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Back-to-front Down-under?

Part-time/Full-time Wage Differentials in Australia*

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ABSTRACT

In 2003, part-time employment in Australia accounted for over 42% of the Australian female workforce, nearly 17% of the male workforce, and represented 28% of total employment. Of the OECD countries, only the Netherlands has a higher proportion of working women employed part-time and Australia tops the OECD league in terms of its proportion of working men who are part-time. In this paper we investigate part-time full-time hourly wage gaps using important new panel data from the new Household Income and Labour Dynamics in Australia Survey. We find that the usual negative part-time wage penalty found in other countries is not found in Australia once unobserved individual heterogeneity has been taken into account. Instead, part-time men and women typically earn an hourly pay premium. This result survives our numerous robustness checks and we advance some hypotheses as to why there is a positive part-time pay premium.

Keywords: part-time, full-time; efficiency hours; gender

JEL Classifications: J16, J22, J31

1. Introduction

In 2003, part-time employment in Australia accounted for over 42% of the Australian female workforce, nearly 17% of the male workforce, and represented 28% of total employment (OECD, 2004). Of the OECD countries, only the Netherlands has a higher proportion of working women employed part-time and Australia tops the OECD league in terms of its proportion of working men who are part-time.[1] Against this background, our aim is to investigate whether there is, in Australia, a pay premium or penalty associated with part-time work relative to full-time work. We also explore the degree to which observed pay gaps differ by gender.

Part-time jobs are often viewed as bad jobs with low pay and little career prospects. Studies based on representative survey data typically find a part-time pay penalty (see for example see Simpson (1986), Main (1988), Blank (1990), Ermisch and Wright (1992) and for a review of US studies see Hirsch (2004)). However more recent analysis by Hirsch (2004), using US panel data, finds little evidence of a pay gap between part-time and full-time women but he does find a part-time pay penalty for men. And Rodgers (2004), using cross-sectional data from Wave 1 of the Household Income and Labour Dynamics in Australia (HILDA) Survey, finds no evidence of a part-time pay penalty for either men or women.

In this paper we use important new panel data from the first two waves of the Household Income and Labour Dynamics in Australia (HILDA) Survey. Two advantages of these data are that they provide a very rich set of controls and they allow for estimation controlling for unobserved heterogeneity. Our estimates reveal that part-time men and women in Australia typically earn a statistically significant hourly wage premium. This result survives our numerous robustness checks and in the conclusion of the paper we discuss some theories consistent with these stylized facts.

To start with, it is useful to review the various hypotheses about the determinants of part-time/full-time wage differentials. Some suggest that part-time work should be associated with a penalty, while others suggest it might command a premium. Although all depend on the interaction of demand and supply factors, we group them below under the broad headings of firms' preferences, workers' preferences, institutional factors and measurement issues.

[1] According to the OECD Employment Outlook (2004:310), the OECD average for 2003 was 25% for women and 7% for men.

1. *Demand side factors*

   1.1. Fixed employment costs might mean firms prefer employees to work longer hours in order to recoup hiring and setup costs. According to this hypothesis, there should be a penalty to part-time work.

   1.2. However the efficiency hours hypothesis can predict the reverse. Suppose there is a hill-shaped relationship between hourly efficiency and the number of hours worked in a day or a week, as suggested by Booth and Ravallion (1993) in the debate on effects of hours cuts. Then part-time work could be associated with a pay premium.

   1.3. Some firms' production schedules require part-time workers for demand peaks. For example, supermarkets and restaurants have variable demand, which might be best met by part-time workers. If there is a fixed supply of such workers, any wage gap will reflect relative demand and supply factors.[2]

   1.4. Firms might have more market power over part-time workers than full-time workers. For example, part-timers might have childcare commitments constraining them to seek work close to home. This could give employers a greater degree of market power over part-time workers.

2. *Supply-side factors*

   2.1 According to human capital theory, individuals who anticipate working part-time will invest less in education than those who intend to work full-time. In addition, part-time workers accumulate experience capital at a lower rate than full-time workers, because they supply fewer hours. We should therefore observe a part-time pay penalty, as part-time workers have lower levels of both human and experience capital.

   2.2 Part-time work might suit worker's heterogeneous preferences (which can of course be affected by policy and we return to this below). Whether or not there are wage differentials depends on the supply of workers who prefer part-time work and the demand for them.

[2] Firms that are able to offer a range of part-time and full-time jobs might face better choice of workers. In countries where part-time work is preferred by many workers, firms could tinker the mix of jobs according to preferences of local labour supply. Firms in good position to offer part-time jobs might be in strong bargaining position to drive down part-time pay. Firms in poor positions might have to increase part-time pay.

3. *Institutional factors*

3.1 Australia has been characterized by a unique award system of rates of pay (see Pocock, 1995, for a summary of how this affected women in particular). Award provisions relating to part-time employment were largely designed to meet the needs of the industry concerned as well as to protect future employment of both part-time and full-time employees (Hawke, 1993). A possible outcome of this might have been that part-time workers received the benefits of the annual award system for covered workers quite independently of their union status. However the importance of this system has diminished over recent time, and in particular with the passage of the 1996 Workplace Relations Act that explicitly emphasized workplace bargaining between employer and employee without union or state intervention.[3]

3.2 In Australia, effective marginal tax rates are high for relatively low-skilled second earners (Apps, 2004). Consequently firms requiring part-time workers may have to pay more to attract marginal workers into market-sector employment.

