thresholds (≥ 3 , ≥ 4 , ≥ 5 , ≥ 6) as well as the industry-level mean flexibility in Appendix Table A8. Although the mean flexibility and the ≥ 4 threshold yield slightly higher first-stage F-statistics, they also exhibit higher R^2 when regressed on industry and year fixed effects, indicating greater collinearity with the fixed effects and less usable within-industry variation. The ≥ 5 threshold strikes the best balance: it delivers a very strong first stage while achieving a lower R^2 , meaning the IV is less collinear with the fixed effects and captures more meaningful within-industry variation over time.

Finally, the first-stage regressions in Table 2 demonstrate strong instrument relevance, with Sanderson–Windmeijer first-stage F-statistics ranging from 1429.53 to 3034.41, well exceeding conventional weak instrument thresholds.⁶ The diagnostic tests (Appendix Table A5) support instrument validity: Kleibergen-Paap underidentification tests strongly reject underidentification ($p < 0.001$), while weak identification statistics substantially exceed Stock-Yogo critical values. These results provide confidence in our identification strategy for isolating the causal effects of union membership and working time flexibility on wages.

Table 2 First-stage Results of IV Estimation

Dependent variable	Men		Women	
	Union	Flexibility	Union	Flexibility
	(1)	(2)	(3)	(4)
Lagged Industry Unionization Rate	0.8897*** (0.0157)	-0.3245*** (0.0682)	0.8329*** (0.0214)	-0.3607*** (0.0966)
Lagged Industry High-flexibility Rate	-0.0293** (0.0115)	2.6788*** (0.0544)	-0.0244 (0.0147)	2.8883*** (0.0757)
The F-value of Sanderson- Windmeijer weak identification test	3034.41***	2370.31***	1511.45***	1429.53***

Note: All regressions include year fixed effects and control variables as specified in the main text. F(excluded) reports the F-statistic for excluded instruments in each first-stage regression. SW F-statistic is the Sanderson-Windmeijer first-stage F-statistic testing the strength of identification for each endogenous variable. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

4. Results

4.1. Union Membership and Working Time Flexibility: Direct Effects on Wages

Table 3 presents the main results examining Equation (1) using multiple estimation strategies. We begin with POLS and FE estimates before proceeding to instrumental variable approaches that treat union membership and flexibility as endogenous variables jointly.

Union membership demonstrates robust and substantial positive effects on wages across all specifications. The FE estimates in Column (2) and (5) consistently show smaller coefficients than OLS estimates in Column (1) and (4), reflecting the effect of removal of individual-specific characteristics and time-invariant unobserved heterogeneity. This suggests that OLS estimates are upward biased due to positive selection into union membership—

⁶ The full first-stage regression result can be found in Appendix Table A4.

workers with higher unobserved productivity are more likely to join unions. The FE estimates rely primarily on the 28% of workers who switch union status (Table A2), excluding the stable union and non-union workers who contribute to between-individual variation in OLS.

The estimates under our preferred IV method in Columns (3) and (6) reinforce the positive association between union membership and wages and are somewhat larger than the corresponding POLS estimates.⁷ For men, the union wage premium increases from 8.4% under POLS to 10.1% under IV estimation. For women, the union wage premium rises from 13.4% under POLS to an impressive 15.1% under IV estimation. Once the potential endogeneity of union membership and working time flexibility is addressed through the IV strategy, the estimated wage gains associated with union membership become slightly larger. The gender difference in union effects, estimated at 5.0% (Column 7), strongly indicates that unions provide larger wage premia for women and help narrow the gender pay gap when flexibility is accounted for.

Working time flexibility exhibits consistently positive relationships with wages across almost all specifications. For men, a one-unit increase in the flexibility scale is associated with wage premia of 4.3% (IV). For women, increased flexibility is associated with a 4.5% wage premium. The gender difference is small and statistically insignificant, suggesting that flexibility has similar average wage returns for men and women.

Nevertheless, the FE estimates reveal much smaller flexibility premia for both genders, suggesting that the cross-sectional flexibility-wage relationship largely reflects selection effects rather than causal impacts.

However, the direct effects presented above reveal an intriguing puzzle that requires

⁷ We do not include individual fixed effects in the IV specifications because our instrumental variables primarily exploit between-individual variation across different industry-region-time contexts. Given that approximately 63% of individuals maintain stable union membership status throughout the observation period (Table A1), controlling for individual fixed effects would eliminate most of the variation that our instruments rely upon. The IV strategy identifies effects by comparing individuals facing different institutional environments.

Table 3 Direct Effects of Union Membership and Flexibility on the log of Real Hourly Wages

	Men			Women			Gender Difference (Female-Male) IV
	POLS (1)	FE (2)	IV (3)	POLS (4)	FE (5)	IV (6)	
Union membership	0.084*** (0.005)	0.033*** (0.008)	0.101*** (0.019)	0.134*** (0.004)	0.025*** (0.007)	0.151*** (0.028)	0.050* (0.030)
Flexibility	0.017*** (0.001)	0.001 (0.001)	0.043*** (0.006)	0.019*** (0.001)	0.005*** (0.001)	0.045*** (0.007)	0.002 (0.009)
Experience	0.024*** (0.001)	0.026*** (0.008)	0.024*** (0.001)	0.014*** (0.001)	0.014* (0.007)	0.014*** (0.001)	
Experience ² /100	-0.038*** (0.001)	-0.037*** (0.003)	-0.038*** (0.002)	-0.021*** (0.001)	-0.025*** (0.003)	-0.021*** (0.001)	
Partnered	0.102*** (0.005)	0.024** (0.009)	0.098*** (0.005)	0.042*** (0.004)	0.031*** (0.009)	0.038*** (0.004)	
Schooling	0.024*** (0.001)		0.023*** (0.001)	0.028*** (0.001)		0.027*** (0.001)	
Children	0.032*** (0.004)	0.018*** (0.004)	0.031*** (0.004)	0.080*** (0.005)	0.034*** (0.006)	0.072*** (0.005)	
Constant	-1.349*** (0.024)	-0.625*** (0.153)	-0.879*** (0.032)	-1.251*** (0.036)	-0.530*** (0.143)	-0.752*** (0.043)	
Individual Effects	No	Yes	No	No	Yes	No	No
Year & Industry Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	50229	50229	50229	51423	51423	51423	101652
R-squared	0.294	0.256	0.286	0.261	0.236	0.248	0.327
Adj. R-squared	0.298	0.277	0.298	0.260	0.235	0.247	0.271

