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The Premium for Part-time Work in Australia

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Abstract

We use fixed effects and difference-in-differences methodologies to investigate the nature of Australia's part-time wage premium, a phenomenon not observed in other countries. Salary sacrifice and non-cash benefits, previously unexplored explanations, are eliminated. The premium is not explained by occupation and it is observed for people with only one change of employment status and for those with multiple changes. We find that changing from full-time to part-time work with the same employer results in a large and sustained increase in the hourly wage, whereas a temporary decrease in the hourly wage accompanies a change from part-time to full-time work with the same employer. Notably, we find no significant wage change when a move between full-time and part-time work is accompanied by a change of employer.

Keywords: Part-time employment, Wage differentials, Labour supply

JEL classifications: J31, J32, J33

1. Introduction

According to the Australian Bureau of Statistics (ABS, 6202.0) approximately 30 per cent of employed Australians currently work fewer than 35 hours per week, nearly double the rate three decades ago. The trend towards part-time work is not unique to Australia, with Chile, Greece, Ireland, Italy and Turkey recently experiencing increasing proportions of employed persons working part-time. Even so, Australia has the third highest rate of part-time employment among the 34 countries in the Organization for Economic Cooperation and Development (OECD), behind only the Netherlands and Switzerland.[1]

[1] In each year since 2009, approximately 37 per cent of workers in the Netherlands were employed for fewer than 30 hours per week. In Switzerland the figure was 26 per cent, in Australia 25 per cent and in the United Kingdom 24-25 per cent. All statistics quoted in this paragraph appear in OECD, 2012. Kishi (2004, p.234) points to the increasing incidence of part-time work in Japan.

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In most OECD countries, including Australia, part-time workers are predominantly women aged 25-54 years. Compared with the OECD average, Australia has slightly smaller proportions of part-time workers who are older than 54 years and who are men aged 25-54, and a slightly larger proportion of part-time workers who are younger than 25 years (OECD, 2010, p.264). Part-time workers in Australia are less unionised than full-time workers to similar extent as in most OECD countries (OECD, 2010, p.223). As in Britain (Manning and Petrongolo, 2008; Connolly and Gregory, 2009) and other OECD countries (Bardasi and Gornick, 2008), in Australia, part-time employees commonly have low paying occupations, such as services and sales, while full-time employees are more typically found in high paying occupations such as managers and professionals. As with occupation, part-time employees in Australia tend to be concentrated in low paying industries, such as retail trade, accommodation and food services, while full-time employees are typically concentrated in high paying industries like manufacturing and mining (ABS, 3659.0).

There is concern in Australia about the high rate of part-time employment because part-time jobs are perceived as paying lower wages, having irregular work schedules, lower job security and fewer career opportunities than full-time jobs (Burgess, 2005; Australian Government, 2003). Whereas in most OECD countries part-timers are entitled to receive the same contractual pay and working conditions as equivalent full-time workers (on a pro-rata basis), in Australia there is no specific statutory requirement to treat part-time and full-time workers equally. All permanent employees in Australia have the same safety net of minimum entitlements for wages, leave, dismissal protection, etc. (OECD, 2010, p.218), but approximately 50 per cent of Australian part-time workers are employed on a casual, rather than permanent, basis. In contrast, less than ten per cent of full-time workers are casual. By definition, casual employees do not have access to paid annual leave or paid sick leave (ABS, 6105.0) although they do receive an hourly wage loading in compensation, and many casuals do not receive paid public holidays, notice of dismissal or redundancy pay (Campbell, 2004).

Approximately half of OECD countries require employers to notify part-time employees who want to work longer hours of full-time vacancies when they arise. Some also require employers to give existing underemployed part-time workers preferential treatment when filling full-time vacancies. Few such rights exist in Australia. Full-time workers who are parents of school-aged children have the right to request part-time work in most OECD countries and employers can only refuse such requests on serious business or operational grounds. Rights to work part-time for non-parents are less common but several countries give specific rights to part-time work to carers of adults, workers who are sick or disabled, those pursuing education or training and older workers. Australia has been one of the least generous of the OECD countries when it comes to the right to request part-time work (OECD, 2010, pp.217-219).

The average hourly wage of part-time workers is indeed lower than that of full-time workers, both in Australia and in many other countries (OECD, 2010, p.221). However, when the attributes of workers and jobs are taken into account, a large, 'adjusted' part-time wage premium is observed in Australia (Booth and Wood, 2008), a result that is at odds with what is observed in the United Kingdom (Connolly and Gregory, 2009) and, to a lesser extent, in the United States (Hirsch, 2005). The objective of this study is to explore the nature of that premium. This is important not only because a substantial, and increasing, number of Australians are working part time, but also because a premium (or a penalty) for part-time work suggests a lack of competition in the labour market and has implications for productivity.

We make four contributions to what is known about the phenomenon. First, we find that the premium cannot be explained by the level of non-cash benefits and salary sacrifice received by full-time and part-time employees. Second, we find that the premium is present across occupations. Third, the premium exists for those who change their employment status multiple times as well as for those with only one such change. Finally, having constructed a time path of wage differentials up to two years after a move between part-time and full-time work, we find the premium persists only for those who remain with the same employer.

2. Economic Theory and Australian Labour-Market Institutions

Economic Theory

The textbook explanation of a part-time wage penalty given homogenous jobs and workers, is the existence of quasi-fixed costs, such as the costs of recruiting and training new employees (Hyclak *et al*. 2005, p.58). The effect of quasi-fixed costs is compounded if part-time workers have higher turnover rates than full-time workers (Casey, *et al*. 1997). A part-time wage penalty will also occur if there is an excess supply of part-time workers who are unwilling or unable to move between part-time and full-time jobs and if employers do not view part-time and full-time workers as perfect substitutes. Many part-time workers have family responsibilities and, therefore, have a high opportunity cost of time spent in employment. The flexible working conditions of part-time jobs make them attractive, in which case the theory of compensating wage differentials suggests there will be a part-time wage penalty. Part-time workers might limit their job search to areas close to home in which case employers with monopsony power in local labour markets are able to offer them hourly wages that are lower than the full-time wage (Ermisch and Wright, 1993, pp.113-115).

On the other hand, a part-time wage premium could be explained by non-cash benefits that form part of total remuneration. Some non-cash benefits constitute a quasi-fixed cost, in which case part-time workers are less likely to receive them and a wage premium may be paid in compensation (Blank, 1990; Montgomery and Cosgrove, 1995). Alternatively, in businesses that face periods of fluctuating demand, part-time workers who are employed to work during peak times only will be more productive than identical full-time workers who would be idle for much of the time. Furthermore, it is possible that part-time workers will be more productive than full-time workers in jobs where productivity declines rapidly with long hours of work (Booth and Ravallion, 1993).

Australian Labour Market Institutions

Unlike other countries, Australia's industrial relations system provides a number of minimum wages in 'award agreements' covering various occupations and industries.[2] The minimum wages set in Australia appear to be among the highest in the world (Waring and Burgess, 2011; Whiteford, 2013, p.26) leading to a wage distribution that has historically been more compressed than in countries such as the United States and the United Kingdom (Gornick and Jacobs, 1996). Although the wage distribution in Australia has widened since the 1990s, in 2005 it was the 13th most compressed of the 28 countries in the OECD (Whiteford, 2013, pp.26-27). This relatively low level of overall wage dispersion implies that any part-time/full-time wage differential in Australia will be smaller than those found elsewhere. Yet the evidence from Booth and Wood (2008) suggests it is larger.

The number of Australian workers hired on a casual basis has been growing, and currently they account for approximately 23 per cent of employed people (ABS, 6359.0). Casual workers do not have access to paid annual leave or paid sick leave and they also lack other rights, such as notice of dismissal. Industrial law requires casual workers be paid a 25 per cent loading on the relevant minimum wage in compensation for lack of such entitlements (Campbell, 2004, p.93).[3] Of course, workers – casual or otherwise – can be paid a wage rate above the legal minimum, in which case the effect of the loading will be reduced and may be eliminated altogether. Nevertheless casual employment provides a reason to expect a part-time wage premium because a larger percentage of part-time than full-time workers are employed on a casual basis and therefore qualify for the casual loading (ABS, 6105). Consequently, casual status should be held constant when estimating the part-time/full-time wage differential.