4. *Measurement issues*

Finally, there are measurement issues to do with the total compensation package received by part-time and full-time workers. Part-time workers might receive non-wage benefits that differ from those offered to full-time workers. While Australian firms do not offer health insurance to their employees, casual workers do frequently receive buy-outs for holiday and sick pay. However we have information about this in our data set, and so we are able to control for casual status in regressions of hourly wages.

2. <u>The Data, Variables and Raw Part-Time Wage Gap</u>

2.1 The data and variables

We use data from Waves 1 and 2 of the new Household Income and Labour Dynamics (HILDA) survey, a nationally representative random-sample survey of private

[3]The 1996 Workplace Relations Act introduced Australian Workplace Agreements, whereby workers might be expected to sign an individual contract and forfeit award coverage. These new workplace agreements were supposed to match the existing award conditions. While the Industrial Relations Commission can review and recommend change, or reject agreements, it is no longer able to set minima for part-time work or set precedents in terms of regulations. In addition, the Act weakened legal recognition of unions with respect to rights of entry and representation.

households in Australia spanning the period 2001-2.[4] Our analysis covers full-time and part-time employees aged between 18 and 60 years in Wave 1, who are not in the armed forces, farming or fisheries, and with valid information on our main variables (hours of work, salary, and whether casual or permanent). Individuals reporting over 100 working hours per week (hours are used to derive hourly wages) were dropped, as were full-time students. Where there were many missing observations for control variables, we created dummy variables indicating their status, to maintain reasonable sample sizes. We use a balanced panel of respondents who are present - and satisfy the selection criteria - in both waves. Our estimating sample comprises 1994 women and 2034 men, representing 3988 person-year observations for women and 4068 person-year observations for men.

Our measure of *part-time work* is based on individual responses for their usual hours of work in their main job (including any paid or unpaid overtime, and work done at the workplace and at home). Part-time workers are defined as those usually working fewer than 35 hours per week (where we follow the Australian Bureau of Statistics cutoff of fewer than 35 hours rather than the OECD cutoff of 30).

The *hourly wage* refers to the hourly wage in the main job.[5] Wave 2 wages were deflated to 2001 (wave 1) levels using the headline Consumer Price Index (CPI) from the Australian Bureau of Statistics. Respondents earning less than $1 or more than $100 have been omitted from the analysis.

Table A.1 in the Data Appendix gives the means of some of the variables of interest, disaggregated by gender and by full-time or part-time employment status. Full-time women work on average 42.09 hours per week and earn an average hourly wage of A$17.22, while part-time women work 20.58 hours to earn A$16.41. This compares with full-time men who work 45.31 hours (earning A$19.75) and part-time men who work on average 20.43 hours (and earn A$16.09). A greater proportion of part-time workers of both sexes are casual (43% of women and 63% of men). Casual workers in Australia are defined as those who are ineligible for sick and holiday pay entitlement and who are

[4] The survey is a longitudinal study of representative households in Australia. Wave 1 comprised 7682 households with 13,969 respondents aged 15 years and over. For Wave 2, interviews were conducted in 7245 households with 11,993 respondents continuing from Wave 1 and 1048 new respondents. The attrition rate between Waves 1 and 2 was 13.2 per cent. HILDA contains four survey instruments; the household form, a household questionnaire, a person questionnaire and the self-completion questionnaire. The information at the household level can be provided by any adult member of the household but preferably a person with knowledge of the household finances. The person-level questionnaires are for all persons aged 15 years and over in the household.

[5] The log wage was calculated from the HILDA variables as log (hourly wage) = log {gross annual salary in main job /[(52.14)*(usual weekly hours in main job including overtime)]}. No specific information is provided on overtime hours and premia. Notice that 52.14 = 365 days per year divided by 7 days per week.

usually paid a wage premium as a compensating differential.[6] Proportionately more full-time workers are on fixed term contracts and working on a regular daytime schedule. Some 94% of full-time women and men have only one job, as compared with 84% and 76% of part-time women and men respectively. Fewer part-time workers of both sexes are on standard daytime schedules and slightly more full-time workers are on fixed term contracts.

There are also some interesting differences across demographic and educational variables. Proportionately more married and cohabiting women work part-time while more married or cohabiting men work full-time. Full-time women and men are on average just over 38 years old, while part-time women are 39 and part-time men 36. Experience and tenure levels are fairly similar for full-time and part-time women, but part-time men are less experienced and have lower tenure levels than full-time men who have the highest levels of these variables. Finally, note that full-time women have higher educational levels than full-time men; 36% of full-time women and 27% of full-time men have at least a university degree and above, as compared with 24% of part-time women and men; ie there are only 3 percentage points difference between the percentage of full-time men with degrees and the percentages of part-time workers with degrees.