Note: Data source is HILDA Survey (Waves 1-22). Dependent variable is log of real hourly wages. Standard errors in parentheses. Columns (1)-(6) show results from separate regressions by gender using POLS, FE, and pooled IV estimation respectively. Column (7) shows gender differences estimated from pooled IV regression with interaction terms of female variable and all independent variables. POLS uses robust standard errors. FE uses within-group estimation with robust standard errors, excluding time-invariant variables (schooling, industry). IV uses one-year-lagged industry-year-region unionization rate and one-year-lagged flexibility rate as instruments for union membership and flexibility respectively. Year and industry effects are controlled for all estimations. All regressions include controls for experience, experience squared, partnered status, and number of children under 5. Industry and year fixed effects included where applicable. ***p<0.01, **p<0.05, *p<0.1.

further investigation. Our results demonstrate that union membership can reduce gender wage disparities in Australia. Yet this raises a fundamental question about the underlying mechanisms: if flexibility could independently enhance wages for women, how might these benefits change when unions promote standardised working arrangements that typically reduce schedule flexibility? What would be the total effect of union membership on gender wage gaps once we allow flexibility to change as union memberships?

To address this puzzle, we proceed with two complementary analyses. First, we examine how union membership systematically affects working time flexibility, particularly for women, which is relevant with institutional theory and prior research (Budd & Mumford, 2004; Heyes, 2011). Second, we estimate the total effects of union membership on wages without controlling for flexibility, allowing us to observe how the potential union-flexibility trade-off influences the overall relationship between unions and gender wage gaps.

4.2. Indirect and Total Effect Analysis

Table 4 presents the crucial test of our central hypothesis by examining both the indirect pathway (Panel A) and total effects (Panel B). This analysis addresses two fundamental questions: first, how do unions systematically affect working time flexibility as shown in Equation (2); and second, whether the substantial direct wage benefits could be offset when flexibility is allowed to vary endogenously, as shown in Equation (3).

Panel A examining Equation (2) provides evidence that union membership systematically constrains working time flexibility across all estimation strategies and gender. For men, the IV estimate in Column (3) shows that union membership reduces flexibility by 0.948 points on the 7-point scale. For women, the flexibility constraints are even more pronounced, with IV estimates indicating a reduction of approximately 1 points in Column (6).

The gender difference in flexibility constraints is not statistically significant (Column 7), confirming that unions impose significant and similar flexibility constraints on both genders. The finding above aligns with institutional theories suggesting that unions' emphasis on collective standardization resist flexibility policies that threaten standardised working hours (Budd & Mumford, 2004).

Panel B examining Equation (3) reveals the crucial evidence for our central hypothesis. For men, the total union wage effect under IV estimation (6.0%) is smaller than the 10.1% direct effect documented in Table 3 (Column 3). For women, the total union wage effect (10.4%) is considerably smaller than the 15.1% direct effect from Table 3 (Column 6). This indicates that flexibility constraints offset approximately 31% of the potential wage benefits that unions could provide to women. Despite the substantial offset, the total effect remains positive and significant, suggesting that women still benefit from union membership on net, but not to the extent that direct wage effects would suggest.

To formally assess whether the estimated union effects differ across specifications, we compute bootstrap-based tests of the difference between direct effect and total effect. The results (Appendix Table A9) show that the difference is statistically significant for both men

and women, indicating that the inclusion of flexibility substantially reduces the estimated union wage premium.

Interestingly, the FE pattern differs because the within-person flexibility effect is small among union switchers.

Most importantly, the gender difference in the total union effect (4.4%) is no longer statistically significant once flexibility is allowed to operate endogenously (Column 7), compared to the significant 5.0% difference in the direct specification (Table 3). While a formal bootstrap test indicates that the reduction in the coefficient is not statistically significant at conventional levels (see Appendix Table A9), suggesting that while flexibility attenuates the estimated gender gap in union effects, it does not significantly alter it.

This overall result helps explain the apparent contradiction between rising female unionization rates and persistent gender wage gaps in unionised sectors. While the ACTU (2024) has highlighted progress in reducing the gender pay gap through union initiatives, our analysis demonstrates that this progress is fundamentally constrained by the inherent trade-off between wage gains and flexibility constraints that unions create.

Table 4 Indirect and Total Effects Analysis

	Men			Women			Gender Difference (Female-Male)
	POLS	FE	IV	POLS	FE	IV	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Union Effects on Working Time Flexibility							
Union membership	-0.702 ^{***} (0.020)	-0.203 ^{***} (0.030)	-0.948 ^{***} (0.076)	-0.625 ^{***} (0.020)	-0.228 ^{***} (0.031)	-0.999 ^{***} (0.113)	-0.052 (0.136)
Other regressors	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual Effects	No	Yes	No	No	Yes	No	No
Year & Industry Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	57807	57807	57760	55749	55749	55731	77795
R-squared	0.106	0.013	0.103	0.138	0.024	0.133	0.831
Adj. R-squared	0.105	0.012	0.102	0.137	0.023	0.131	0.120
Panel B: Total Effects of Union on the Log of Real Hourly Wages (without Flexibility controlled)							
Union membership	0.074 ^{***} (0.005)	0.035 ^{***} (0.007)	0.060 ^{***} (0.017)	0.120 ^{***} (0.004)	0.023 ^{***} (0.007)	0.104 ^{***} (0.026)	0.044 (0.031)
Other regressors	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual Effects	No	Yes	No	No	Yes	No	No
Year & Industry Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	57489	57489	57489	57219	57219	57219	114708
R-squared	0.219	0.075	0.219	0.161	0.060	0.160	0.319
Adj. R-squared	0.218	0.073	0.218	0.159	0.059	0.159	0.274

Note: Standard errors in parentheses. Panel A presents indirect analysis with working time flexibility (1-7 scale) as the dependent variable, while Panel B shows total effects analysis with log of real hourly wages (inflation-adjusted using OECD CPI, base = 2015) as the dependent variable. Columns (1)-(6) display results from separate regressions by gender using POLS, FE, and IV estimation respectively. Column (7) presents gender differences estimated from pooled IV regressions with interaction terms. Gender differences represent coefficients of interaction terms of female variable and all independent variables. IV specifications use industry-year-region unionization rates as instruments for union membership. Year and industry effects are controlled for all estimations. All regressions include comprehensive controls for work experience, experience squared, educational attainment, partnered status, number of children under age 5, and year fixed effects.

