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Public-Private Wage Gap in Australia: Variation Along the Distribution

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Abstract

Previous research on public-private wage differentials in Australia is scarce and has focused on the central parts of the conditional wage distribution. Using the first six waves of the Household, Income and Labour Dynamics in Australia (HILDA) survey, this study applies quantile regression models to examine whether the sectoral wage effect varies along the wage distribution. For females, we find public sector wage premiums for almost the entire wage distribution and the premiums are relatively stable except at the extremities of the distribution. For males, the premiums decrease monotonically and are negative for the top half of the conditional wage distribution. The decomposition results show that the observed differences in individuals and job characteristics account for a substantial proportion of the overall sectoral wage gap.

JEL Codes: J31, J45
Keywords: wage gap, quantile regression, decomposition

1. Introduction

There are many reasons why public and private sectors workers can be paid differently. First, public sector could set wages in a non-competitive way due to the monopolistic power of governments in setting prices and taxes for the provision of public services (Reder 1975). Second, public sector may be driven by objectives such as vote and/or budget maximisation rather than profit maximisation. Wages in public sector may also be used to achieve other considerations such as equity and fairness. Third, the institutional environment for wage setting may differ between public and private sectors. For example, there could be an imperfect labour market in the public sector. Union density is often higher in the public sector than in the private sector. Consequently, union may have a stronger bargaining power in securing higher wages for public sector employees in a collective bargaining industrial framework. Fourth, productivity-related characteristics of employees in the two sectors may be different. If public sector employees are relatively skilled, they require higher remunerations.

Study of public-private pay gap has important policy implications on a wide range of labour market issues. For example, higher wages to public sector employees may justify outsourcing of some government functions to private sector, and may potentially crowd out recruitment effort of private sector, forcing it to raise wages in order to compete for employees in the labour market.

Earlier studies on the public-private wage differentials mostly focus on the mean of the wage distribution. International evidence suggests that, relative to private sector, on average there is a wage premium for public sector (between 3 to 11 per cent) and the premium is often found to be higher for females than for males (See Borland and Gregory (1999) for a detailed review).[1] Recently, quantile regressions have been increasingly used to examine whether the public-private earnings differentials vary along the earnings distribution. The volume of this literature includes, for instance, Melly (2002) on Germany; Poterba and Rueben (1994) on the U.S.; Mueller (1998) on Canada; and Blackaby *et al.* (1999) on the U.K.; Lucifora and Meurs (2006) on Italy, France and the U.K.; Bonjour (1999) on Switzerland. Typically, these studies find lower pay dispersion in the public sector. Also, they find that public sector employees at the lower end of the wage distribution enjoy a wage premium relative to private

[1] However, Adamchik and Bedi (2000) find a private sector wage advantage in Poland.

sector employees; but the reverse holds for employees at the upper end of the wage distribution. In addition, female public sector employees are often found to enjoy a premium across almost the entire wage distribution; while male employees in the public sector suffer wage penalty over a large part of the wage distribution.

Only a few studies have examined public-private earnings differentials in Australia. Among the few, they give conflicting evidence about whether the observed earnings differentials are attributable to the sectoral effect. For instance, using the 1993 Training and Education Experience Survey collected by the Australian Bureau of Statistics, Borland *et al.* (1998) show that, relative to full-time private sector employees, the average weekly earnings of full-time public sector employees were 10 to 15 per cent higher for males and 20 to 25 per cent higher for females. However, they find that the differentials can all be explained by observed differences in productivity-related individual and job characteristics, suggesting there is no sectoral effect for Australian workers. On the other hand, using the 1985 Australian Longitudinal Survey data, Vella (1993) finds a significant wage premium for young female government employees aged 15 to 26 years relative to their private sector counterparts even after controlling for observed heterogeneity.

The Australian public sector has gone through significant reform and its size, measured in terms of employment, has reduced significantly since mid-1980s. The employment share of the public sector (including commonwealth, state/territory, and local governments) has dropped from 25.4 per cent in 1985 to about 16 per cent in 2005 (Krvger 2006). It is a result of a combination of privatisation, outsourcing, reduction of permanent employment, increase in part-time, causal and contract employment, and technological changes. Among the OECD countries, Australia and the U.K. are the only two countries that have significantly reduced the share of public sector. The public sector size in Australia is much smaller than other OECD countries such as Sweden and Denmark (over 30 per cent in 1998) (Jürges 2002), and the U.K. (about 20 per cent in 2005). Nonetheless, the Australian government remains a major employer, with over 1.3 million employees in 2006, approximately 18 per cent of the workforce (Industry Skills Councils 2006).

Australia has also been undergoing significant changes in industrial relations since the early 1990s. Through a ruling of the Australian Industrial Relations Commission, the Australian wage setting started to shift from industry-based awards towards enterprise-based (or workplace-based) agreements (Waddoups 2005). The introduction of the Workplace Relations Act (WRA) in 1996 further legitimised this practice. As a result of the WRA, the proportion of workers covered by the traditional award system has fallen dramatically. For example, in May 2000 only 23.2 per cent of employees were paid under an award compared to 67.6 per cent in May 1990 (Department of Employment and Workplace Relations 2002).[2] Reflecting these industrial relations reforms, the Public Service Act 1999 provides the most significant and extensive deregulation. For instance, it gives agency heads direct power to manage staff using merit principle to maximise agency performance, shifting the focus to individual agency for wage determinations (Australian Public Service Commission 2003).