High effective marginal tax rates for a second income earner in a family also provide a reason to expect a part-time wage premium in Australia. If a second income is earned, a family may no longer be eligible for certain tax exemptions and social programs such as the Family Tax Benefit, thereby increasing the effective tax rate. Employers who want to hire part-time workers will need to pay them a premium to compensate for the high effective marginal tax rate (Booth and Wood, 2008). Therefore, a control for marital status should be used when estimating part-time/full-time wage differentials.

In Australia, certain industrial awards specify that all workers in a given industry be paid penalty rates for working outside of normal hours, such as evenings, weekends and public holidays. Both part-time and full-time workers can receive penalty rates, but if they are more common for one group than the other, this will give rise to a part-time/full-time wage differential. Hence, there is a need to control for working a standard versus nonstandard work schedule in estimating part-time/full-time wage differentials.

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[2] Minimum wages have remained a key feature of wage-setting procedures in Australia despite changes to the industrial relations system over the past decade, the most notable being the introduction of the Work Choices legislation in 2006, which attempted to decentralise wage bargaining and reduce award coverage, and its replacement in 2009 with the Fair Work Act, which at least partially reversed some of these changes.

[3] See Fair Work Australia, Miscellaneous Award 2010 (s.10.4).

Unpaid overtime is common in Australia (Wooden, 2001) and is another factor that might contribute to a part-time wage premium. Salaried individuals may work long hours at certain stages of their careers, on the expectation of being rewarded by promotion later on. High levels of unpaid overtime may also be worked by individuals who simply enjoy their jobs (Hirsch, 2005). Unpaid overtime is likely to be more common among full-time workers because many part-time workers intentionally choose to work short hours so they can engage in family-related activities. Indeed, a survey conducted by The Australia Institute found that, on average, full-time employees work 70 minutes of unpaid overtime per day while part-time employees work 23 minutes per (working) day (Fear and Denniss, 2009). Individuals who work long hours each week for a constant weekly salary will have a calculated hourly wage (usual weekly wages and salaries divided by hours usually worked per week) that is lower than the hourly wage in their employment contracts, so with unpaid overtime more common among full-time than among part-time workers a part-time wage premium will be observed. Furthermore, people employed on part-time contracts who regularly work a lot of overtime, are likely to be classified as full-time based on the hours they report usually working per week. This is particularly important for Australian studies since the upward trend in unpaid overtime appears to be more widespread in Australia than in other OECD countries (Campbell, 2007).

In summary, both economic theory and Australian labour market institutions offer several explanations for the existence of a part-time/full-time wage differential but neither unambiguously predicts either a part-time wage penalty or a part-time wage premium.

3. Literature Review

Cross-Sectional Studies

Early cross-sectional studies using United States data, such as Owen (1978) and Long and Jones (1981), found adjusted part-time wage penalties in excess of ten per cent after controlling for various worker and job characteristics, but not for self-selection into part-time work. Later studies that corrected for self-selection bias using a Heckman-type correction procedure (Heckman, 1979) typically also found significant part-time wage penalties (Simpson, 1986; Main, 1988; Ermish and Wright, 1993). An exception was Blank (1990), who controlled for selection into part-time work and found a small part-time wage premium for women.

Rodgers (2004) estimated part-time/full-time wage differentials for Australian men and women. After controlling for worker and job attributes, and self-selection into non-employment, part-time and full-time employment, the author found part-time wage premiums of nine and three per cent for women and men, respectively. However, neither premium was statistically significant.

There have also been studies of specific occupations. Most found part-time wage penalties (Montgomery and Cosgrove, 1995; Lettau, 1997; Manning and Petrongolo, 2008), but one found a small part-time wage premium (Hirsch and Schumacher, 1995). The three studies from the 1990s also investigated the receipt of fringe benefits, such as medical insurance, by full-time and part-time workers and found that part-time workers had lower total compensation than equivalent full-time workers.

There have been two cross-national comparisons of part-time/full-time wage differentials, both based on cross-sectional data from the Luxemburg Income Survey (LIS). Gornick and Jacobs (1996) found part-time wage penalties for both genders in the United States, Canada and Australia, and for women, but not men, in the United Kingdom. Bardasi and Gornick (2008) found part-time wage penalties for women in Canada, Germany, Italy, the United Kingdom and the United States but not Sweden.

Longitudinal Studies

Recent studies have used longitudinal data to control for unobservable worker characteristics. They have found small, or non-significant, part-time wage differentials.

In the United States, Hirsh (2005) used panel data from the Current Population Survey Outgoing Rotation Group and found a part-time wage premium of less than two per cent for both men and women. He also examined possible asymmetries of moving between full-time and part-time work. For women, he found a premium for changing from full-time to part-time work that was larger in magnitude than the penalty suffered by changing from part-time to full-time work, while the reverse was true for men. On the other hand, Hirsch found part-time wage penalties of around five per cent for workers changing occupation and industry at the same time as changing their part-time/full-time status. Unlike British studies that focus on the lack of availability of part-time work in highly paid industries and occupations, he suggests that the wage differential probably reflects workers moving into and out of career jobs (Hirsch, 2005, p.542).

In the United Kingdom, Connolly and Gregory (2009) used the New Earnings Panel to estimate part-time/full-time wage differentials for women and to assess their respective wage growth over time. With experience and job tenure as the only controls, the authors found a part-time wage penalty of 11 per cent but this fell to just two per cent when controls for work experience and tenure were broken down, first into their part-time and full-time components, and then into high, medium and low skilled occupations. The authors also found that a move from full-time to part-time work was associated with an immediate wage penalty of seven per cent, which was brought about by an occupational downgrading and/or a change of employer, rather than by the part-time nature of the job. The penalty continued for four years, but became smaller each year. On the other hand, women switching from part-time to full-time employment earned a premium, but only after the first year of full-time employment.

Booth and Wood (2008) used longitudinal data from the first four waves of the Household, Income and Labour Dynamics in Australia (HILDA) Survey to control for time-invariant, unobserved heterogeneity. They found that part-time women who were not on casual contracts earned approximately ten per cent more per hour than their full-time counterparts. The premium was 15 per cent for women who were part-time and casual. In contrast, part-time, non-casual men earned a 15 per cent premium over comparable full-time men, but the premium was ten per cent for men who were part-time and casual. Therefore, part-time wage premiums appear to be much higher for both men and women in Australia than in the United States and the United Kingdom. The authors conducted several robustness tests and explored the potential for asymmetric wage effects of moving between full-time and part-time employment, all of which were consistent with their primary results. With a 12-year panel of HILDA data now available, we further explore the magnitude and nature of the part-time premium in the remainder of this paper.