*** p<0.01, ** p<0.05, * p<0.10.

4.3. Heterogeneity Analysis

4.3.1. Wage distribution

Since unions impose similar flexibility constraints on both genders, this raises a deeper theoretical question: why do equal flexibility constraints generate differential impacts on gender wage gaps? To understand the mechanisms behind this asymmetric impact, we next examine how union and flexibility effects vary across the wage distribution, which may reveal why equal constraints produce unequal outcomes. Figure 3 reports the heterogeneous effects of union membership and working time flexibility across the wage distribution, estimated using an IV-based distribution regression framework. Following the intuition of Foresi & Peracchi (1995), I create a series of binary response models at each percentile threshold of the wage distribution from the 10th to the 90th, increasing in steps of 5 percentile points:

$$Y_{itp} = 1\{\ln(wage_{it}) > w_{pt}\} \quad (4)$$

For each percentile threshold w_{pt} , binary variable Y_{itp} equals one if the log wage of individual i at year t is above the threshold at p percentile and zero otherwise.

Then, I estimate the following instrumental-variable binary regression:

$$P(Y_{itp} = 1 | Un_{it}, Fx_{it}, X_{it}) = F(\theta_p + \theta_{1p}Un_{it} + \theta_{2p}Fx_{it} + \theta_{3p}X_{it}) \quad (5)$$

where $F(\cdot)$ can be interpreted as a cumulative distribution function. In practice, a linear probability model is implemented with clustered standard errors at the individual level. The estimated marginal effects of union membership (θ_{1p}) and flexibility (θ_{2p}) are allowed to vary across wage percentiles, capturing potential heterogeneity along the wage distribution.

We extend the Foresi & Peracchi (1995) approach by employing the same instrumental variables as described in Section 3.4. to address potential endogeneity of both union membership and working time flexibility. This framework enables us to examine not only how the effects of unions and flexibility vary across the wage distribution but also whether these patterns differ by gender.

Panel A of Figure 3 shows that unions provide a significant positive wage premium for most workers, especially between the 30th and 60th percentiles. This highlights unions' strong role in raising wages for middle-income earners. At the bottom of the distribution, the union

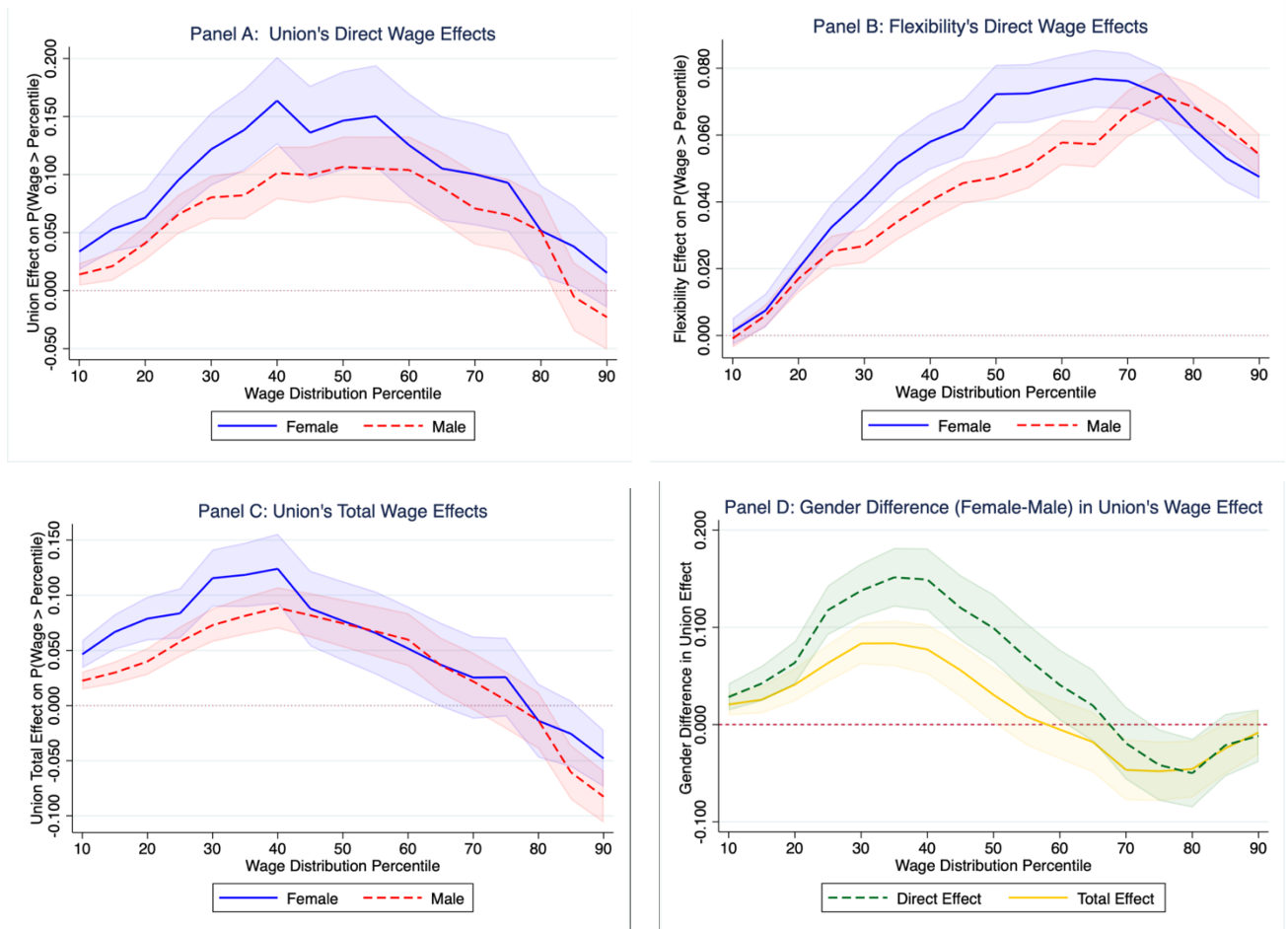
effect is small because statutory minimum wages already act as wage floors, leaving less space for additional bargaining gains. At the top of the distribution, however, the effect diminishes and even turns

negative, suggesting that high-income workers rely less on collective bargaining and more on individual negotiation power. Thus, middle-income workers are the group most likely to benefit from collective wage negotiations, as they are neither protected by binding wage floors nor endowed with high individual bargaining leverage.