This study contributes to the Australian literature on public-private pay gap in two ways. First, we examine wage differentials in a changed industrial relations environment where decentralised wages settings are more extensive than in the period covered by previous studies. We expect that the new industrial relations system may lead to differential sectoral effects across the conditional wage distribution. The experience of the U.K. suggests that decentralisation in wage setting and higher employer’s autonomy in wage determination have contributed larger public-private wage differences especially in the lower part of the wage distribution (Bender and Elliott 1999; Blackaby *et al.* 1999; and Disney and Gosling 2003). Focusing on public sector alone, Bender (2003) finds that the pay distribution has narrowed at the low end but has widened at the upper end after the first round of enterprise bargaining in Australia. As a result, pay inequality in the public sector has grown. This would have implications on the wage differentials between public and private sector workers. Second, using quantile regression models, we examine how the wage gap varies across the conditional wage distribution rather than only estimate the gap at the mean as earlier Australian studies did. The results from quantile regressions would provide a more complete description of the sectoral wage differentials.

The rest of the paper is arranged as follows. Section 2 describes quantile regression models and the semi-parametric decomposition method. Section 3 discusses the data

[2] In May 2000, 35.2 per cent of employees were on registered collective agreements, 1.5 per cent on unregistered collective agreements, and 40 per cent were covered by individual agreements (Department of Employment and Workplace Relations 2002).

source and model specification. Section 4 presents estimation results. Finally, in Section 5, we set out our conclusions.

2. Method

2.1 Quantile regression

To investigate whether the public-private pay gap vary at different points of the conditional wage distribution, we employ the quantile regression models of Koenker and Bassett (1978). Following Buchinsky (1998), we specify the θ^{th} ($0 < \theta < 1$) conditional quantile of the distribution of the (log) wage *w*, conditional on a vector of covariates *x*, as

(1) $Q_{\theta}(w \mid x) = x\beta(\theta)$.

Equation (1) assumes a linear relationship between the population conditional quantile of *w*, $Q_{\theta}(w \mid x)$, and the covariates *x*. For a random sample of (w_i, x_i) for $i = 1,..., N$, equation (1) implies

(2) $w_i = x_i\beta(\theta) + \varepsilon_{\theta_i}$, with $Q_{\theta}(\varepsilon_{\theta_i} \mid x)$ =0,

where ε_{θ_i} is the error term of the θ^{th} conditional (on x_i) quantile. In quantile regressions the only distributional assumption on ε_{θ_i} is that the θ^{th} conditional (on x_i) quantile of the error term equals zero.

For a given $\theta \in (0,1)$, $\beta(\theta)$ can be estimated by

(3) $$\hat{\beta}(\theta) = \arg\min_{\beta} \frac{1}{N}\sum_{i=1}^{N}(w_i - x_i\beta)(\theta - 1(w_i \le x_i\beta)),$$

where $1(\cdot)$ is the indicator function (Koenker and Bassett 1978). $\beta(\theta)$ is estimated separately for each $\theta \in (0,1)$.

Following the tradition, we first estimate a single equation quantile regression model of the form similar to equation (2),

(2') $w_i = \alpha(\theta)P_i + x_i\beta(\theta) + \varepsilon_{\theta_i}$, with $Q_{\theta}(\varepsilon_{\theta_i} \mid P_i, x_i)$ =0,

where P_i is a dummy variable equals to one if individual *i* works in public sector and zero otherwise; x_i is a vector of other variables that are expected to affect wages, such as education and experience. The quantile regression coefficients can be interpreted as the rates of return to the respective characteristics at the specific quantile of the conditional wage distribution (Buchinsky 1998; Koenker 2005). Therefore, $\alpha(\theta)$ measures the public sector wage premium (or penalty if it is negative) at the θ^{th} conditional quantile of wages and $\beta(\theta)$ measures the effect of other variables at that point of the conditional wage distribution. If the public sector wage premium is the same across the conditional wage distribution, we would expect $\alpha(\theta)$ not to vary for different θs. On the other hand, if being a public sector employee has no effect on wages, then $\alpha(\theta)$ should not be significantly different from zero for any θ.

The single equation model in equation (2') assumes that the wage determination process is identical for both public and private sector workers. However, test results shown later suggest that the assumption is violated: the wage determinants affect public and private workers differently. To account for the differences in the returns to wage determining factors between public and private sector workers, separate wage equations for each group are required. As in the OLS framework, after estimating the wage equation separately for public and private sector workers using quantile regressions, the differences at various quantiles of the wage distributions between the two groups of workers can be decomposed into the difference due to observed characteristics and the difference in returns to those characteristics.

2.2 Decomposition in quantile regression

A decomposition method for quantile regression models was initially developed by Machado and Mata (2005). Here we use a modified procedure proposed by Melly (2005) and Autor *et al.* (2005). In the modified procedure, instead of randomly drawing θ and *x*, we simply estimate quantile regressions for a large number of selected θs, such as $\theta_1, \theta_2, ..., \theta_J$, and use the observed sample *x* to form required marginal distributions of wages. In summary, the following steps are involved in decomposing the wage gap between public and private sector workers at different points of the wage distributions.