2.2 The distribution of hours and hourly wages

Figures 1(a) and 1(b) give the distribution of usual weekly hours worked in the main job for women and men respectively (where observations are pooled across waves). For both men and women there is a spike at 40 hours per week, but female hours are also more dispersed across the lower part of the distribution. In addition, there are small spikes at five hourly intervals, as is usual in reported hours per week.

[6] Wooden and Warren (2003) emphasize the importance of distinguishing between causal work, fixed term contracts and temporary agency work. HILDA is particularly appropriate for studying part-time/full-time wage gaps, owing to the fact that usual pay and hours information is given for the main job (facilitating classification into part-time based on hours in main job), and that there is a very rich set of other controls potentially affecting wage determination, including casual and contract status. Clearly the richer the set of controls, the lower is unobserved heterogeneity.

Figure 1(a)

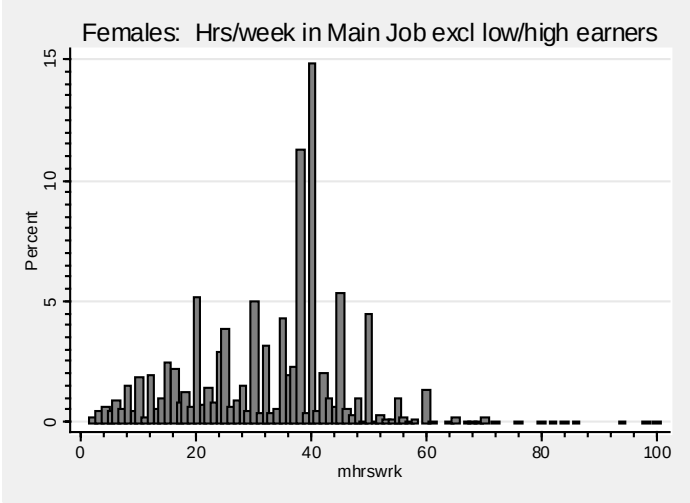
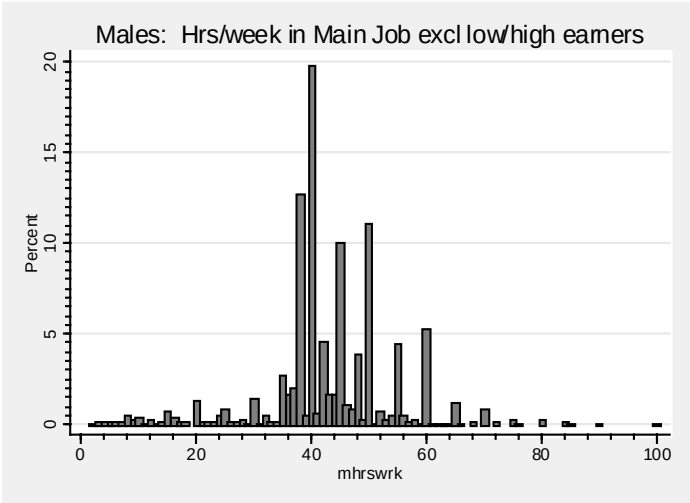


Figure 1(b)



Figures 2(a) and 2(b) show mean hourly wages by hours worked across the distribution (from 0 to 60 hours). Inspection of these figures reveals first, that there is considerable noise at hourly intervals not divisible by five (fewer workers are observed at these points) and also at the bottom of the hours distribution. Second, there is a relatively flat profile especially over the interval 5-50 hours for women. This is in contrast to results for the US found by Hirsch (2004), where hourly wages increased across the hours distribution and especially so for men. Third, men have a slight jump in hourly wages at around 35-40 hours per week but the effect is barely discernible for women. Finally, there is a slight tendency for hourly wages to decline with hours worked for women supplying more than 50 hours per week and men supplying more than 40 hours per week, perhaps reflecting the fact that some of these workers are not paid for overtime hours (as happens with some salaried workers).

Figure 2(a)

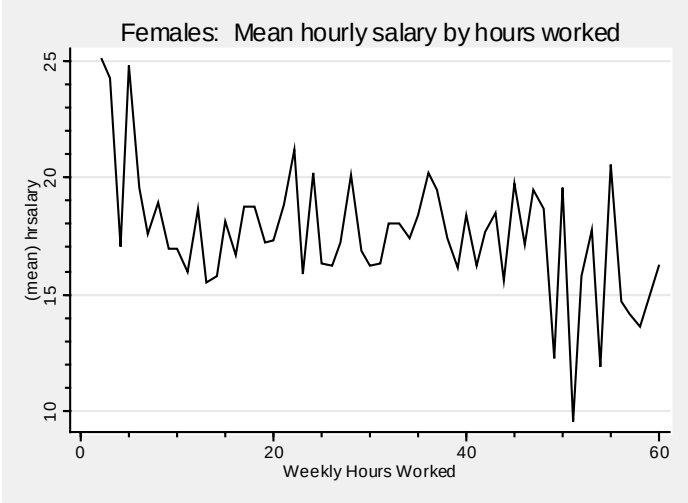
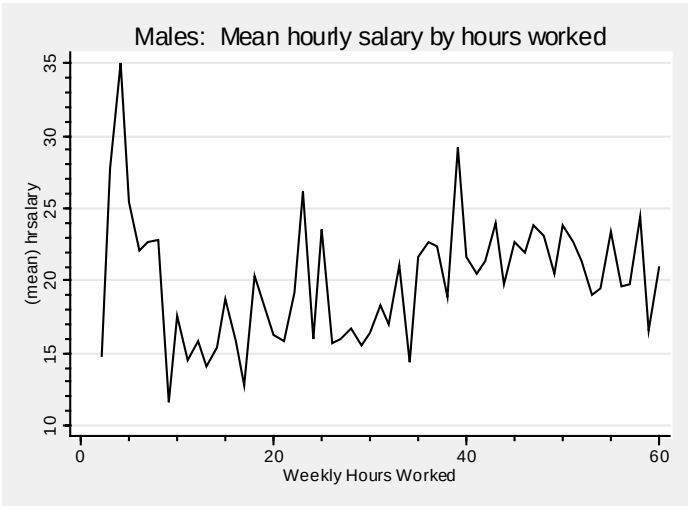