Figure 3's Panel B reports the heterogeneous effects of working time flexibility on wages. Flexibility premia are small or absent at the bottom of the wage distribution, possibly reflecting a "flexibility stigma" (Correll et al., 2014; Williams et al., 2013), where workers in low-wage jobs face penalties for requesting flexible arrangements. The flexibility effect rises steadily across the distribution, peaking around the 50th to 70th percentiles, implying that upper-middle-income earners capture the largest wage premia from flexibility. However, this premium declines sharply among the highest earners, likely because their total compensation relies more on unique skills and equity incentives rather than contracted hours.

Also, flexibility generates systematically higher wage returns for women than men across most of the earnings distribution, supporting Goldin and Katz's (2016) finding that appropriately structured flexibility can enhance rather than constrain women's earning potential. While unions impose similar flexibility constraints on both genders, these constraints create asymmetric economic costs. Women, who demonstrate higher dependency on flexibility for wage optimization, experience larger indirect wage losses when unions limit schedule autonomy. This explains why unions' gender-equalizing benefits are partially offset in total effects.

Figure 3 Heterogeneous Union and Flexibility Effects Across the Wage Distribution



Note: Data source is Household, Income, and Labour Dynamics in Australia (HILDA) Survey (Waves 5–22). Shaded areas represent 90% confidence intervals computed using robust standard errors clustered at the individual level. Gender differences in Panel D represent coefficients of interaction terms of female variable and all independent variables. All regressions include controls for experience, education, partnership status, children, industry, and year fixed effects.

Figure 3's Panel C shows the total effect of union membership, combining both wage gains and flexibility losses. Compared to Panel A, the estimated wage premium is substantially smaller, suggesting the presence of a trade-off: while unions raise wages, they also limit working time flexibility, and this offsetting pattern is more pronounced for middle-income workers.

Finally, Panel D compares the gender difference of union's direct effects with total effects. Women generally benefit more wage gains from unions than men, especially around the 30th to 50th percentiles. However, once flexibility is controlled for, this gendered benefits narrow noticeably. It reflects that the attenuation on the gender wage gap is most pronounced in the

middle of the wage distribution, consistent with a stronger offsetting pattern in that range.

Overall, unions help raise wages, particularly for middle-income workers, but their benefits are partially offset by reduced working time flexibility. Flexibility premia are concentrated among workers in the 50th-70th percentiles of the wage distribution. And women derive substantially higher wage returns from flexibility than men across most of the distribution. This differential dependency on flexibility explains why equal union constraints produce asymmetric gender effects. Thus, gender differences in union effects are driven by differential reliance on flexibility for wage optimization.

4.3.2. Family Structure

While the wage-distribution analysis shows where the union-flexibility trade-off is most pronounced, it does not fully explain which women are most exposed to this trade-off. I therefore further examine family structure, since partnership status and the presence of young children may affect women's flexibility needs and returns to union membership. To examine this, we estimate our direct effects model across four distinct family configurations. Table 5 presents the IV estimation results, revealing distinct patterns based on partnership status and the presence of young children.

Table 5 Union and Flexibility Direct Effects on Women by Family Structure under IV estimation

	(1) Single, No Child	(2) Single w/ Child	(3) Partnered, No Child	(4) Partnered w/ Child
Union	0.098** (0.049)	0.090 (0.060)	0.141** (0.060)	0.245*** (0.086)
Flexibility	0.028 (0.018)	0.016 (0.017)	0.024 (0.022)	0.068*** (0.014)
<i>N</i>	12890	559	32182	5777

Note: IV estimates for women only. All regressions include controls for experience, education, partnership status, children, industry, and year fixed effects. Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Union wage premia are concentrated among partnered women, particularly those with children. Partnered mothers gain the largest and most significant premium (24.5%), followed by partnered, childless women (14.1%). In contrast, single women receive smaller and less precise benefits: childless singles gain about 9.8%, while single mothers receive a premium of 9.0%, which is statistically insignificant.

One explanation is that partnered women — especially mothers — may find unions’ standardised work arrangements more compatible with their household division of labour. In dual-income families, having a predictable schedule can facilitate coordinated caregiving responsibilities. Conversely, single mothers may require more individualised flexibility than unions typically provide, which limits their ability to fully benefit from collective agreements.

Flexibility effects follow a different pattern. While flexibility is not significantly related to wages for single women and childless women, it is clearly valuable for partnered mothers, who experience the largest and most significant return (6.8%).

The absence of flexibility premia for childless women may reflect a stronger “flexibility stigma” (Correll et al., 2014; Williams et al., 2013), as single mothers and childless women may reflect unequal access to high-quality flexible jobs. In contrast, for mothers, flexibility use is socially legitimised as necessary for caregiving and may even receive policy and workplace support (Glass & Estes, 1997).

Overall, the results suggest that union standardisation primarily benefits partnered women, especially mothers, while flexibility is also most valuable for partnered mothers. Single mothers, however, remain disadvantaged under both mechanisms, highlighting the need for integrated policy approaches that combine collective wage protection with caregiving-compatible flexibility.

4.4. Robustness Test

As an additional robustness test, we construct an alternative IV using the three-year rolling average of the industry union rate. This specification captures unionisation patterns over a longer time span and smooths short-term fluctuations in industry unionisation.

Table 6 compares our main results with this alternative specification. Robustness checks using a three-year rolling average of lagged industry unionisation rates produce similar results. The direct union premium remains larger for women than for men, while the gender difference becomes statistically insignificant in the total-effect specification. First-stage and diagnostic tests reported in Appendix Table A6 and Table A7 also remain strong, suggesting that the main findings are not driven by the specific one-year lag structure of the instrument.