Step 1: Estimate $\beta^p(\tau_j)$ and $\beta^n(\tau_j)$, for $\tau_j \in (0,1)$ and $j = 1,..., J$, using the public sector workers and private sector workers respectively, to form $\{x_i^p \beta^p(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$ and $\{\{x_i^p \beta^n(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$, where x_i^p refers to the observed characteristics of public sector worker i; x_i^n refers to the observed characteristics of private sector worker i; N_p and N_n refer to the numbers of public and private sector workers respectively. $\{x_i^p \beta^p(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$ provide the predicted wage density of public sector employees; $\{\{x_i^p \beta^n(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$ provide the counterfactual wage density of public sector workers that would arise if they retained their own characteristics but were paid as private sector workers.

Step 2: Estimate the θ^{th} quantile of the sample $\{x_i^p \beta^p(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$, denoted as $Q_\theta(x^p, \beta^p(\tau))$, and of the sample $\{\{x_i^p \beta^n(\tau_j)\}_{j=1}^{J}\}_{i=1}^{N_p}$, denoted as $Q_\theta(x^p, \beta^n(\tau))$.

Step 3: Obtain $Q_\theta(x^p, \beta^p(\tau))$ - $Q_\theta(x^p, \beta^n(\tau))$. This difference represents the wage gap attributable to the differences in the returns to observed characteristics at the θ^{th} quantile, i.e. the public sector wage effects.[3]

To estimate the standard errors and confidence intervals of the sectoral wage effects, the bootstrap method can be used to replicate the above procedure. In this study 100 replications are carried out to estimate the confidence intervals and repeated observations for the same person in different waves (i.e. clustering) are taken into account in re-sampling.

3. Data and model specification

3.1 Data source

The empirical analysis is based on the first six waves (2001–2006) of the Household, Income and Labour Dynamics in Australia (HILDA) survey. The survey is a national household panel survey focusing on families, income, employment and well-being.[4] The first wave was conducted between August and December 2001. Then, 7683 households representing 66 per cent of all in-scope households were interviewed, generating a sample of 15,127 persons 15 years or older and eligible for interview. Of them, 13,969 were successfully interviewed. Subsequent interviews for later waves were conducted about one year apart.

[3] An alternative decomposition using $Q_\theta(x^n, \beta^p(\tau))$ and $Q_\theta(x^n, \beta^n(\tau))$ shows a similar result.

[4] Detailed documentation of the survey is in Wooden, Freidin and Watson (2002).

The HILDA survey contains detailed information on individuals' current labour market activity including labour force status, earnings and hours worked, and employment and unemployment history. For those employed, information on job characteristics, such as the size of the workplace and the industry to which the employee belongs is also collected. The wages used in this study refer to hourly wages derived from pre-tax total weekly earnings and hours worked in the main job.[5,6] To avoid the effect of irregular reporting of weekly earnings and hours worked, we excluded those whose hourly wage rate is less than \$5. One comparative advantage of HILDA is that the earnings data are not grouped, thus avoiding possible measurement error due to grouped data. To increase the sample size and thus the accuracy of the estimated distribution, we pool the six waves of HILDA survey currently available. Wages are deflated to the first quarter of 2001 using quarterly wage growth rates for males and females separately. Another reason for pooling the data is that sufficiently large sample sizes are important in bootstrapping the standard errors of the decomposition results.[7] Pooling six waves of HILDA raises two econometric issues. One relates to repeated observations, as most individuals are surveyed more than once. The other is an increase in real wages over time. We include year dummies and use bootstrap methods that account for clustering in the empirical work to address these issues.

Our sample includes those wage earners who worked in non-agricultural industries. It includes males aged between 25 and 64 (inclusive) years and females aged 25 to 61 (inclusive) years. Full-time students are excluded. There are 18,570 individuals: 9,713 males and 9,257 females. About 22 per cent of males and 33 per cent of females in the samples are public sector employees. The summary statistics of the samples are presented in the appendix Table A1.

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[5] We use hourly wages in this study to avoid complications arising from the potential effects of unions on hours worked (Andrews et al. 1998).

[6] Using hourly wages in all jobs produces virtually the same results.

[7] The bootstrapping method is difficult to carry out if the sample size is too small. It is because sampling draws did not always contain observations that had the characteristics used in the model if only one wave data were used. For example, since only a few private workers are indigenous in any one wave, a redrawn private worker sample may not have an indigenous worker. As a result, the original model that includes indigenous status as a covariate cannot be estimated using this redrawn sample. While STATA goes ahead to estimate βs by automatically dropping these variables, the number of variables for public and private samples, x^p *and* x^n respectively, will no longer be the same. As a result, one could not calculate the counterfactual wages of public sector employees in bootstrapping, since $x^p\beta^n$ becomes unconformable. Pooling the six waves of data helps to avoid the problem.

3.2 Distribution of wages

To have a better grasp of the wage distribution across sectors, we estimate the wage density using the kernel estimator and present the results in Figure 1. For both males and females, wages in the public sector appear to have a higher mean than in the private sector. For males the dispersion of wages in the private sector appears to be larger than in the public sector; the opposite is true for females.