Figure 2(b)



3. <u>Part-Time/Full-Time Wage Gap Estimates</u>

3.1 The econometric model

Our estimating equation, which incorporates the influences of various observed and unobserved characteristics on the log of hourly wages, is given by:

$$\ln w_{it} = X'_{it}\beta + \alpha P_{it} + \mu_i + \varepsilon_{it} \qquad (1)$$

where $i=1,\ldots,N$ represents the number of individuals at each wave and $t=1,2$ is the number of waves. Note that $\boldsymbol{X}_{it}$ is a vector of characteristics that influence the outcome variable w_{it}; the associated parameter vector is $\boldsymbol{\beta}$; P_{it} denotes part-time employment status; μ_i is an unobserved individual-specific effect; and ε_{it} is a random error term. The parameter of interest is α.

Cross-sectional estimation of equation (1) is likely to produce biased estimates of α, since individuals are likely to self-select into full-time employment status based on unobservable factors. Suppose that μ_i denotes an individual's ability in market production relative to home production and μ_i is fixed over time. Suppose further that this is negatively correlated with self-selection into part-time jobs and positively correlated with hourly wages. Then the estimated coefficient for part-time work in a cross-sectional regression will be negatively biased through the omission of any control for unobserved μ. Once control has been taken of unobserved heterogeneity through estimation of equation (1) using appropriate panel data models, we would expect that the part-time wage gap would become larger, *ceteris paribus.* Since failure to control appropriately for unobservables will result in omitted variable bias to the coefficient α, we utilise panel data techniques to control for unobserved heterogeneity. We then compare these with the estimates obtained from ordinary least squares (OLS) using pooled person-year observations.

3.2 The estimates

We estimate the models separately by gender. The results are reported in Table 1 for the five different specifications described in the notes under the table. For each specification we give the pooled estimates, the random effects (RE) estimates and the fixed effects (FE) estimates.[7] For readers interested in the impact of other variables, we report in Appendix Table A.2 the full set of estimates from Specification [3]. The first panel of Table 1 gives the estimates from Specification [1], which includes only a constant and the part-time work dummy variable. The pooled OLS estimates show a statistically significant negative effect of part-time work on hourly wages. The RE estimate is positive but insignificant for women and negative and statistically significant for men. However, the FE estimate – preferred across all specifications based on the tests reported under Table A.2 in the Appendix – is statistically significant and positive for both men and women.

[7] FE is consistent when μ_i and the x_{ijt} are correlated, whereas consistency of random effects estimation hinges on orthogonality of μ_i and the x_{ijt} . Random effects estimation assumes that the distribution of the random effects conditional on the covariates has a standard normal distribution with zero mean and constant variance. This assumption could be relaxed by allowing for correlation between the unobserved heterogeneity and included covariates by, for example, including time-means of the covariates as additional regressors (Chamberlain, 1980). However, since we only have two waves of data, there is not enough variation in the time-means of the covariates to enable us to account for possible correlation. We have therefore not followed this route.

Table 1: Estimates of Part-time/Full-time Log Wage Differential

	Women		**Men**	
Specification	Coeff	t-stat	Coeff	t-stat
1. Part-time and constant only				
Pooled	-.048***	3.91	-.205***	8.72
RE	.005	0.38	-.040	1.81
FE	.127***	6.14	.128***	4.39
2. base plus individual characteristics				
Pooled	-.002	0.14	-.052**	2.00
RE	.032**	2.28	.033	1.39
FE	.128***	5.94	.146***	4.83
3. (2) plus education dummies				
Pooled	.017	1.38	-.048**	1.97
RE	.041***	3.14	.034	1.48
FE	.128***	5.92	.152***	5.00
4. (3) plus firm attributes and industry dummies				
Pooled	.039***	3.15	.004	0.17
RE	.057***	4.31	.052***	2.33
FE	.130***	5.95	.146***	4.79
5. (4) plus occupational dummies				
Pooled	.066***	5.43	.041*	1.79
RE	.076***	5.90	.069***	3.09
FE	.134***	6.09	.147***	4.79