Table 6 Robustness Test – Lagged Industry Unionization Rate

	Main IV (one-year-lagged)				Robustness (rolling rate of past 3 years)			
	Direct		Total		Direct		Total	
	Men	Women	Men	Women	Men	Women	Men	Women
Union	0.101*** (0.019)	0.151*** (0.028)	0.060*** (0.018)	0.104*** (0.026)	0.107*** (0.018)	0.136*** (0.026)	0.064*** (0.016)	0.086*** (0.024)
Flexibility	0.043*** (0.006)	0.045*** (0.007)			0.044*** (0.006)	0.045*** (0.007)		
Gender difference of Union	0.050* (0.030)		0.044 (0.031)		0.039* (0.030)		0.018 (0.018)	
<i>N</i>	50229	51423	57489	57219	50220	51395	57421	57190

Note: Main IV uses one-year lagged industry unionization rate, robustness test uses three-year rolling average of lagged industry unionization rates. Both specifications use one-year-lagged industry flexibility prevalence (≥ 5) as the flexibility instrument. All regressions include comprehensive controls for work experience, experience squared, educational attainment, partnered status, number of children under age 5, year and industry fixed effects. Standard errors clustered at the individual level are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The consistency of results across different instrument definitions validates our identification strategy and strengthens confidence in the causal interpretation.⁸

5. Conclusion

This paper examines the trade-off between union wage gains and working time flexibility and its implications for gender pay disparities in Australia using longitudinal data from the HILDA Survey (2005-2022). It addresses a comparatively less-studied question: whether flexibility acts as an indirect channel through which unions influence gender wage disparities. By considering both the gender-equalising wage benefits associated with unionisation and the extent to which reduced working time flexibility offsets these gains, the paper provides a more comprehensive assessment of the wage implications of union membership for men and women.

The results show that union membership is associated with higher wages for both men and women. When working time flexibility is held constant, the preferred IV estimates indicate union wage premia of 10.1% for men and 15.1% for women. However, union membership is

⁸We do not replace the main IVs with smoothed multi-year averages, as this would shorten the effective sample period by approximately two years. Given the study period of 2005–2022, such a restriction would result in a non-negligible reduction in sample coverage. Therefore, we retain the baseline IV specification to preserve both sample size and temporal variation.

also associated with lower perceived working time flexibility, with reductions of a similar magnitude for both genders. Once flexibility is allowed to operate as part of the overall relationship between union membership and wages, the estimated union premium falls to 6.0% for men and 10.4% for women. The gender difference in the union wage premium also becomes statistically insignificant. These findings suggest that reduced flexibility partially offsets the wage gains associated with union membership and weakens unions' potential contribution to narrowing gender wage disparities.

Heterogeneity analysis highlights substantial variation in outcomes across both the earnings distribution and family structures. Union wage benefits are concentrated among middle-income workers, with limited effects at the lower end and even negative impacts at the top. Flexibility premia are concentrated among middle-income workers. And women derive substantially higher wage returns from flexibility than men across most of the distribution, meaning that equal flexibility constraints impose asymmetric economic costs by gender. The differential dependency on flexibility explains why equal union constraints partially offset the gender-equalizing effect of unions. The results also vary across family structures: partnered mothers appear to benefit most from both union membership and flexibility, whereas single mothers benefit less from either mechanism. This pattern highlights a persistent institutional gap where neither collective wage protection nor existing flexibility provisions adequately meet their needs. The results are also robust to alternative instrument definitions.

However, women may nevertheless value unions because the net wage premium remains positive and because unions provide non-wage protections that are not captured in the wage analysis.

From a policy perspective, the findings suggest that strengthening collective wage protection and promoting working time flexibility should not be treated as separate objectives. Institutional arrangements that preserve the wage-setting benefits of collective bargaining while allowing greater scope for schedule accommodation may be particularly important for workers with caregiving responsibilities.

More broadly, the results show that the distributional consequences of unionisation extend

beyond wages alone. While unions raise wages and can reduce gender wage disparities, part of these gains may be offset through reduced working time flexibility. Accounting for both channels therefore provides a more complete assessment of how union institutions shape gendered labour-market outcomes.