4. The Models

First, we replicate Booth and Wood's (2008) results using a reduced-form model, which is estimated separately for men and women using pooled OLS and fixed-effects, both with clustered standard errors, which are robust to both heteroskedasticity and correlation over time for the same individual. Model 1 is:

$$\log(wage_{i,t}) = \beta_1 P_{i,t} + \beta_2 C_{i,t} + \beta_3 (P^*C)_{i,t} + \mathbf{X}_{i,t}\boldsymbol{\gamma} + \mu_i + \varepsilon_{i,t} \tag{1}$$

where $P_{i,t}$ is a binary variable equal to one if person i is employed part-time in year t; $C_{i,t}$ equals one if person i is employed on a casual contract in year t (zero otherwise); $(P^*C)_{i,t}$ is the interaction between part-time and casual status; $\mathbf{X}_{\mathbf{i,t}}$ is a set of observed characteristics of employee i and his or her job in year t (including an intercept); μ_i represents unobserved time-invariant attributes of employees; and $\varepsilon_{i,t}$ is an idiosyncratic error. The parameters of interest are β_1, which is the average log-wage differential between part-time, non-casual workers and full-time, non-casual workers, and $\beta_1 + \beta_3$, which is the average log-wage differential between part-time, casual workers and full-time, casual workers. Like Booth and Wood (2008), we do not model selection into employment with a separate equation. However, the fixed-effects estimation procedure will eliminate selection bias that operates purely through the individual-specific error, μ_i. In other words, if employed people have time-invariant characteristics that make them different from people who are not employed then the FE estimators will be consistent.[4]

Model 1 assumes that the effect on wages of changing into, and out of, part-time jobs is symmetric, as is the effect of changing into, and out of, casual jobs. These assumptions are relaxed in Model 2 which, like Model 1, controls for time-invariant, unobserved heterogeneity among employees.

$$\begin{aligned}\Delta\log(wage_{i,t\text{-}1/t+k}) = {} & \beta_1^a Pjoiner_{i,t-1/t} + \beta_2^a newemp_{i,t-1/t+k} + \beta_3^a (Pjoiner_{i,t-1/t} * newemp_{i,t-1/t+k}) + \\ & \beta_4^a Cstayer_{i,t-1/t+k} + \beta_5^a Cleaver_{i,t-1/t+k} + \beta_6^a Cjoiner_{i,t-1/t+k} + \mathbf{\Delta X}_{i,t-1/t+k}\boldsymbol{\gamma}^a + \Delta\varepsilon^a_{i,t-1/t+k}\end{aligned} \tag{2a}$$

$$\begin{aligned}\Delta\log(wage_{i,t\text{-}1/t+k}) = {} & \beta_1^b Pleaver_{i,t-1/t} + \beta_2^b newemp_{i,t-1/t+k} + \beta_3^b (Pleaver_{i,t-1/t} * newemp_{i,t-1/t+k}) + \\ & \beta_4^b Cstayer_{i,t-1/t+k} + \beta_5^b Cleaver_{i,t-1/t+k} + \beta_6^b Cjoiner_{i,t-1/t+k} + \mathbf{\Delta X}_{i,t-1/t+k}\boldsymbol{\gamma}^b + \Delta\varepsilon^b_{i,t-1/t+k}\end{aligned} \tag{2b}$$

Equation (2a) is estimated using only those observations where the individual works full time in year t-1 (full-time starters) and Equation (2b) is estimated using only those observations where the individual works part-time in year t-1 (part-time starters).

[4] Fixed-effects estimation does not eliminate all forms of selection bias but using a Heckman-like correction to Model 1 would require (strictly exogenous) instrumental variables, which are difficult to find. Weak instruments introduce problems of their own. (See, Vella, 1998, pp.156-157).

Full-time starters are divided into full-time stayers, who continue to work full-time (the control group), and part-time joiners, who switch to part-time work in year t (the treatment group, *Pjoiner*). Part-time starters are divided into part-time stayers, who continue to work part-time (the control group), and part-time leavers, who switch to full-time work in year t (the treatment group, *Pleaver)*.[5] Model 2 also contains controls for changes in casual status that occur simultaneously with the wage change. Casual stayers (*Cstayer*) are casual in both years t-1 and t+k; casual leavers (*Cleaver*) are casual in year t-1 and non-casual in year t+k; casual joiners (*Cjoiner*) are non-casual in year t-1 and casual in year t+k. The excluded category consists of people who are non-casual in both periods. Changes from year t-1 to year t+k in other control variables, $\boldsymbol{\Delta X}$, are also included.

Model 2 is used to investigate whether the effect of changing to, or from, part-time work is affected by a concurrent change of employer (*newemp*). Equations (2a) and (2b) are estimated with the interaction parameter, β_3, set to zero, and also with no constraint on β_3 thereby allowing the effect of changing employment status to be different for those who simultaneously change employers and those who do not. k = 0, 1, 2 …. traces out the effect on wages several years after changing between full-time and part-time employment.

5. The Data

The models are estimated using a 12-year panel of unit-record data from Release 12 of the HILDA Survey, which is a representative sample of people living in Australian households. (See, Summerfield *et al*. 2012, pp.116-118 for details). People aged 15 years and older have been interviewed each year and data on family and household characteristics, income and work have been recorded. These responding people constitute the unbalanced panel used in this study.[6]

The HILDA data have several major advantages for this study compared with other Australian data sets. The data allow an estimate of each employee's usual hourly wage in his or her main job, which is the job from which the most pay is usually received each week. Usual, weekly, gross wages and salary in the main job, adjusted for inflation using the consumer price index, was divided by hours per week usually worked in the main job. Hours and wages and salary are recorded as continuous variables. The use of usual gross earnings and usual hours of work is preferred to gross earnings and hours of work in the week prior to the interview, which may be an atypical week for the worker. As in previous Australian studies, in this paper a part-time worker is defined as someone who usually works less than 35 hours per week in his or her main job. We use hours of work in the main job, rather than in all jobs combined, because we are interested in the nature of part-time jobs.

The HILDA data set provides a considerable amount of demographic data on employees, such as age, sex, marital status, education and geographic location. There are also data on the attributes of respondents' jobs, such as casual status, type of work schedule, occupation, industry, workplace size and whether in the public or private sector.

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[5] Restrictions concerning other changes in employment status are imposed on the sample when Model 2 is estimated.

[6] A detailed discussion of the HILDA survey can be found in Wooden and Watson (2007).

The major disadvantage of the HILDA data for the purpose of this study is that hours worked includes paid and unpaid overtime, but unpaid overtime hours are not separately recorded. Although our calculated wage reflects the 'true' hourly payment for what is actually worked, the inclusion of unpaid overtime in hours worked will result in a calculated hourly wage that is lower than the wage stated in the employment contract. The second deficiency of the HILDA data is that work experience, which is reported as years in paid work, makes no distinction between full-time and part-time work experience. The same applies to job tenure. This is likely to lead to an understatement of the full-time/part-time wage differential and would be of concern if a part-time penalty had been found, as its size would have been overstated. But in Australia a part-time premium has been found, and it is likely to be an understatement.

6. Descriptive Statistics

To verify Booth and Wood's (2008) results with the 12-year panel we applied the same exclusions: individuals must be employees, aged 18 to 64 years, not full-time students, not in the armed forces, farming or fishing; they must work less than 100 hours per week and earn an hourly wage between one and 160 dollars.[7] Individuals must also have valid data on the control variables, details of which are given below. To estimate Model 1 we used a sample consisting of people who satisfy these conditions in at least two consecutive waves of HILDA data, thereby ensuring that the wage differentials in the data are not due to any deterioration in human capital associated with a substantial break in employment. Model 2 was estimated using a sample comprised of people who satisfy these conditions in at least four consecutive waves.

Five specifications of Model 1, the same ones reported by Booth and Wood (2008), were estimated. In the first specification $X_{i,t}$ includes only dummy variables for (all but one of) the years covered by the longitudinal data and provides an estimate of the raw part-time/full-time wage differentials for men and women. Specification 2 also includes binary variables for holding a single job, whether the main job involves a regular day-time schedule, is a fixed-term contract, and was acquired through a labour-hire firm. Also included are the individual's country of birth, age, marital status and geographic location. Education and quadratics for job tenure and employment experience are added to form Specification 3. Workplace size, union membership, public/private sector, and industry are appended to form Specification 4 and occupation is added to form Specification 5. After deleting observations with missing data on one or more of these control variables, the sample used to estimate all specifications of Booth and Wood's model contains 5,319 men with 29,199 observations (2,494 part-time and 26,705 full-time) and 5,497 women with 28,937 observations (12,287 part-time and 16,650 full-time).[8]

[7] Very few people in the data set usually work 100 or more hours per week in their main job. These outliers would be likely to cause errors in the calculation of the hourly wage. The majority of people with hourly wages greater than $160 only earned these high wages in one wave, worked less than five hours work per week and/or had low levels of education. Many individuals earning less than one dollar per hour recorded high weekly hours and/or low weekly wages.