Figure 1: Distribution of wages in the public and private sectors by sex

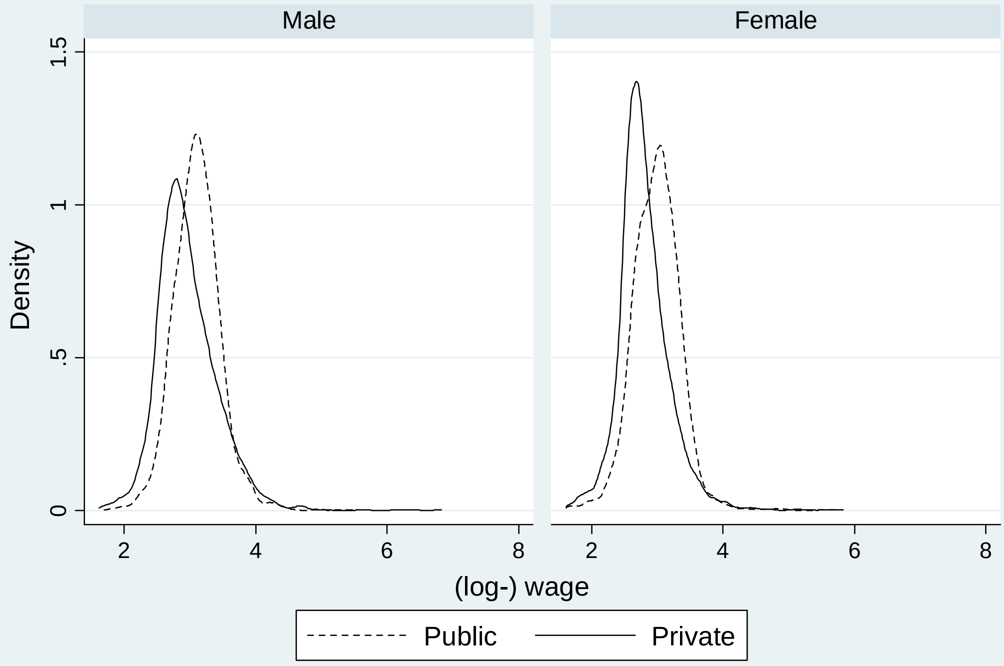
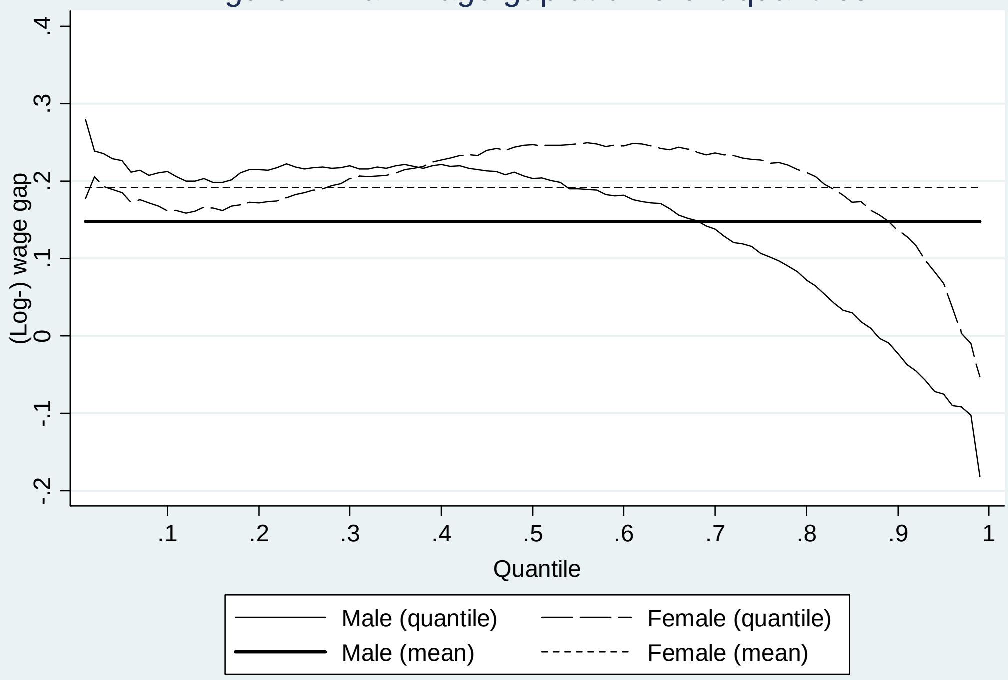


Figure 2 shows the raw wage gap between public and private sectors at different percentiles and at the mean. On average, (log) wages of male workers in the public sector are 15 percentage points higher than male workers in the private sector; female workers in the public sector have a wage 19 percentage points higher. These mean wage gaps are in line with that found in Borland et al (1998). Clearly, the gap is not uniform across the wage distribution. For males, the gap decreases from the bottom up to the 15^{th} quantile; it then increases until to the 23^{th} quantile and becomes relatively flat up to the 40^{th} quantile. After that it falls monotonically. The gap for females increases initially and then falls until to the 12^{th} quantile; from the 13^{th} quantile the gap increases up to about the 60^{th} quantile and thereafter falls monotonically. For males the gap is positive from the bottom up to about the 88^{th} quantile of the wage distribution; for females, the gap is positive for almost the entire wage distribution. The variation of the wage gap along the wage distribution provides a case for using quantile regressions to analyse the public-private wage differentials.