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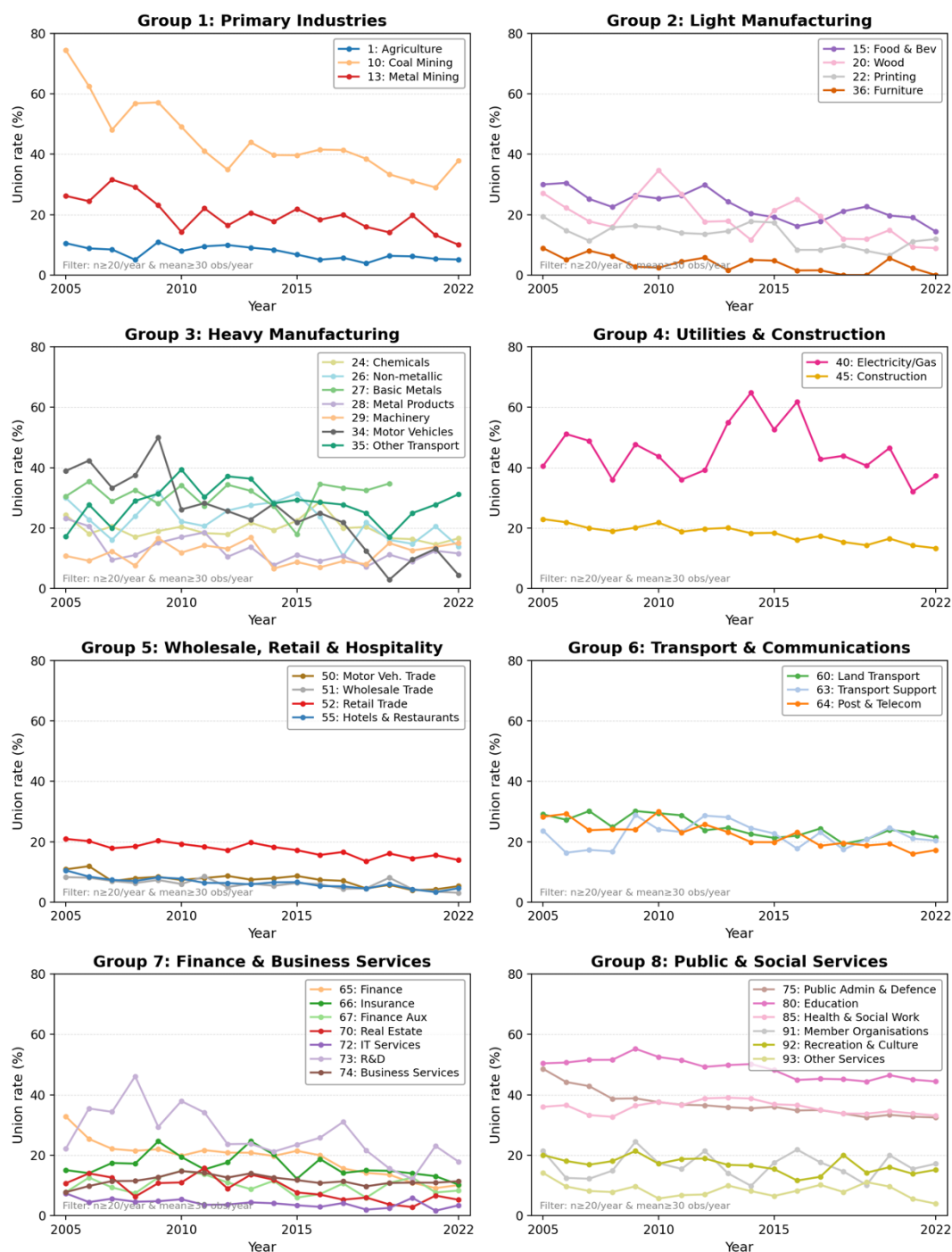
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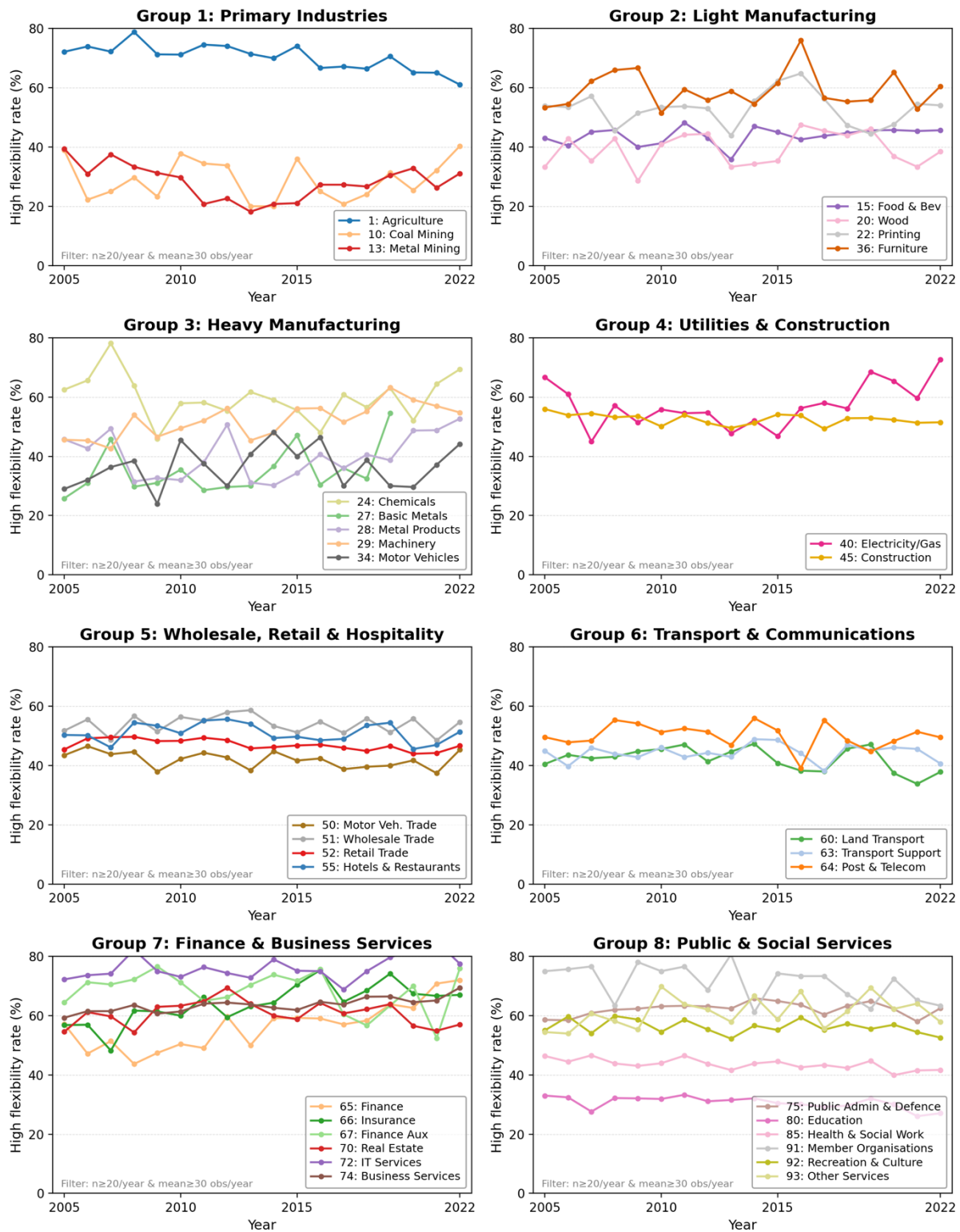
Appendix

Figure A 1 Industry-Level Unionization Rates, 2005–2022



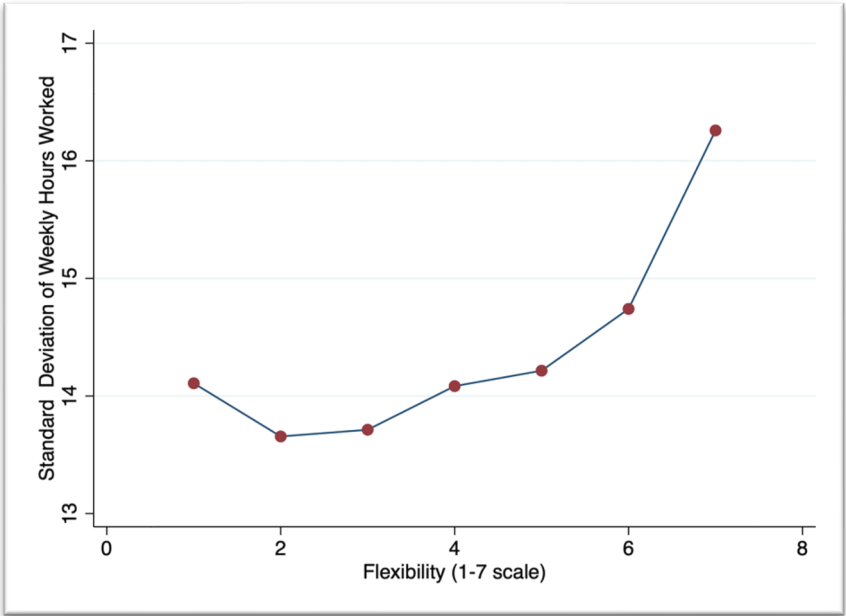
Note: Source: HILDA Survey, Waves 5–22. Each line represents a two-digit ISIC industry. Industries are grouped into eight broad sector categories. Industry-year observations with fewer than 20 respondents and industries with an average of fewer than 30 observations per year are excluded. The figures show substantial variation in unionisation/high-flexibility rates both across industries and within industries over time.