[8] There were 37,184 and 38,021 person-year observations for males and females, respectively, after imposing the same exclusions as Booth and Wood (2008). These were reduced to 34,734 and 35,510 as a result of requiring valid observations on all control variables. The requirement that there be at least two consecutive observations per person produced the final sample of 29,199 male-year observations and 28,937 female-year observations.

Descriptive statistics underlying Model 1 are presented in table 1. They show that average hourly wages are lower for part-time men and women than for their full-time counterparts but the differential is statistically significant for men only. We also see that part-time and full-time employees are systematically different, suggesting that at least part of the 'raw' part-time wage penalty can be attributed to differences in the observable (productivity-related) characteristics of the two groups and the attributes of the jobs they hold. For example, part-time men and women are more likely to be employed on a casual basis, less likely to hold only one job, less likely to work a regular day-time schedule and less likely to be employed on a fixed-term contract than their full-time counterparts.

Table1 - Descriptive Statistics for Full-time and Part-time Men and Women, Model 1

	Men		*Women*	
Variable	*Full-time*	*Part-time*	*Full-time*	*Part-time*
Hourly wage ($)	31.75	27.98 ***	27.21	26.76
Usual hours per week	44.73	22.42 ***	41.76	21.45 ***
Casual[a]	0.08	0.57 ***	0.06	0.39 ***
One job only[a]	0.95	0.79 ***	0.95	0.85 ***
Regular day-time schedule[a]	0.80	0.59 ***	0.85	0.73 ***
Fixed-term contract[a]	0.09	0.07 *	0.11	0.08 ***
Employed via labour hire firm[a]	0.03	0.05 **	0.02	0.03
Age (years)	39.29	39.81	39.25	42.00 ***
Married[a]	0.56	0.43 ***	0.46	0.64 ***
Cohabitating[a]	0.13	0.10 **	0.16	0.10 ***
Widowed, divorced or separated[a]	0.06	0.06	0.14	0.11 *
Never married[a]	0.25	0.41 ***	0.25	0.14 ***
Urban location[a]	0.72	0.69	0.74	0.65 ***
Inner regional location[a]	0.19	0.22	0.17	0.24 ***
Outer regional location[a]	0.08	0.08	0.08	0.09
Remote location[a]	0.01	0.01	0.01	0.01
Australian born[a]	0.75	0.76	0.74	0.79 **
Born in other English speaking country[a]	0.11	0.09	0.10	0.09
Born in a non-English speaking country[a]	0.14	0.15	0.16	0.12 **
Tenure with current employer (yrs)	7.68	5.01 ***	7.19	6.43 **
Experience (yrs)	20.60	19.73	18.22	18.98 *
Postgraduate degree[a]	0.06	0.04	0.05	0.03 ***
Graduate diploma/certificate[a]	0.06	0.07	0.09	0.07
Bachelor degree[a]	0.16	0.15	0.23	0.16 ***
Advanced diploma/diploma[a]	0.09	0.08	0.12	0.10 *
Certificate III or IV[a]	0.29	0.19 ***	0.14	0.18 ***
Year 12[a]	0.16	0.22 **	0.17	0.16
<=Year 11 or unknown[a]	0.19	0.26 **	0.20	0.31 ***
Trade union member[a]	0.33	0.26 **	0.33	0.28 ***
Public sector job[a]	0.23	0.23	0.35	0.31 **
Firm size <20[a]	0.31	0.41 ***	0.26	0.40 ***
Firm size 20-99[a]	0.31	0.31	0.34	0.32
Firm size 100-499[a]	0.24	0.18 ***	0.23	0.16 ***
Firm size >=500[a]	0.15	0.09 ***	0.17	0.11 ***
Person-year observations	26,705	2,494	16,650	12,287
Number of people	5,319		5,497	

Source: HILDA, Waves 1-12.
*, ** and *** indicate a difference that is significantly different from zero at the 5%, 1% and 0.1% levels.
Notes: [a] Measured as a proportion. [b] Means are weighted by cross sectional probability weights.

To identify the effect of employment status on wage rates using longitudinal data there must be a reasonable number of changes in status between adjacent years. Table 2 presents transition matrices for Model 1, men and women. The off-diagonal entries show the number of observations involving a change of status. Compared with the data used by Booth and Wood (2008), our data show many more transitions. Notably, where there is no distinction between moving into a particular employment state and moving out of it, the smallest number of transitions in our data set is (24+28=) 52 for men, and (53+48=) 101 for women, who move between part-time-non-casual and full-time-casual employment. The corresponding numbers in Booth and Wood's (2008) study were 17 and 24.

Table 2 - Transition Matrices Showing Changes of Employment Status, Model 1

	Men			
Year t-1 ↓ / *Year t* →	*Part-time & casual*	*Part-time & non-casual*	*Full-time & casual*	*Full-time & non-casual*
Part-time & casual	623	74	192	226
Part-time & non-casual	57	506	28	237
Full-time & casual	158	24	854	608
Full-time & non-casual	163	226	450	19454
	Women			
Year t-1 ↓ / *Year t* →	*Part-time & casual*	*Part-time & non-casual*	*Full-time & casual*	*Full-time & non-casual*
Part-time & casual	2610	568	227	393
Part-time & non-casual	315	4883	48	874
Full-time & casual	173	53	317	284
Full-time & non-casual	233	861	186	11415

Source: HILDA, Waves 1-12.
Note: Numbers are unweighted counts of transitions, based on pairs of consecutive waves.

7. Results

Model 1

Part-time/full-time wage differentials in each of the five specifications of Model 1 are presented in table 3. Full results for Specification 3, the specification preferred by Booth and Wood (2008), appear in appendix A1. The Breusch-Pagan Lagrange multiplier test leads to the rejection of the hypothesis that $\mu_i = 0$ in Model 1, indicating that either random effects or fixed effects is preferred to pooled OLS. The Hausman test indicates that fixed effects is preferred to random effects in all 5 specifications. Consequently, the discussion below is focussed on the fixed-effects estimates.

The results from the fixed-effects model are remarkably stable across the five specifications and indicate large part-time wage premiums for both men and women. Specification 3 indicates that there is a statistically significant part-time wage premium of 12.5 per cent for men employed on casual contracts and 17.0 per cent for men who are non-casuals. For women, there is a part-time wage premium of 13.2

per cent for casuals and 10.3 per cent for non-casuals, both of which are statistically significant. These results are very similar to those found by Booth and Wood (2008). Our fixed-effects coefficients on the P*C interactions in Model 1 have the same sign (negative for men and positive for women) as those of Booth and Wood (2008) but are not statistically significant at the five per cent level. In this respect our results differ from those of Booth and Wood (2008) who report t-statistics for their interactions that are statistically significant at the five per cent level, whereas ours have p-values of 0.089 for men and 0.101 for women.[9] Therefore, we conclude that the part-time wage premiums for casuals and non-casuals are not significantly different, neither for men nor for women. Adding industry and occupation dummies has little effect on the fixed-effects coefficients. To explore the nature of the premium produced by fixed effects, we conduct some sensitivity tests on Specification 3.