Figure 2: Raw wage gap at different quantiles



3.2 Model Specification

The specification of the wage equation is an extension of the standard Mincer model of wage determination (Mincer 1974). Essential to his model are human capital variables. Therefore we include in the wage equation four education dummies (degree, other post-school qualification, year 12, and year 11 and below), work experience (lifetime employment and its square) and a dummy on whether one has long-term health conditions (representing health capital). In addition to human capital, variables on the following characteristics are also included in the model: demographic characteristics (three dummies for whether one is born in Australia, an immigrant from an English speaking or an immigrant from a non-English speaking country; a race dummy to identify whether an individual is an Aborigine or Torres Strait Islander; and a marital status dummy); and employment characteristics (three dummies to identify casual, part-time or full-time employment); and three occupation dummies for white-collar workers (managers and professionals), other white-collar workers and blue-collar workers. To control for heterogeneity of local labour markets and the differential effects of regional living costs on wages, we also include six state dummies and a dummy indicating capital city residence. There are six dummies to identify workplace size ranging from less than 20 to over 500 employees. The positive relationship between workplace size and wages is well documented (Idson and Feaster 1990; Morissette 1993; Miller and Mulvey 1996). Increasing monitoring costs (which result in higher wages according to efficiency wage theories), greater importance of workplace-specific human capital and teamwork are some explanations discussed in the literature. A union membership dummy variable is used to capture the union wage effect. A positive relationship between union membership and wages is often found in the literature. Finally, year dummies are included to control for the trend of increasing real wages over the six waves of the HILDA data.

Summary statistics for the variables used are presented in appendix Table A1. The sample means reveal very little that is not already well known. For instance, public sector employees enjoy higher wages; larger workplaces (generally) have a higher incidence of public sector employees; public sector employees tend to participate in the workforce longer and more educated; have white-collar type of occupation; tend to be union members; are less likely to be migrants from non-English speaking countries; are more likely to be from New South Wales or the Australian Capital Territory and Victoria, but are less likely to hold casual and part-time jobs. There is some evidence of gender differences. As expected, more females have casual or part-time jobs. This is especially apparent among private sector workers. Also, more females are degree holders and with white-collar jobs than are their male counterparts irrespective of which sector they are employed. More female public sector employees are immigrants from non-English speaking countries.

3.3 Econometric issues

The estimation of a sectoral wage gap typically involves two complications resulting from two selection processes. One is the problem of sample selection arising from the work choice decision; the other is the selection into different sectors. If these two selection processes are determined by some unobserved factors that also affect wages, the public-private wage differentials estimated from models that do not account for these possibilities are likely to be biased. Our approach for accounting for sample selection in quantile regressions follows Buchinsky (1998, 2001). That is, we first estimate a single index selection equation using semi-parametric procedures (Frölich 2006; Klein and Spady 1993); a power series of the predicted index is then included in the wage equation.[8] In our case we found two terms were sufficient to account for sample selection. However, we could not account for the potential endogenous sector selection of workers due to lack of valid identifying instruments.[9] Accordingly, the results reported here must be interpreted with caution.

4. Results

4.1 Single equation estimation

Figure 3 presents the coefficient estimates and their 95 per cent confidence intervals for the public sector dummy variable from both the OLS model and quantile regressions. The quantile regressions are estimated at each 0.01 percentile point. For ease of reading, Table 1 lists the coefficient estimates for the public sector dummy at selected percentiles and also the estimates from OLS for males and females.[10]

The OLS estimates show that male workers in the public sector earn a wage that is 3 per cent lower than their counterparts in the private sector, while for female workers in the public sector their wages are about 4 per cent higher than female workers in the private sector. The OLS results are comparable with other studies. Take the U.K. as an example, public servants enjoy a wage premium of 5 per cent on average relative to comparable private sector employees (e.g. Rees and Shah 1995). It ranges from 2 to 5 per cent for males; but much higher for females (15-18 per cent).

[8] The results of the semi-parametric estimation of the selection equation are reported in Appendix Table 3, together with the probit estimates for comparision.

[9] Dustmann and van Soest (1998) and Melly (2005) use the father's public sector employment as an instrument. Goddeeris (1998) use political activities in college and self-reported political orientation to control sample selection. Unfortunately, HILDA does not collect such information. Occupation of parents when the person was 14 years is available in the data. We attempted to use parental occupation as instruments, but found parental occupation was generally insignificant in explaining males' sector choice, while for females mother's, but not fathers', occupation was sometimes significant. In addition, we are sceptical about the validity of parental occupation as instruments. Parental occupation is likely to be affected by unobserved ability which, in turns, is likely to be highly correlated between parents and their children.

[10] Coefficient estimates for other variables are not shown here but are obtainable on request from the authors.