Figure A 2 Industry-Level High Flexibility Rates, 2005–2022



Note: Source: HILDA Survey, Waves 5–22. Each line represents a two-digit ISIC industry. Industries are grouped into eight broad sector categories. Industry-year observations with fewer than 20 respondents and industries with an average of fewer than 30 observations per year are excluded.

Figure A 3 Relationship between Working time Flexibility and Variation in Weekly Hours



Note: Data source is Household, Income, and Labour Dynamics in Australia (HILDA) Survey (Waves 5–22). The horizontal axis reports the working time flexibility scale, ranging from 1 to 7, where higher values indicate greater perceived flexibility. The vertical axis reports the standard deviation of weekly hours worked within each flexibility category.

Table A1 Variables Description

Variables Description		
Variable name	Definition	Measurement
Wage	real hourly wage of main job for individual	log of real hourly wages (current weekly gross wages in main job divided by hours per week usually worked in main job) ⁹
Union membership	whether respondent is a member of a trade	binary: 1 = union member, 0 = non-union
Working time	perceived and formal flexibility in working	Likert scale (strongly disagree to strongly agree: 1-7)
Gender	female or male	binary: 1 = female, 0 = male
Experience	work experience	years since left full-time education
Schooling ¹⁰	years of schooling	imputed education-equivalent years
Partnered	whether respondent has a partner	binary: 1 = yes (including legally married and De facto), 0 = no
Children	number of dependent children aged under 5 (includes partner's children)	Min: 0; Max:4
Unearned income	the part of household income from household assets or spousal earnings	log of individual unearned income (total household income minus the person's earned income)

⁹ We apply a logarithmic transformation to real hourly wages for two key reasons: first, this allows coefficients to be interpreted as percentage changes in wages, facilitating more intuitive economic interpretation; second, the raw wage data exhibits substantial variability (standard deviation of 38), and the logarithmic transformation helps normalise the distribution and reduce the influence of extreme values in the regression analysis. And the main job refers to the job in which a respondent usually works the most hours during their reference week. And Nominal hourly wages were converted to real terms using the OECD annual Consumer Price Index (CPI, base = 2015) to account for inflation over 2005–2022.

¹⁰ Education-equivalent years are imputed from respondents' highest educational qualification following Zawadzki and Wojdyło (2018): 18 years for Masters/Doctorate, 16 for Graduate Diploma/Certificate, 15 for Bachelor/Honours, 13 for Diploma/Advanced Diploma, 12 for Certificate III/IV and Year 12, and 11 for Year 11 or below.

Instrument variables	Lagged Industry unionization	the degree of union presence within an individual's industry	One-year- lagged the ratio of
	Lagged Industry high-	the industry-level prevalence of high-perceived flexibility	One-year- lagged the proportion

Table A 2 Union Membership Patterns by Gender in Australia

Gender	Always	Never	Switched	Total
Male	8.78%	62.88%	28.34%	100%
Female	8.51%	63.22%	28.27%	100%

Note: This table shows the distribution of union membership patterns across individuals' entire observation period. "Always" indicates individuals who are union members in all observed periods (min=1, max=1). "Never" indicates individuals who are never union members in any observed period (min=0, max=0). "Switched" indicates individuals who change union membership status during the observation period (min=0, max=1). The analysis excludes individuals with missing union membership data in all periods.

Table A3 Summary Statistics of Industry-Level Unionization and Flexibility Rates

Statistic	Union rate	Flexibility rate
Mean (across industries)	0.184	0.508
SD (between-industry)	0.122	0.098
SD (within-industry)	0.115	0.120
Min	0	0
Max	1	1

Note: This table reports summary statistics for industry-level unionization rates and flexibility prevalence used as instrumental variables. “Between-industry” standard deviation captures cross-industry heterogeneity, while “within-industry” standard deviation reflects temporal variation within industries over time. The analysis covers 60 industries observed over multiple years. The comparable within- and between-industry variation indicates that the instruments exploit both cross-sectoral and temporal changes rather than relying solely on static industry differences.

Table A4 First-stage Results of IV Estimation (Main Results)

	Men		Women	
	Union	Flexibility	Union	Flexibility
	(1)	(2)	(3)	(4)
Lagged Industry	0.8897***	-0.3245***	0.8329***	-0.3607***
Unionization Rate	(0.0157)	(0.0682)	(0.0214)	(0.0966)
Lagged Industry High- flexibility Rate	-0.0293**	2.6788***	-0.0244	2.8883***
	(0.0115)	(0.0544)	(0.0147)	(0.0757)
Children 0-4	-0.0021	0.0239	-0.0027	0.3132***
	(0.0033)	(0.0155)	(0.0042)	(0.0193)
Experience	0.0123***	0.0017	0.0062***	0.0086***
	(0.0005)	(0.0028)	(0.0006)	(0.0029)
Experience ²	-0.0185***	0.0110**	-0.0061***	-0.0071
	(0.0010)	(0.0050)	(0.0011)	(0.0054)
Partner	0.0025	0.1160***	0.0000	0.1686***
	(0.0047)	(0.0220)	(0.0043)	(0.0200)
Schooling	0.0012	0.0413***	0.0077***	0.0311***
	(0.0007)	(0.0035)	(0.0008)	(0.0036)
Year Effects	Yes	Yes	Yes	Yes
Observations	50,245	50,245	51,408	51,408
F(excluded)	1659.64***	1300.35***	766.21***	758.39***
SW F-statistic	3034.41***	2370.31***	1511.45***	1429.53***