Table 3 - Part-time/Full-time Wage Differentials, Model 1

		Men		*Women*	
Spec.	*Casual Status*	*Pooled OLS*	*Fixed Effects*	*Pooled OLS*	*Fixed Effects*
1	Non-casual	-0.155 **	0.155 ***	-0.009	0.103 ***
		n.s.	*	n.s.	n.s.
	Casual	-0.085 **	0.101 ***	0.026	0.124 ***
2	Non-casual	-0.131 *	0.159 ***	-0.045 **	0.100 ***
		n.s.	n.s.	*	n.s.
	Casual	-0.072 **	0.113 ***	0.019	0.129 ***
3	Non-casual	-0.116 *	0.170 ***	-0.001	0.103 ***
		n.s.	n.s.	n.s.	n.s.
	Casual	-0.077 **	0.125 ***	0.029	0.132 ***
4	Non-casual	-0.046	0.171 ***	0.022 *	0.107 ***
		n.s.	n.s.	n.s.	n.s.
	Casual	-0.015	0.130 ***	0.042	0.139 ***
5	Non-casual	-0.011	0.173 ***	0.044 ***	0.110 ***
		n.s.	n.s.	n.s.	n.s.
	Casual	-0.002	0.132 ***	0.047 *	0.140 ***

Source: HILDA, Waves 1-12.
*, ** and *** beside a coefficient indicates significantly different from zero at the 5%, 1% and 0.1% levels. Asterisks or n.s. on the line between coefficients indicates a significant difference, or no significant difference (n.s.), between the PT/FT wage differential for non-casual and casual employees.
Note: Standard errors are cluster robust.

Sensitivity Tests on Model 1, Specification 3

Overall, the part-time wage premium is robust to the sensitivity tests we performed. First, we graphed the average hourly wage against weekly hours of work and found

[9] Aside from the larger size of our sample, a possible explanation is that we used the ABS definition of casual status with the survey question being 'Does your employer provide you with paid holiday/ sick leave?' whereas Booth and Wood (2008) used a self-identification measure based on the survey question 'Looking at (showcard) which of the following best describes your current contract of employment? Employed on a fixed term contract; employed on a casual basis; employed on a permanent or ongoing basis; other.'

that the highest wage rates, for both men and women occur at fewer than ten hours per week, while the lowest wage rates are for people working more than 70 hours per week. To test whether the observed part-time premiums are being driven by these extreme observations, the fixed-effects model was re-estimated after dropping 1,694 observations with weekly hours of work below ten or above 70. Another 210 observations with an hourly wage less than five dollars were also dropped as they seem unrealistically low. The results, which appear in Panel A of table 4, indicate that outliers explain some of the part-time premium, particularly for men and women on casual contracts. When the outliers are excluded there is a 14.8 per cent part-time wage premium for non-casual men and a significantly smaller 7.8 per cent part-time wage premium for casual men. For non-casual women the part-time wage premium is 9.8 per cent, while for casual women the premium is 8.4 per cent, which are not significantly different. Additional sensitivity tests (discussed below) also exclude the 1,904 outliers from the sample.

The second test indicates that the part-time wage premium is not sensitive to the 35-hour cut-off. Specification 3 of the fixed-effects model was re-estimated with the cut-off set at the OECD definition of part-time employment, 30 hours per week (see Panel B), and at the definition consistent with most Australian awards, 38 hours per week (Panel C). Significant part-time premiums are observed at both cut-offs, for men and women, casual and non-casual.

Next, we found that measurement errors in hours of work are not responsible for the part-time wage premium. Weekly work hours recorded in the HILDA data are clustered at 38 and at the end of five hourly intervals, which is suggestive of measurement errors. With the hourly wage calculated as usual weekly gross wages and salary in the main job divided by hours per week usually worked in the main job, an error in weekly hours worked, random or otherwise, will lead to an error in the calculated hourly wage, and possibly to an error in part-time status. If hours of work are overstated (understated), then the hourly wage will be understated (overstated) and the employee might be classified as full (part) time. We tested whether our results are driven by errors in recorded hours worked per week by re-estimating the model after dropping the 641 observations that involve a change of employment status resulting from a change in weekly work hours that is five or less in absolute value. This eliminates small changes in hours worked that are likely to result in a change of employment status. The impact is seen in Panel D of table 4; the part-time wage premiums remain for men and women, casual and non-casual, and they are slightly larger than those in Panel A.

We also found that the part-time wage premium is fairly constant across age groups for women, but for men the premium is somewhat related to age and casual status. Men tend to work part time when they are young or old, whereas women tend to work part-time when they are middle aged. To test whether the part-time wage premium is concentrated in particular age groups we re-estimated Specification 3 of the fixed-effects model separately for men and women in three age groups: younger than 30 years, between 30 and 50 years inclusive, and older than 50 years. The results are given in Panel E of table 4. There is a statistically significant part-time wage premium in all cases except for middle-aged men employed on casual contracts.

Table 4 - Sensitivity Analysis of the Part-time/Full-time Wage Differential, Model 1, Specification 3

	Sensitivity to	*Casual Status*	*Men*	*Women*
A	Outliers, namely	Non-casuals	0.148 ***	0.098 ***
	hours <10 or >70 or		**	n.s.
	wage < $5 or >$160	Casuals	0.078 ***	0.084 ***
B	Definition of PT work	Non-casuals	0.182 ***	0.119 ***
	= 30 hours/week		*	n.s.
		Casuals	0.106 ***	0.099 ***
C	Definition of PT work	Non-casuals	0.122 ***	0.097 ***
	= 38 hours/week		**	n.s.
		Casuals	0.061 ***	0.087 ***
D	Random errors in	Non-casuals	0.157 ***	0.101 ***
	self-reported hrs/week		**	n.s.
		Casuals	0.082 ***	0.087 ***
E	Age group	Non-casuals	0.224 ***	0.114 ***
	<30 years old		**	n.s.
		Casuals	0.091 **	0.111 ***
		Non-casuals	0.099 ***	0.090 ***
	30-50 years old		n.s.	n.s.
		Casuals	0.050	0.084 ***
		Non-casuals	0.149 ***	0.125 ***
	>50 years old		n.s.	n.s.
		Casuals	0.134 **	0.114 *
F	Number of PT/FT changes	Non-casuals	0.186 ***	0.096 ***
	0 or 1		n.s.	n.s.
		Casuals	0.133 ***	0.131 ***
		Non-casuals	0.156 ***	0.110 ***
	0 or >1		n.s.	n.s.
		Casuals	0.102 ***	0.130 ***
G	Occupation	Non-casuals	0.184 ***	0.127 ***
	Managers & professionals		n.s.	n.s.
		Casuals	0.169 **	0.102 *
		Non-casuals	0.210 ***	0.145 *
	Technical, trades, community &		**	*
	personal service workers	Casuals	0.071 *	0.062 n.s.
		Non-casuals	0.028 n.s.	0.079 ***
	Clerical, admin & sales workers		n.s.	n.s.
		Casuals	0.062 n.s.	0.071 ***
		Non-casuals	0.111 **	0.027 n.s.
	Machine operators, drivers & labourers		n.s.	n.s.
		Casuals	0.083 *	0.066 n.s.
H	Time period	Non-casuals	0.143 ***	0.096 ***
	2001-2004		n.s.	n.s.
		Casuals	0.093 **	0.097 ***

Table 4 - Sensitivity Analysis of the Part-time/Full-time Wage Differential, Model 1, Specification 3 (continued)

	Sensitivity to	*Casual Status*	*Men*	*Women*
	2005-2008	Non-casuals	0.145 ***	0.127 ***
			n.s.	n.s.
		Casuals	0.097 **	0.121 ***
	2009-2012	Non-casuals	0.175 ***	0.134 ***
			n.s.	n.s.
		Casuals	0.128 ***	0.128 ***
I	>= 2 consecutive obs	Non-casuals	0.137 ***	0.086 ***
			**	n.s.
		Casuals	0.072 ***	0.071 ***
J	Salary Sacrifice & Non-Cash Benefits	Non-casuals	0.181 ***	0.150 ***
			n.s.	n.s.
		Casuals	0.139 ***	0.150 ***

Source: HILDA, Waves 1-12.
*, ** and *** beside a coefficient indicates significantly different from zero at the 5%, 1% and 0.1% levels. Asterisks or n.s. on the line between coefficients indicates a significant difference, or no significant difference (n.s.), between the PT/FT wage differential for non-casual and casual employees.
Notes: [a] Standard errors are cluster robust. [b] The 1,904 outliers on hours and wage were excluded from all sensitivity tests.