Notes:
(i) Specification [1] contains only a constant and the part-time employment status dummy.
(ii) Specification [2] also contains personal characteristics (onejob daywork contract casual tempagency, State/Territory, marital status, Australian-born, born in English speaking country, urban dummies, tenure and its squared, age and its square, experience and its square). The base is full-time worker in shift work on a permanent contract, not a casual worker, in NSW in a remote area, single, born in nes country.
(iii) Specification [3] is as for [2] but with educational dummies added with the base being "missed high school".
(iv) Specification [4] adds in firm attributes and industry dummies (union member, establishment size, public sector, one-digit industry dummies). The base is someone who is not a union member, who works in a very small private sector establishment (fewer than 20 employees) in 'other services' industry.
(v) Specification [5] adds in occupational dummies with the base being elementary clerical.
(vi) The full set of estimates from Specification [3] are reported in Appendix Table A.2.
(vii) *** denotes significance at 1% level; ** at 5% level and * at 10% level.

Next consider Specification [2], containing all the additional individual characteristics as detailed in the notes under Table 1, including controls for casual status and for contract type. As shown in the second panel of Table 1, the RE and FE estimates of the part-time coefficient are now both positive and – for the FE model – statistically significant for both men and women.

Each of the additional specifications incrementally adds in blocks of explanatory variables, in the order given in the notes under the table. The FE estimates are preferred for all specifications and they show that part-time workers – once other observable and unobservable characteristics have been taken into account - earn a pay premium of between 13 and 15 log points over and above comparable full-time colleagues. We included occupational status dummy variables in the last specification, Specification [5], since occupation is potentially endogenous. Note that our estimated parameter of interest does not alter with the inclusion of this set of dummies.

The FE estimates are identified from individuals changing their employment status between the two waves. For women, 147 changed from part-time to full-time status and 119 from full-time to part-time, yielding a total of 266 changers. For men, there are 113 changers, comprising 64 men from part-time to full-time and 49 from full-time to part-time. It is interesting that the majority of people who changed employment status did so without changing employer.[8]

Even if one is skeptical about the FE estimates because of their reliance on a relatively small number of changers, the RE estimates – while considerably smaller – nonetheless suggest a clear pay premium for part-time work. From Table 1, it can be seen that the RE estimates of the part-time premium range from approximately 3 to 6% for women, and 3 to 7% for men. Even the OLS pooled estimates suggest a pay premium in Specifications [4] and [5].

Later we will speculate on why there is a part-time pay premium in Australia, but first we report the results of our robustness checks.

3.3 Robustness checks

Table 2 displays the results of a number of different robustness checks. To save space, we report the results only for Specification [3]. In the first panel of Table 2, we report the results of our investigation into measurement error. Our definition of part-time work was based on usual hours of work in the main job being less than 35 hours – the

[8] Of the 147 women who changed employment status from part-time to full-time (between Waves 1 and 2), 121 were asked if they changed employer. Some 74% stayed with the same employer. Of men in the same situation, 64% stayed with the same employer. For the 119 women who changed from full-time to part-time, 96 were asked if they changed employer and 70% stayed with the same employer. For men, 67 out of 96 were asked this question and 70% stayed with the same employer. We also investigated those few cases who were not asked this question about job change and found that the majority of these were characterized by long tenures. Tables A.3.a and A.3.b report the mean characteristics of the job changers at waves 1 and 2 respectively.

ABS definition. But this could be subject to measurement error. Those who are at the margin might have been erroneously classified as part-time when they were actually full-time, or vice versa. To take this possibility into account, in our first robustness check we dropped all cases whose reported usual hours of work in their main job lay in a band of 6 hours around 35 hours. Thus we dropped men and women whose usual hours lay between 32 and 37 hours inclusive. The pooled OLS, RE and FE estimates from this exercise are reported in the first panel of Table 2. Again the preferred estimates are the FE, which are now slightly larger compared to Specification [3] in Table 1 and still statistically significant.

Table 2: Robustness Checks of Part-time/Full-time Log Wage Differential

	Women		**Men**	
Specification	Coefficient	t-statistic	coefficient	t-statistic
1. Omitting PT/FT borderline cases (hrs>=32 and hrs<=37 dropped)				
Pooled	.028**	(2.04)	-.044*	(1.68)
RE	.047***	(3.19)	.041	(1.61)
FE	.151***	(5.13)	.202***	(5.63)
Person-year observations	*3485*		*3779*	
2. Omitting cases < 5 hours				
Pooled	.016	(1.26)	-.050**	(2.04)
RE	.041***	(3.11)	.031	(1.36)
FE	.128***	(6.01)	.149***	(4.89)
Person-year observations	*3938*		*4056*	
3. Omitting cases < 10 hours				
Pooled	.013	(1.05)	-.065***	(2.59)
RE	.036***	(2.81)	.024	(1.06)
FE	.115***	(5.57)	.144***	(4.77)
Person-year observations	*3789*		*4009*	
4. All hours, PT Dummies				
Pooled: 0-9	.139***	(5.04)	.076	(1.43)
10-19	.031*	(1.65)	-.078*	(1.75)
20-29	.018	(1.19)	-.042	(1.27)
30-34	-.021	(1.12)	-.090**	(2.18)
RE: 0-9	.201***	(7.13)	.202***	(4.19)
10-19	.096***	(4.99)	-.024	
20-29	.027*	(1.71)	.066**	(2.24)
30-34	-.004	(0.22)	-.034	(0.97)
FE: 0-9	.437***	(10.32)	.375***	(6.02)
10-19	.294***	(9.68)	.127**	(2.49)
20-29	.093***	(3.69)	.196***	(5.29)
30-34	.060**	(2.33)	.055	(1.30)
Person-year observations	*3988*		*4068*	