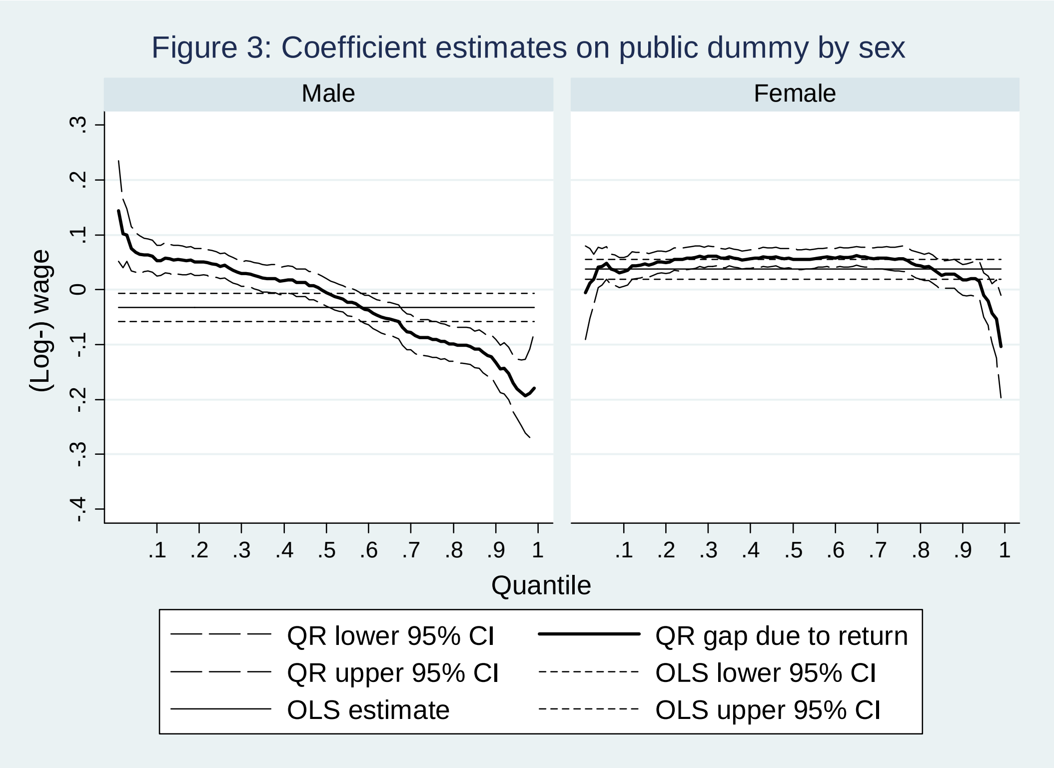


Table 1: Estimates of public-private wage gap in a single equation model

	Males		Females	
	Coef	s.e.	Coef	s.e.
OLS	-0.0321**	0.0133	0.0387***	0.0095
Quantile regression				
0.1	0.0529***	0.0143	0.0333**	0.0137
0.2	0.0507***	0.0124	0.0517***	0.0107
0.3	0.0293**	0.0117	0.0632***	0.0095
0.4	0.0170	0.0126	0.0589***	0.0082
0.5	-0.0052	0.0129	0.0553***	0.0089
0.6	-0.0368***	0.0137	0.0615***	0.0089
0.7	-0.0777***	0.0165	0.0578***	0.0098
0.8	-0.0985***	0.0164	0.0461***	0.0123
0.9	-0.1322***	0.0210	0.0199	0.0153

The story is quite different from the quantile regression estimates. For males the sectoral effect exhibits a monotonic decrease across almost the entire conditional wage distribution. For the lower one third of the conditional wage distribution, a positive effect is found, while for the upper 60 per cent of the conditional wage distribution, the effect is found to be negative. The magnitude of the negative effect is fairly large at the upper end of the conditional wage distribution. For example, in the top 20 per cent of the wage distribution, the negative effect is about 10 per cent or more. The OLS estimate is close to the estimate at the 60th quantile, but is far off those at the bottom and top ends of the conditional wage distribution. For females the quantile regression estimates are relatively stable and positive over almost the entire wage distribution and in the range of 3 to 6 per cent. The estimates at the very bottom and top ends of the conditional wage distribution are insignificant. Again, the OLS estimate for females provide misleading inference as to the effect at other parts of the conditional wage distribution. Using German data, Jürges (2002) also find that in contrast to males, female wage earners in the public sector enjoy a positive wage premium. A negative wage premium is only observed at very high quantiles. Poterba and Rueben (1994) estimate a single log wage equation with a public sector dummy using quantile regressions for the U.S.. They also report negative public sector wage premiums at the upper tail of the wage distribution, while a positive premium is evident at the lower end.

4.2 Quantile regression decomposition

The single equation estimation results must be interpreted with caution, because they rely on the assumption that the wage determination process is identical for both public and private sector workers. This assumption may be violated if being a public sector employee also affects the returns to factors such as education. To see whether the model should be estimated separately for each group of workers, we experimented through making interactions of each independent variable with the public sector dummy. If the interaction terms are jointly significant, the independent variables affect public and private sector workers differently. The test statistics are reported in appendix Table A2. The results in general reject the hypothesis that workers in both sectors are subject to the same wage determination process. Therefore, the sectoral wage effects estimated using the single equation model are likely to be misleading; separate wage determination equations for public and private sector and decomposition methods are required to provide a more reliable picture of the public-private sectoral wage differentials.

To generate the samples for decomposition purposes, we estimate models for quantiles at [0.001, 0.003... 0.997, 0.999] and at the median. There are 501 regressions for each gender and sector group and thus it is not possible to report all the estimation results.[11] In the followings we focus on the decomposition results.

[11] Selected quantile regression results, together with OLS estimates, can be obtained from the authors.

Using the procedure described in Section 2, we decompose the difference between the quantiles of the distribution into the components explained by the differences in inter-sectoral distribution of characteristics (e.g. personal and job characteristics) and different returns to these characteristics. It is the latter component that can be interpreted as the sectoral effects, because otherwise there should not be any difference in the returns. For this reason the reported results focus on the gap due to return differences.

Figure 3 shows the wage gap attributable to return differences at each 0.01 percentile point, together with bootstrapped 95 per cent confidence intervals. In bootstrapping the 95 per cent confidence intervals, 100 replications were used and the clustering of the observations resulting from the panel data was also taken into account. For comparison, the horizontal line in the figure shows the sectoral wage effect estimated using the OLS decomposition method. The OLS estimate is computed as $\bar{x}^p(\beta^p - \beta^n)$, using the Blinder-Oaxaca decomposition (Blinder, 1973; Oaxaca, 1973), where $\bar{x}^p$ refers to the means of the public sector worker sample; β^p and β^n refer to the OLS coefficient estimates from public sector workers and private sector workers respectively. Again, for ease of reading Table 2 presents the results for selected quantiles.