Panel F of table 4 shows the premium exists for people who change employment status once only and also for people who change their employment status multiple times. These results were obtained from separate fixed-effects estimations using the 4,505 men who never change employment status and (a) the 510 men who change employment status once only, and (b) the 304 men who changed multiple times. Fixed-effects were also estimated using the 3,754 women with no change of status and (a) the 1,064 women who changed employment status exactly once, and (b) the 679 women with multiple changes.

We estimated the part-time/full-time wage differential separately for four occupational groups (see Panel G of table 4). For men there is a part-time premium for all occupations except clerical, administrative and sales workers. Women in all occupations experience a part-time premium except machine operators, drivers and labourers. For technical, trades, community and personal service workers the part-time premium is significantly larger for non-casuals than for casuals.[10]

The part-time/full-time wage differential was estimated separately for the periods 2001-2004, 2005-2008 and 2009-2012, the latter being after the global financial crisis (GFC) hit the Australian economy (see Panel H of table 4). The part-time premium is observed for both men and women in all three periods and it is largest in the post-GFC period.

We also tested the sensitivity of our results to the constraint that individuals contribute data to the sample only if they have valid observations in at least two

[10] We also estimated the part-time/full-time wage differential separately for 19 industry groups but in most cases the samples were too small for the results to be meaningful.

consecutive waves of HILDA data. This requirement eliminated 5,259 observations on men and 5,953 observations on women, so it is relevant to ask whether doing so introduced bias into our results. Panel I of table 4 shows that the part-time wage premium remains for men and women, casual and non-casual, although in each case it is a little smaller than in Panel A.

Salary Sacrifice and Non-Cash Benefits Included in the Hourly Wage

The results discussed above, and the analysis of Booth and Wood (2008), are based on a measure of weekly labour income that excludes salary sacrifice and non-cash benefits.[11] If full-time employees receive larger amounts from these sources than part-time employees then their exclusion from the hourly wage could explain the part-time wage premium. This possibility was explored using data from Waves 10 through 12 of our sample, which are the only waves in which salary sacrifice and non-cash benefits are recorded. Outliers defined in Panel A of table 4 were excluded. The subset contains data on 3,217 males (8,353 male-years) and 3,244 females (8,331 female-years).

The same specifications of Model 1 were estimated but the set of control variables in Specifications 2 to 5 was expanded to include a binary variable equal to one if the individual supervises others in his or her main job, a higher wage being expected for supervisors, ceteris paribus. We also included three dummy variables indicating whether the individual's pay is set by a collective agreement, an individual agreement, a combination of the two, or (the excluded category) by an industrial award. These variables account for the possibility suggested by Booth and Wood (2008, p.132) that the premium results from employers paying wages that are above awards rates to employees on collective or individual agreements.

The results indicate that the premium is not a compensating differential for salary sacrifice and non-cash benefits, it is not explained by the method of setting pay, nor is it a loading for supervisory duties. The fixed-effects estimates in Specification 3 indicate substantial, statistically significant part-time wage premiums (see Panel J, table 4). The premium is 13.9 per cent for men employed on casual contracts and 18.1 per cent for men who are non-casuals. For women, there is a part-time wage premium of 15.0 per cent for both casuals and non-casuals.

Model 2

The fixed-effects assumption of symmetry is relaxed in the difference-in-differences equations of Model 2. We modified the control variables, dropping those that showed little variation 'within' employees and simplifying others. This resulted in a set of controls for changing casual status, changing employer, moving between single and multiple jobs, shifting between standard and non-standard work schedules, changing from a fixed contract to some other type of employment contract, as well as for joining or leaving a trade union, moving between public-sector and private-sector jobs, changing between jobs with and without supervisory duties, and two controls for switching to a job with a small firm and to a job with a large firm.

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[11] For information on the two variables, see Kecmanovic and Wilkins, 2011.

With the sample consisting of at least four consecutive observations on each individual, we were able to estimate Model 2 with k=0, 1 and 2, obtaining estimates of wage changes from the year immediately prior to a change of employment status, to the year immediately after, as well as one and two years beyond. We consider three cases. In Case 1 the control group consists of people who do not change employment status between years t-1 and t, while the treatment group consists of people who do change status between years t-1 and t; no further restrictions are placed on either group. To determine how sensitive the results are to the definitions of the treatment and control groups, two additional sets of restrictions were imposed on the sample. In Cases 2 and 3, the control group consists of people who are never observed to change employment status. In Case 2, the treatment group has no change of employment status prior to that which occurs between years t-1 and t. In Case 3, the treatment group has exactly one change of status prior to year t+2, namely the change that occurs between years t-1 and t. Table 5 shows the numbers of transitions between full-time and part-time employment, with and without a change of employer, for men and women in each of Cases 1, 2 and 3.

Table 5 - Transition Matrices Showing Changes of Employment Status, Model 2

		Men					
		All		*No change of employer*		*Change of employer*	
Year t-1 ↓	*Year t* →	*Part-time*	*Full-time*	*Part-time*	*Full-time*	*Part-time*	*Full-time*
Case 1							
Part-time		531	319	481	198	50	121
Full-time		237	12,944	133	11,524	104	1,420
Case 2							
Part-time		209	178	194	97	15	81
Full-time		183	11,537	93	10,315	90	1,222
Case 3							
Part-time		209	144	194	73	15	71
Full-time		45	11,537	26	10,315	19	1,222
		Women					
		All		*No change of employer*		*Change of employer*	
Year t-1 ↓	*Year t* →	*Part-time*	*Full-time*	*Part-time*	*Full-time*	*Part-time*	*Full-time*
Case 1							
Part-time		4,057	826	3,748	627	309	199
Full-time		658	7,324	490	6,546	168	778
Case 2							
Part-time		2,152	550	2026	404	126	146
Full-time		342	4,827	241	4,286	101	541
Case 3							
Part-time		2,152	320	2,026	215	126	105
Full-time		175	4,827	128	4,286	47	541

Source: HILDA, Waves 1-12.
Note: Numbers are unweighted counts of transitions, based on pairs of consecutive waves.

To begin, however, Model 2 with k=0 was estimated using the same sample of two or more consecutive observations for each person as was used to estimate Model 1.[12] The results appear in table 6, labelled Case 0. They are consistent with the results from Model 1 in that they show substantial premiums for part-time work and penalties for full-time work, for both men and women. The results, however, are not symmetric. For men the premium for part-time work is smaller (12.0 per cent) than the penalty for full-time work (14.4 per cent). For women the reverse is observed; the premium for part time work (14.2 per cent) is larger than the penalty for full-time work (11.6 per cent).

Table 6 - Part-time/Full-time Wage Differentials, Model 2

			Men	*Women*
Case	*Change of Status*	*Period*	*Cumulative Effect*	*Cumulative Effect*
0	Pjoiner	t-1/t	0.120 ***	0.142 ***
0	Pleaver	t-1/t	-0.144 ***	-0.116 ***
1	Pjoiner	t-1/t	0.114 ***	0.133 ***
		t-1/t+1	0.015	0.054 ***
		t-1/t+2	0.040	0.047 ***
1	Pleaver	t-1/t	-0.126 ***	-0.106 ***
		t-1/t+1	-0.050	-0.024
		t-1/t+2	-0.039	-0.017
2	Pjoiner	t-1/t	0.093 **	0.145 ***
		t-1/t+1	-0.009	0.052 *
		t-1/t+2	0.016	0.041 *
2	Pleaver	t-1/t	-0.133 **	-0.098 ***
		t-1/t+1	-0.085 *	-0.023
		t-1/t+2	-0.063	-0.011
3	Pjoiner	t-1/t	0.176 *	0.112 ***
		t-1/t+1	0.049	0.076 **
		t-1/t+2	0.068	0.056 *
3	Pleaver	t-1/t	-0.090 *	-0.050 *
		t-1/t+1	-0.075	-0.021
		t-1/t+2	-0.099	-0.003

Source: HILDA, Waves 1-12.
*, ** and *** indicates significantly different from zero at the 5%, 1% and 0.1% levels. Note: OLS estimates with cluster robust standard errors.