Table 2: Robustness Checks of Part-time/Full-time Log Wage Differential (continued)

	Women		**Men**	
Specification	Coefficient	t-statistic	coefficient	t-statistic
5. All hours, PT Dummies, FT Dummy				
Pooled: 0-9	.205***	6.82	.090*	1.67
10-19	.097***	4.37	-.063	1.41
20-29	.083***	4.29	-.029	0.84
30-34	.044*	1.95	-.076*	1.82
FT 35-44	.088***	5.43	.023*	1.80
RE: 0-9	.286***	9.34	.246***	5.07
10-19	.181***	7.95	.018	0.46
20-29	.111***	5.51	.110***	3.61
30-34	.080***	3.59	.009	0.25
FT 35-44	.108***	6.75	.062***	5.16
FE: 0-9	.576***	12.60	.466***	7.46
10-19	.432***	12.29	.215***	4.16
20-29	.232***	7.50	.289***	7.49
30-34	.201***	6.35	.145***	3.36
FT 35-44	.163***	7.49	.121***	7.67
Person-year observations	*3988*		*4068*	

As a second robustness check, we restored those observations described above and instead dropped from our estimating sub-sample all individuals whose usual hours were less than five. We did this to eliminate from the sub-sample the noisy observations illustrated in Figure 2. In the second panel of Table 2, we report our estimates obtained from this sub-sample. Again we find that the estimates are similar to those for Specification [3] in Table 1, being approximately 13 and 15% for women and men respectively.

Third, we repeated this procedure on a further-reduced sub-sample in which we dropped all individuals whose usual hours were less than ten. These estimates are reported in Panel 3 and the FE estimates are approximately 12% and 15% for women and men respectively – ie slightly smaller for women.

Fourth, we restored those observations described above and – on our original sample – we re-estimated Specification [3] but replaced the single part-time dummy variable with four different dummies, representing very short hours (PT0-9), short hours (PT10-19); intermediate hours (PT20-29); and longer part-time hours (PT30-35). This quite flexible specification for hours of work has the advantage of allowing the returns to differ across the hours distribution. These estimates are reported in the bottom panels of

Table 2. Again the FE estimates are preferred. For both women and men, the part-time premium is largest for individuals working very short hours or working short hors, as compared with those whose working hours are closer to full-time. Of course there is a smaller number of workers making transitions across the various hours-of-work dummies so we recommend treating these estimates with caution. Nonetheless, they are interesting since they suggest that the pay premium declines with hours of work.

Finally, we repeated the estimation with an additional dummy variable, FT35_44 (representing usual weekly hours in the range 35-44), with the base being 'more than 44 hours'. Again the FE estimates are preferred, using the appropriate specification checks, although once again we would emphasise caution in interpreting these results because of the small number of changers across the hours categories. However, overall, the results in Table 2 confirm our findings in Table 1: that part-time workers – once other observable and unobservable characteristics have been taken into account - earn a pay premium over and above comparable full-time colleagues.

3.4 What Impact do other forms of flexible work have on wages?

We now consider the impact of other forms of flexible work, in particular whether or not the individual is a casual worker, a temporary agency worker, or working in the day-time/shift etc., and how this differs across gender. To conserve space, we discuss only the results from Specification [3] – reported in full in Appendix Table 2 - but note that the results for these variables are very similar across specifications. Again the FE estimates are preferred and so we focus only on these. For women, part-time status and short-term contract are the only two employment type variables that are statistically significant. Women on short term contracts are paid approx 5% less than are women on permanent contracts (the RE estimate is very similar). For men, contract type is statistically insignificant but being a casual worker reduces hourly pay by approximately 5 log points while being a temporary agency worker is associated with a wage premium of 14 log points (there is no effect for women).

4. Discussion and Conclusions

In the preceding analysis we established that, once account is taken of unobserved heterogeneity using panel data, part-time workers receive an hourly pay premium. The magnitude of this is between 13 and 15%, *ceteris paribus*. We now consider why the observed part-time pay gaps in Australia are so large.

Hypotheses consistent with a positive premium for part-time work are the following. First, according to the efficiency hours hypothesis, part-time workers may be more productive because they are more focused on their jobs for a shorter time period each day and therefore are on the rising part of the hours-productivity hill. For this reason part-time workers might earn a pay premium that could swamp any pay penalty due to lower levels of experience capital.

Second, part-time premia might reflect equilibrium market clearing rates where part-time jobs are plentiful but the supply is constrained. However this explanation does not seem plausible in the Australian context, where there is a high proportion of workers in part-time work.

Third, the part-time pay premium could reflect the high effective marginal tax rates faced by relatively low-skilled second earners in Australia (Apps, 2004). As a consequence, firms with strong demand for part-time workers may have to pay more to attract these workers.