Note: This table reports first-stage regression results for the IV estimation of the direct effect model. The dependent variables are lagged union membership (columns 1 and 3) and lagged working time flexibility (columns 2 and 4) for men and women respectively. IV for union membership is rolling average of industry unionization rate in past 3 years. IV for flexibility is lagged industry high-flexibility rate in past 3 years. All regressions include year fixed effects and control variables as specified in the main text. F(excluded) reports the F-statistic for excluded instruments in each first-stage regression. SW F-statistic is the Sanderson-Windmeijer first-stage F-statistic testing the strength of identification for each endogenous variable. Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A5 Instrumental Variable Diagnostics (Main Results)

	Men	Women
Underidentification test:		
Kleibergen-Paap LM	1868.93***	1188.93***
(p-value)	(0.000)	(0.000)
Weak identification test:		
Cragg-Donald F	1021.87	604.20
Kleibergen-Paap F	1147.88	691.96
Stock-Yogo 10% critical	7.03	7.03
Individual instrument strength:		
First-stage F (Union)	3034.41	1511.45
First-stage F (Flex)	2370.31	1429.53
Anderson-Rubin test:		
F-statistic	36.24**	30.53***
(p-value)	(0.000)	(0.000)

Note: This table reports diagnostic tests for the instrumental variable estimation of the direct effect model. The Kleibergen-Paap LM statistic tests for underidentification (H_0 : model is underidentified). The Cragg-Donald F and Kleibergen-Paap F statistics test for weak identification, with the Stock-Yogo 10% critical value (7.03) providing the benchmark for rejecting weak instruments when there are 2 endogenous variables. First-stage F-statistics test the relevance of instruments for individual endogenous variables (threshold: 10.0). The Anderson-Rubin test provides a test of the joint significance of endogenous regressors that is robust to weak instruments. All test statistics strongly reject underidentification and weak identification concerns, confirming the validity of our instrumental variable approach. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A6 First-stage Results of IV Estimation of Direct Effect (Robustness Test)

	Men		Women	
	Union	Flexibility	Union	Flexibility
	(1)	(2)	(3)	(4)
Three-year Rolling Average of Lagged Industry Unionization Rates	0.823*** (0.017)	-0.148** (0.063)	0.810*** (0.022)	-0.082 (0.090)
Lagged Industry High-flexibility Rate	-0.036** (0.013)	3.378*** (0.046)	0.810*** (0.015)	3.492*** (0.066)
Children 0-4	-0.012* (0.007)	0.039 (0.029)	-0.001 (0.009)	0.249*** (0.037)
Experience	0.015*** (0.001)	-0.037*** (0.006)	0.003** (0.001)	-0.042*** (0.007)
Experience ²	-0.024*** (0.002)	0.064*** (0.009)	-0.002 (0.002)	0.068*** (0.011)
Partner	0.018** (0.008)	0.171*** (0.033)	-0.018*** (0.007)	0.195*** (0.029)
Schooling	-0.001 (0.001)	0.052*** (0.006)	0.008*** (0.002)	0.027*** (0.007)
Year & Industry Effects	Yes	Yes	Yes	Yes
Observations	19,587	19,587	21,893	21,893
F(excluded)	1192.55***	2729.96***	674.65***	1434.05***
SW F-statistic	2304.28***	5270.97***	1332.44***	2790.11***

Note: This table reports first-stage regression results using alternative unionization instrument (three-year rolling average of lagged industry unionization rates). All regressions include year and industry fixed effects and control variables as specified in the main text. Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A7 Robustness Test Instrumental Variable Diagnostics (rolling rate of past 3 years)

	Men	Women
Underidentification test:		
Kleibergen-Paap LM	1436.94***	1032.04***
(p-value)	(0.000)	(0.000)
Weak identification test:		
Cragg-Donald F	1200.61	653.41
Kleibergen-Paap F	1150.05	670.49
Stock-Yogo 10% critical	7.03	7.03
Individual instrument strength:		
First-stage F (Union)	2304.28	1332.44
First-stage F (Flex)	5270.97	2790.11
Anderson-Rubin test:		
F-statistic	16.28***	38.03***
(p-value)	(0.000)	(0.000)

Note: This table reports diagnostic tests using alternative union instrument. The Kleibergen-Paap LM statistic tests for underidentification. The Cragg-Donald F and Kleibergen-Paap F statistics test for weak identification, with the Stock-Yogo 10% critical value (7.03) providing the benchmark for rejecting weak instruments when there are 2 endogenous variables. First-stage F-statistics test the relevance of instruments for individual endogenous variables (threshold: 10.0). The Anderson-Rubin test provides a test of the joint significance of endogenous regressors that is robust to weak instruments. All test statistics strongly reject underidentification and weak identification concerns, confirming the validity of our alternative instrumental variable approach. *** p<0.01, ** p<0.05, * p<0.1.

Table A8 Diagnostic Comparison of Alternative Instruments for Flexibility

IV Type	SW F-stat (Men)	SW F-stat (Women)	R ² w/ Industry & Year FE
Score ≥ 3	2053.34	1104.09	0.5153
Score ≥ 4	2470.72	1296.95	0.4949
Score ≥ 5 (chosen)	2370.31	1429.53	0.4573
Score ≥ 6	1595.80	1222.42	0.3893
Mean Flexibility	2936.58	1741.40	0.5153

Note: All models include industry and year fixed effects. R² is from regressing each IV on industry & year FE; lower R² indicates less collinearity and more usable within-industry variation.

Table A9 Bootstrap tests of differences between direct and total union effects (IV estimates)

Test	Δ (Direct – Total)	SE	p-value	95%CI
Panel A: Difference between Direct and Total Effects				
Male	0.041	0.014	0.003***	[0.012, 0.067]
Female	0.047	0.016	0.004***	[0.012, 0.077]
Panel B: Change in Gendered Difference				
	0.006	0.020	0.764	[-0.032, 0.043]

Note: This table reports bootstrap-based tests of differences between the estimated union coefficients across model specifications. Panel A examines whether the union effect differs between specifications that include and exclude working time flexibility, estimated separately for men and women. Panel B evaluates whether the gender difference in union effects changes across these specifications. The reported difference is defined as $\Delta = \beta_{\text{Direct}} - \beta_{\text{Total}}$. All estimates are based on IV regressions using the same estimation sample. Standard errors and confidence intervals are computed from 300 bootstrap replications clustered at the industry level based on the same estimation sample. All underlying regressions include the full set of controls, industry and year fixed effects, and robust standard errors. Delta-method and SUEST tests yield similar conclusions that the difference is not statistically significant.