The results for Cases 1-3 also appear in table 6.[13] Full results for Case 1 are given in appendix A2. The wage changes with k=0 are all large (greater than 10 per cent), highly statistically significant and display the same form of asymmetry as in Case 0. What is noteworthy about Case 1 for men, and for women for switch from part-time to full-time work, is that by year t+1, the effect of a change in employment status is practically small and not significantly different from zero. For women who change from full-time to part-time work, the effect one and two years later is statistically significant

[12] The outliers specified in Panel A of table 4 were excluded. Salary sacrifice, non-wage benefits and type of contract were not included because these variables are available only in recent waves of HILDA data.
[13] Full results for all cases in tables 6 and 7 are available from the authors on request.

but about one third the size of the initial effect of 13.3 per cent. As in Case 1, there is a substantial, asymmetric, immediate effect on the wage of changing employment status in Cases 2 and 3 but there is no evidence that it is sustained, except for women who switch from full-time to part-time employment. For these women the effect after two years is statistically significant and 4.1 per cent (Case 2) and 5.6 per cent (Case 3).

Finally, we investigate whether the effect on the wage of changing between full-time and part-time work is different for those who change employers and those who do not. Assuming (a) unpaid overtime is more common among full-time, than among part-time, workers and (b) people who change between full-time and part-time work with the same employer are more likely to retain the same contracted hourly wage than people who simultaneously change employer, then any change in the hourly wage – as measured with HILDA data – will be inversely related to the change in hours worked. Furthermore, as Hirsch (2005, p.541) points out, a change in full-time/part-time status is less likely to be mismeasured for workers who simultaneously change employer.

The variable $newemp_{i,t-1/t+k}$ in Model 2 equals zero if the individual has the same employer in every year from t-1 through t+k, and equals one otherwise.[14] The results are presented in table 7. In Case 0, the contemporaneous wage premium associated with changing to part-time employment is statistically significant only if the employer is unchanged. There is a statistically significant contemporaneous wage penalty for changing to full-time employment for both men and women, with or without a change of employer. For men, the penalty is larger with a change of employer; for women, it is larger with no employer change.

In Cases 1, 2 and 3 there is stronger evidence of a wage change when the individual remains with the same employer when changing employment status than when he or she simultaneously changes employer. In all three cases, women experience a wage premium associated with switching to part-time work provided they stay with the same employer. The premium decreases over time but after two years it is at least eight per cent and statistically significant. A sustained wage premium is also observed for men who switch to part-time work without changing employers in Case 1. In Cases 2 and 3 the statistical evidence is weaker, possibly because of small numbers of males changing employment status, particularly in Case 3 where only 19 men change to part-time work with a different employer. On the other hand, when a change to part-time employment is accompanied by a change of employer there is no evidence of a sustained part-time wage premium for men, and only a little evidence of a part-time wage premium for women in Case 2.

In all three cases, women experience an immediate wage penalty associated with switching to full-time work if they remain with the same employer. However, with time the penalty quickly decreases and becomes statistically non-significant. For men who do not change employers, there is evidence of a temporary full-time wage penalty in Cases 1 and 2, but not Case 3. On the other hand, when a change to full-time employment is accompanied by a change of employer there is little evidence of a full-time wage penalty for women. For men, the evidence is mixed. A large, statistically significant, immediate full-time wage penalty is observed in Cases 2 and 3, but not Case 1.

[14] A change of job within the same firm does not constitute a change of employer.

Table 7 - Part-time/Full-time Wage Differentials, With and Without a Change of Employer, Model 2

			Men		Women	
			Cumulative Effect		Cumulative Effect	
Case	Change of Status	Period	Same Employer	Different Employer	Same Employer	Different Employer
0	Pjoiner	t-1/t	0.182 ***	0.021	0.163 ***	0.060
0	Pleaver	t-1/t	-0.137 ***	-0.174 *	-0.121 ***	-0.099 **
1	Pjoiner	t-1/t	0.163 ***	0.044	0.152 ***	0.061
		t-1/t+1	0.085 **	-0.055	0.085 ***	-0.018
		t-1/t+2	0.115 **	-0.020	0.077 ***	-0.011
1	Pleaver	t-1/t	-0.123 ***	-0.138	-0.113 ***	-0.075 *
		t-1/t+1	-0.067 *	-0.008	-0.031 *	-0.002
		t-1/t+2	-0.062	0.002	-0.029	0.008
2	Pjoiner	t-1/t	0.142 ***	0.033	0.159 ***	0.104 **
		t-1/t+1	0.056	-0.063	0.085 **	-0.013
		t-1/t+2	0.109 *	-0.048	0.083 ***	-0.027
2	Pleaver	t-1/t	-0.110 *	-0.264 **	-0.106 ***	-0.054
		t-1/t+1	-0.062	-0.163	-0.022	-0.027
		t-1/t+2	-0.035	-0.135	-0.020	0.008
3	Pjoiner	t-1/t	0.209 **	0.128	0.131 ***	0.052
		t-1/t+1	0.145 *	-0.073	0.118 **	-0.014
		t-1/t+2	0.163	-0.042	0.109 **	-0.035
3	Pleaver	t-1/t	-0.056	-0.253 **	-0.061 **	-0.011
		t-1/t+1	-0.052	-0.143	-0.023	-0.017
		t-1/t+2	-0.077	-0.147	-0.011	0.012

Source: HILDA, Waves 1-12.
*, ** and *** indicates significantly different from zero at the 5%, 1% and 0.1% levels. Note: OLS estimates with cluster robust standard errors.

8. Concluding Remarks

The only previous longitudinal study of the full-time/part-time wage differential in Australia found large part-time wage premiums for both men and women, a result that is at odds with studies from other countries. Using 12, rather than four, years of longitudinal data we find part-time wage premiums for men and women on casual employment contracts of 12.5 and 13.2 per cent, respectively, and for non-casual men and women the premiums are 17.0 and 10.3 per cent, respectively.

Having verified the results of the earlier study, we then proceeded to investigate the nature of the premium. The premium does not result from the loading received by employees on casual employment contracts, many of whom work part-time. The premium is not explained by outlying observations on weekly hours of work or the hourly wage, although it is reduced, particularly for people on casual contracts, when outliers are excluded from the data set. Nor is the premium an artefact of how part-time employment is defined. Random errors in reporting weekly work hours theoretically could explain the premium but empirically they do not appear to do so. Age sheds only a little light on the nature of the part-time wage premium as it is detected for all age groups, except men aged 30-50 years who are employed on casual contracts. Similarly, the premium is observed in three out of four occupational groups, the exceptions being men in clerical, administrative or sales occupations, and women who are machine operators, drivers or labourers. People who change employment status once only and people who change multiple times both experience a part-time wage premium. The premium is observed throughout the period 2001-2012 and has increased since the GFC.

We found no evidence that the part-time wage premium is a payment to part-time workers in compensation for receiving less in non-cash benefits or salary sacrifice than their full-time counterparts. When wages and salaries include salary sacrifice and are augmented with the value of non-cash benefits, the fixed-effects estimates indicate a statistically significant part-time wage premium of at least 14 per cent for both men and women, regardless of whether or not they are employed on casual contracts. Nor is there evidence in support of Booth and Wood's (2008) conjecture that the premium results from employers paying wages that are above awards rates to employees on collective or individual agreements. The premium is observed whether or not the payment method (collective agreement, individual agreement, a combination of the two, or industrial award) is held constant.