The OLS decomposition shows that for male workers the contribution of returns differences to the wage gap is -0.05, which is significant at the five per cent significance level, implying that male workers in the public sector earn 5 per cent less on average than a comparable worker in the private sector. This estimate is larger in size than that found in a single equation model. For females, the OLS decomposition shows that the sectoral wage effect is about 0.03, lower than that found in the single equation model. For both males and females our OLS decomposition results are different from Boland e*t al.* (1998). They find that the entire observed wage gap can be attributed to the differences in observed individual and job characteristics. The differences of the results might not come as a surprise for at least two reasons. First, Borland *et al.* (1998) use data collected in 1993, about 10 years earlier than the data used in this study. In 1993 wage setting was very much controlled by the award system, but the data in this study covered the period when enterprise-bargaining is wide-spread. Second, the macroeconomic conditions are very different between the two periods. Year 1993 was the time when recession hits the bottom with an unemployment rate of over 12 per cent. The period studied here is characterised by a booming economy with an unemployment rate less than a half of the 1993 level. As a result, the relative wage structure between public and private sectors might have changed, provided that the two sectors have responded to the economic boom differently in terms of how wages are set to attract skilled workers.

Figure 4: Wage gap due to difference in returns by sex

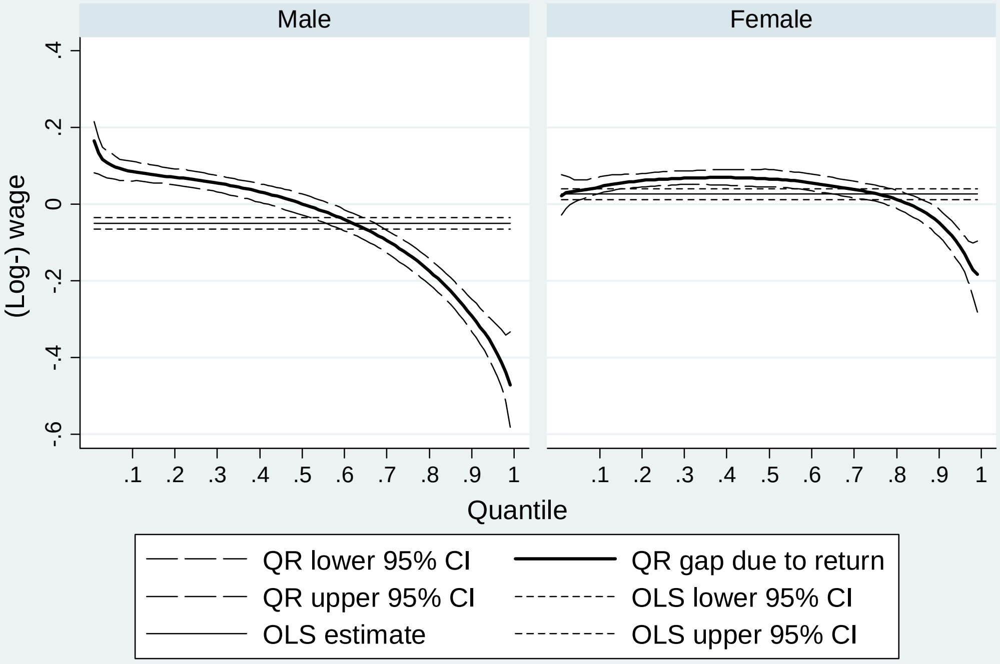


Table 2: Decomposition of the public-private wage gap

	Males			Females		
	Raw gap	Gap due to diff in returns		Raw gap	Gap due to diff in returns	
		Estimates	95% CI		Estimates	95% CI
OLS	0.1483	-0.0493	(-0.0641, -0.0344)	0.1917	0.0263	(0.0122,0.0404)
Quantile Regression						
0.1	0.2125	0.0853	(0.0610, 0.1121)	0.1619	0.0455	(0.0287, 0.0713)
0.2	0.2146	0.0701	(0.0502, 0.0927)	0.1716	0.0617	(0.0454, 0.0799)
0.3	0.2196	0.0552	(0.0085, 0.0519)	0.2031	0.0679	(0.0519, 0.0874)
0.4	0.2214	0.0327	(-0.0049, 0.0535)	0.2273	0.0698	(0.0495, 0.0908)
0.5	0.2030	0.0010	(-0.0277, 0.0264)	0.2469	0.0659	(0.0458, 0.0909)
0.6	0.1819	-0.0397	(-0.0695, -0.0141)	0.2454	0.0557	(0.0349, 0.0790)
0.7	0.1380	-0.0950	(-0.1270, -0.0675)	0.2367	0.0384	(0.0166, 0.0598)
0.8	0.0721	-0.1744	(-0.2091, -0.1423)	0.2115	0.0125	(-0.0106, 0.0369)
0.9	-0.0235	-0.2922	(-0.3325, -0.2477)	0.1364	-0.0483	(-0.0851, -0.0132)

The quantile regression decomposition results show that the sectoral effects are positive for the quantiles from the bottom up to the 50th quantile for male workers and up to the 82nd quantile for females. The positive effects are significant at the 5 per cent significant level for the quantiles from the bottom up to the 41st for males and for the quantiles from the 4th to the 77th for females. For males the significant positive effect decreases monotonically from 16 per cent at the bottom to about 3 at the 41st quantile; for females the significant positive effect initially increases from 3 per cent at the 4th quantile to about 7 per cent at the 40th quantile and falls thereafter to about 2 per cent at the 77th quantile. For males the negative effect becomes significant from the 58th quantile onwards and the significant negative effect increases from 3 per cent to 47 per cent. For females the negative effect becomes significant from the 89th quantile onwards; the significant negative effect increases from about 4 per cent to about 18 per cent. Similar effect patterns are found for German workers (Jürges 2002; Melly 2004). Clearly, the estimates from the OLS model cannot reveal the variation of the sectoral effect across the wage distribution, as found from the quantile regression models. In particular, the opposite sectoral effects at the lower end and the upper end of the wage distribution cannot be inferred from the OLS models.