Finally, Australia is characterized by a unique award system of rates and this may mean the part-time workers receive the benefits of the annual award system for covered workers. But high awards rates for part-time workers cannot be the whole story. If award rates were too high, employers would substitute full-time for part-time work. This is clearly not happening in Australia, where the proportion of part-time work has been steadily increasing over time.

Of necessity our discussion of the causes of the part-time pay premia has been speculative. We hope that future work - with more waves of the HILDA data - might be able to investigate some of these hypotheses more fully. However a full investigation of some of the hypotheses relating to part-time pay premia – such as efficiency hours - would require linked employer-employee data, with information about firm's production technology in addition to employee characteristics.

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Appendix Table A.1: Means for Full-time and Part-time (Main Job) Men and Women for Selected Variables

	Women		Men	
	Full-time	Part-time	Full-time	Part-time
Work Attributes				
Log hourly earnings	2.846	2.798	2.983	2.778
Usual hours per week in main job	42.092	20.582	45.308	20.431
Casual	0.063	0.429	0.060	0.632
Fixed term contract	0.115	0.087	0.096	0.083
Employed through a labour hire firm	0.025	0.037	0.030	0.060
One job only	0.946	0.840	0.940	0.758
Regular daytime schedule (main job)	0.834	0.671	0.779	0.537
Tenure with current employer (years)	6.865	5.268	7.733	3.203
Tenure in current occupation (years)	8.956	8.814	10.137	5.212
Trade union member	0.374	0.281	0.376	0.224
Public sector	0.375	0.305	0.247	0.212
Demographics				
Age	38.240	39.460	38.719	35.383
Experience	17.592	17.012	20.440	15.822
Married	0.465	0.619	0.610	0.406
Cohabiting	0.156	0.100	0.123	0.083
Urban	0.677	0.584	0.662	0.630
Inner regional	0.218	0.285	0.238	0.264
Outer regional (base is remote/very remote)	0.084	0.116	0.080	0.086
Australian born	0.772	0.796	0.769	0.761
Born in English speaking country (not Oz)	0.104	0.101	0.117	0.086
Education dummy variables:				
postgraduate degree (masters or doctorate)	0.040	0.019	0.046	0.035
grad diploma, grad certificate	0.098	0.080	0.059	0.048
Bachelor degree	0.220	0.144	0.164	0.159
advanced diploma, diploma	0.128	0.095	0.087	0.093
certificate iii or iv	0.113	0.122	0.300	0.159
certificate i or ii	0.064	0.073	0.037	0.053
certificate not defined	0.038	0.069	0.028	0.033
Year 12	0.119	0.144	0.103	0.229
Year 11	0.163	0.227	0.174	0.184
Missed high (base)	0.016	0.026	0.006	0.008
Number of person-year observations	*2174*	*1814*	*3671*	*397*

Appendix Table A.2 (a) WOMEN, Specification [3]

	Pooled OLS		Random Effect		Fixed Effect	
	Coefficient	t-stat	Coefficient	t-stat	Coefficient	t-stat
emppt	0.0171	1.38	0.0413***	3.14	0.1285***	5.92
onejob	-0.0413**	2.31	-0.0283	1.60	-0.0122	0.50
daywork	-0.0268**	2.09	-0.0212	1.58	-0.0129	0.62
contract	-0.0396**	2.19	-0.0460***	2.65	-0.0489**	2.16
casual	-0.0661***	4.37	-0.0596***	3.85	-0.0242	1.06
tempagency	0.0621**	1.97	0.0537*	1.79	0.0354	0.92
vic	-0.0662***	4.53	-0.0658***	3.73	0.0613	0.59
qld	-0.0789***	5.07	-0.0799***	4.25	-0.0496	0.48
sa	-0.1139***	5.61	-0.1105***	4.49	0.1527	0.73
wa	-0.0812***	3.92	-0.0798***	3.19	0.0162	0.10
tas	-0.0469	1.41	-0.0501	1.26	-0.0071	0.04
nt	0.0062	0.10	0.0309	0.43	0.3346	1.19
act	-0.0158	0.44	-0.0106	0.24	0.2434	0.88
marr	0.0529***	3.15	0.0450**	2.31	0.0336	0.63
cohab	0.0335*	1.73	0.0266	1.24	0.0083	0.20
wds	0.0361*	1.70	0.0294	1.20	0.0163	0.27
born_oz	0.0418**	2.40	0.0387*	1.83	(dropped)	
born_engsp	0.0708***	3.07	0.0697**	2.48	(dropped)	
urban	0.0358	0.88	0.0331	0.72	-0.0029	0.03
innreg	0.0005	0.01	0.0057	0.12	0.0181	0.18
outreg	-0.0161	0.37	-0.0240	0.50	-0.0448	0.46
tenure	0.0110***	4.80	0.0090***	3.58	-0.0017	0.38
tensq	-0.0002***	2.25	-0.0001	1.48	0.0001	0.42
hgage	0.0154***	2.59	0.0138*	1.91	-0.0568	1.10
agesq	-0.0002***	2.70	-0.0002**	2.03	0.0009	1.56
exper	0.0144***	4.17	0.0161***	3.84	-0.0329	0.78
expersq	-0.0003***	3.75	-0.0003***	3.40	-0.0015**	2.39
postgrad	0.2447***	7.25	0.2740***	6.81	0.2018	1.05
graddip	0.2339***	10.18	0.2513***	9.33	0.0096	0.09
bachelor	0.1666***	8.73	0.1809***	8.12	-0.1103	1.16
cert3or4	-0.1243***	5.79	-0.1076***	4.32	0.0170	0.22
cert1or2	-0.1667***	6.67	-0.1476***	5.05	-0.0350	0.38
certnd	-0.1015***	3.68	-0.0919***	2.82	-0.0967	0.78
year12	-0.0739***	3.51	-0.0546**	2.31	-0.0456	0.87
year11	-0.2042***	10.57	-0.1764***	8.16	0.0417	0.89
wave2	0.0196*	1.85	0.0195**	2.53	0.1013**	2.21
_cons	2.3869***	22.20	2.3783***	18.54	4.6331***	3.24