Difference-in-difference models reveal that both men and women experience substantial and sustained hourly wage increases when they switch from full-time to part-time work with the same employer. Smaller but statistically significant hourly wage reductions are experienced immediately after changing from part-time to full-time work without a change of employer but two years later the decrease is small and not statistically significant. There is scant evidence of a part-time wage premium (or penalty) when the change of employment status is accompanied by a change of employer.

The results from our dynamic model are consistent with Booth and Wood's (2008, p.132) conjecture that full-time workers are more likely to work unpaid hours than part-time workers. Notably, the one study that found a part-time wage penalty with longitudinal data, Connolly and Gregory (2009), used contracted hours of work to calculate the hourly wage. We were unable to test the unpaid-overtime hypothesis directly because paid and unpaid hours of work are not distinguished in the HILDA data; that is a task for future research into Australia's premium for part-time work.

Appendix

Appendix A1- Full Estimates of Model 1, Specification 3

	Men		*Women*	
	Pooled OLS	*Fixed Effects*	*Pooled OLS*	*Fixed Effects*
Part-time	-0.116 *	0.170 ***	-0.001	0.103 ***
Casual	-0.009	0.037 **	-0.059 *	0.021
Part-time#casual	0.039	-0.044	0.030	0.029
One job only	0.017	-0.029 **	-0.029	-0.011
Standard schedule	-0.061 ***	-0.013	-0.037 **	-0.016
Fixed contract	0.031 *	0.005	0.035 **	0.008
Labour hire firm	0.068 *	0.067 ***	0.107 **	0.052 **
Vic	-0.020	0.036	-0.064 ***	0.022
Qld	-0.021	-0.037	-0.052 ***	0.033
SA	-0.099 **	-0.040	-0.075 **	0.003
WA	0.022	0.019	-0.008	-0.025
Tas	-0.084 *	-0.012	-0.017	-0.054
NT	0.028	0.044	0.014	-0.011
ACT	0.110 **	0.066	0.102 ***	-0.028
Married	0.149 ***	0.040 **	0.083 ***	0.037 *
De facto	0.118 ***	0.026 *	0.088 ***	0.039 **
Sep, divorced, widowed	0.097 **	0.042 *	0.058 **	0.033
Born other ES country	0.036	-	0.024	-
Born NES country	-0.100 ***	-	-0.048 **	-
Inner regional	-0.070 ***	-0.018	-0.047 ***	-0.046 **
Outer regional	-0.044	0.008	-0.066 **	-0.043
Remote area	0.069	0.075	0.009	0.078
25-29 yrs old	0.038	0.011	0.053 **	0.012
30-34 yrs old	-0.006	-0.031	0.021	-0.007
35-39 yrs old	-0.057	-0.077 **	-0.015	-0.025
40-44 yrs old	-0.095	-0.109 ***	-0.032	-0.054
45-49 yrs old	-0.175 *	-0.139 ***	-0.066	-0.072 *
50-54 yrs old	-0.227 *	-0.138 ***	-0.100 *	-0.087 *
55-64 yrs old	-0.229 *	-0.146 ***	-0.098 *	-0.090 *
Tenure	0.008 ***	-0.001	0.010 ***	0.001
Tenure-sq	0.000 *	0.000	0.000 **	0.000
Experience	0.027 ***	0.037 **	0.024 ***	0.034 **
Experience-sq	0.000 ***	-0.001 ***	0.000 ***	-0.001 ***
Postgraduate	0.556 ***	0.096	0.466 ***	0.168 ***
Grad dip/cert	0.489 ***	0.093 *	0.390 ***	0.123 **
Bachelor	0.460 ***	0.006	0.358 ***	0.088 *
Diploma	0.262 ***	0.061	0.180 ***	0.061
Certificate 3/4	0.163 ***	0.076 **	0.070 ***	0.035
Year 12	0.153 ***	-0.103 *	0.103 ***	0.018
Constant	2.673 ***	2.881 ***	2.719 ***	2.668 ***
Person-yr observations	29,199	29,199	28,937	28,937
F-statistic	49.718 ***	39.948 ***	64.550 ***	30.442 ***
Breusch-Pagan χ^2 stat		33,800.2 ***		17,065.2 ***
Hausman χ^2 stat		702.9 ***		527.3 ***

Source: HILDA, Waves 1-12.
*, ** and *** beside a coefficient indicates significantly different from zero at the 5%, 1% and 0.1% levels.
Note: Standard errors are cluster robust.

Appendix A2 - Full Estimates of Model 2, Case 1

	Men			Women		
	Coef t-1/t	*Coef t-1/t+1*	*Coef t-1/t+2*	*Coef t-1/t*	*Coef t-1/t+1*	*Coef t-1/t+2*
Pjoiner	0.114 ***	0.015	0.040	0.133 ***	0.054 ***	0.047 ***
Cleaver	-0.071 ***	-0.067 ***	-0.083 ***	-0.067 **	-0.069 **	-0.064 *
Cjoiner	0.016	0.049 *	0.044 *	0.027	0.019	-0.001
Cstayer	0.005	0.024	0.031	-0.035	-0.026	-0.037
Δemployer	0.038 ***	0.053 ***	0.062 ***	0.036 **	0.045 ***	0.050 ***
Δmultijobs	-0.023	-0.026	-0.025	-0.007	-0.003	-0.009
Δschedule	0.013	-0.008	-0.016	-0.030 *	-0.029 *	-0.045 **
Δfcontract	-0.002	0.000	-0.004	-0.005	0.000	-0.004
Δunion	0.033 ***	0.036 ***	0.035 ***	-0.011	0.006	0.007
Δpublic	0.012	0.019	0.018	0.039 **	0.032 *	0.057 ***
Δsupervise	-0.001	0.010	0.012	0.012	0.014 *	0.014 *
Δsmallfirm	-0.012	-0.029 **	-0.037 **	-0.008	-0.016	-0.027
Δbigfirm	0.029 *	0.046 ***	0.064 ***	0.028 *	0.024	0.043 **
constant	0.017 *	0.035 ***	0.060 ***	0.013	0.044 ***	0.063 ***
N	13181	13181	13181	7982	7982	7982
F-statistic	4.779 ***	7.516 ***	7.835 ***	7.790 ***	4.550 ***	6.293 ***
	Men			*Women*		
	Coef t-1/t	*Coef t-1/t+1*	*Coef t-1/t+2*	*Coef t-1/t*	*Coef t-1/t+1*	*Coef t-1/t+2*
Pleaver	-0.126 ***	-0.050	-0.039	-0.106 ***	-0.024	-0.017
Cleaver	-0.016	-0.023	-0.017	-0.027	-0.038 *	-0.020
Cjoiner	0.028	-0.086	-0.083	0.069	0.037	0.016
Cstayer	-0.019	0.007	-0.004	0.011	0.020	0.010
Δemployer	0.050	0.064	0.077 *	0.053 **	0.054 ***	0.045 **
Δmultijobs	-0.006	-0.024	0.004	-0.047 **	-0.048 **	-0.040 **
Δschedule	-0.050	-0.064	-0.024	-0.002	-0.006	-0.017
Δfcontract	0.088 *	0.149 **	0.091	0.006	0.006	0.007
Δunion	0.094 *	0.084 *	0.070	0.003	-0.005	0.020
Δpublic	0.110	0.064	0.076	-0.013	-0.008	0.002
Δsupervise	-0.001	0.002	-0.006	0.010	0.010	-0.007
Δsmallfirm	0.042	0.016	-0.017	-0.039	-0.060 **	-0.035
Δbigfirm	0.140 *	0.116 *	0.048	0.043	0.010	0.037
constant	-0.044 ***	-0.004	0.031	-0.014	0.003	0.018
N	850	850	850	4883	4883	4883
F-statistic	2.485 ***	2.071 **	1.635 ***	5.220 *	2.427 ***	2.264 ***

Source: HILDA, Waves 1-12.
*, ** and *** beside a coefficient indicates significantly different from zero at the 5%, 1% and 0.1% levels.
Note: Standard errors are cluster robust.

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