Notes: Breusch Pagan Lagrangian multiplier test for Random Effect model has the value of χ^2=420.17 which is highly significant at the 5% level significance. Hausman test statistics of χ^2=109.85 is highly significant at 5% level. t-value of each variables are reported below the corresponding coefficients. *, **, *** denote significant at 0.10, 0.05, and 0.01 level respectively.

Appendix Table A.2 (b) MEN, Specification [3]

	Pooled OLS		Random Effect		Fixed Effect	
	Coefficient	t-stat	Coefficient	t-stat	Coefficient	t-stat
emppt	**-0.0482****	1.97	0.0339	1.48	0.1522***	5.00
onejob	-0.0001	0.01	-0.0238	1.15	-0.0513**	1.98
daywork	-0.0551***	3.89	-0.0299**	2.13	0.0053	0.27
contract	0.0400**	1.93	0.0173	0.98	0.0087	0.42
casual	-0.0243	1.04	-0.0526***	2.57	-0.0547**	2.20
tempagency	0.0641*	1.85	0.1118***	3.75	0.1389***	3.92
vic	-0.0207	1.28	-0.0205	0.99	0.0626	0.61
qld	-0.0767***	4.50	-0.0874***	4.02	-0.3381***	2.76
sa	-0.1760***	7.69	-0.1740***	5.95	0.0919	0.51
wa	-0.0890***	4.03	-0.0895***	3.17	-0.0588	0.33
tas	-0.0925**	2.40	-0.1175***	2.41	-0.5248**	2.09
nt	-0.0619	0.67	-0.1007	0.93	-0.0895	0.34
act	0.0521	1.27	0.0399	0.78	-0.3175*	1.74
marr	0.1615***	8.79	0.1424***	6.56	0.0318	0.64
cohab	0.1291***	5.82	0.1026***	4.38	0.0396	1.06
Wds	0.1044***	3.60	0.1007***	3.12	0.0432	0.75
born_oz	0.0941***	4.80	0.0906***	3.58	(dropped)	
born_engsp	0.1683***	6.74	0.1715***	5.30	(dropped)	
urban	-0.1089**	2.49	-0.1135**	2.44	-0.2028***	2.60
innreg	-0.2053***	4.57	-0.1824***	3.86	-0.1913**	2.56
outreg	-0.2176***	4.61	-0.1889***	3.90	-0.1246*	1.74
tenure	0.0080***	3.47	0.0073***	2.84	0.0015	0.32
tensq	-0.0000	0.17	-0.0000	0.15	0.0001	0.58
hgage	0.0360***	4.13	0.0362***	3.26	0.0099	0.15
agesq	-0.0005***	4.78	-0.0005***	3.75	-0.0000	0.08
exper	0.0092*	1.95	0.0112*	1.86	-0.0044	0.11
expersq	-0.0000	0.09	-0.0001	0.39	-0.0006	0.78
postgrad	0.2365***	6.70	0.2584***	5.89	0.0382	0.23
graddip	0.1604***	5.05	0.1713***	4.38	-0.0882	0.67
bachelor	0.1384***	5.58	0.1488***	4.89	-0.0930	0.90
cert3or4	-0.1610***	7.14	-0.1419***	5.15	-0.0815	0.87
cert1or2	-0.2671***	7.38	-0.2523***	5.68	-0.2699**	2.11
certnd	-0.2330***	5.83	-0.2099***	4.30	-0.1607	1.20
year12	-0.1219***	4.53	-0.0972***	3.06	-0.1280*	1.75
year11	-0.2520***	10.36	-0.2179***	7.49	0.0889	1.22
wave2	0.0196*	1.66	0.0211***	3.02	0.0560	1.25
_cons	2.2624***	15.54	2.2336***	12.31	3.3471**	2.14

Notes: Breusch Pagan Lagrangian multiplier test for Random Effect model has the value of χ^2=837.99 which is highly significant at the 5% level significance. Hausman test statistics of χ^2=123.6 is highly significant at 5% level. t-value of each variables are reported below the corresponding coefficients. *, **, *** denote significant at 0.10, 0.05, and 0.01 level respectively.