Table 2 also shows that the part of wage gap due to differences in observed individual and job characteristic is substantial. The proportion of the observed wage gap attributable to the sectoral effect is relatively small. This suggests that public sector employees have individual and job characteristics that are more conducive to higher remuneration. In Table 2 the quantile where the largest proportion of the gap (40 per cent for males and 36 per cent for females) can be attributed to the sectoral effect is the 10th and the 20th quantiles for males and females respectively. This finding is in line with that of Melly (2005).

Comparing Figure 3 with Figure 2 and Table 2 with Table 1, we find that the patterns of the estimated effects are similar between the single equation and separate equation models, but the magnitude of the estimated effects differs. The estimates from the separate equation models are generally larger than that in the single equation models.

5. Conclusion

Using the first six waves of the HILDA survey, this paper employs both OLS and quantile regressions and a semi-parametric decomposition method to examine the sectoral wage gap at the mean and over the entire conditional wage distribution. Unlike earlier Australian studies, using OLS models, we found a significant negative sectoral effect for males and a significant positive effect for females after controlling for observed individual and job characteristics, although the size of the effect is small.

Using quantile regressions, we found a significant wage premium for the public sector at the lower part of the conditional wage distribution and a significant wage penalty at the upper part of the distribution, a result similar to a number of international studies. The public wage premium at the lower end of the conditional wage distribution might be due to the more effective implementation of equal opportunity and anti-discriminatory policies in the public sector, since the government may use public sector pay to achieve objectives such as equity and to be a 'good' employer (Bender and Elliot 1999). A commonly cited reason for public wage penalty at the upper end of the conditional wage distribution is public opposition to high pay for public servants (Katz and Kreuger 1991; Lucifora and Meurs 1999), while private sector is not subject to such opposition. This allows the private sector to use high pay to attract high-skilled workers. Higher private sector remuneration could also be compensating differentials that private sector employers use to reduce the turnover rate of high skilled employees and/or for less pleasant work environment. For example, some studies find that overall satisfaction in the public sector is higher than in the private sector (Gardner and Oswald 1999; Jürges 2001). Nation-wide skill shortage and the booming economy may be another reason for the much higher wages in private sector than in public sector at the upper end of the wage distribution. Bargain and Melly (2008) also find a positive effect of the economic upturn on private sector wages in France. They attribute the sectoral wage differential to the sensitivity of private sector wage (and the lack of sensitivity of the public sector) to macro shocks. The currently booming Australian economy is largely driven by rapid increases in export of raw material and commodities produced by private sector. The booming of the mining and related industries not only creates high demand for high-skilled workers, but also generates large revenue for the industries. This means that these private sector employers can afford to pay high wages to attract employees needed.

It is not clear why female private sector employees are only paid more than their public sector counterparts at the very top end, whereas male private employees are rewarded higher than their public sector counterparts for a larger part of the wage distribution. One possible explanation is that labour market discrimination against women is more widespread in the private sector than in the public sector, affecting most except a few women at the very top end. The alternative explanation could be that the distributional differences of men and women across industries in the private sector may lead to different patterns of the sectoral wage effect between men and women if more men are in high-pay industries. Finally, different unobservables between men and women could also be an attributing factor. For instance, relative to men, women in general may not be good at bargaining for themselves (Babcock and Lashever 2003). Only the few female executives who have acquired the bargaining skills gain higher pays as the specific salary levels in the private sector are more likely to be determined by negotiation. The exact reasons for the difference of the sectoral wage effects between males and females require further investigation.

The decomposition results indicate that differences in observed characteristic explain a substantial proportion of the overall public-private sector wage gap. The sectoral effect only accounts for a relatively small proportion with its impact mostly confined to the lowest end of the conditional wage distribution. In other words, public sector employees have characteristics that are more conducive to higher remuneration.

This study has limitations. First, due to the data constraint the problem of selection into different sectors could not be dealt with here. If wages and selection into a particular sector are affected by some correlated unobservables, the estimates reported here might be biased. By not allowing endogenous sector choices, we may underestimate the mean premium as shown by Melly (2005). Second, the quantile regression results rely on the assumption that the covariates, particular sectoral status, are not related to the mean of the unobservables. The estimated sectoral wage effects would be biased if this assumption does not hold. Third, as our data have no information on work effort (often lower in the public sector) and non-wage benefits (often higher in the public sector), our results are likely to underestimate the true public sector premium. Fourth, Chatterji *et al.* (2007) find that workplace characteristics such as presence of performance related pay, company pension schemes and family-friendly employment practices (eg. paternity leave and maternity leave with pay) are important in explaining the public-private wage gap. Due to data availability, we could not include job related characteristics other than workplace size. Finally, we do not distinguish public sector employees employed by the federal government from those employed by the state or local governments. Large wage differences could exist between different levels of government employees (Poterba and Rueben 1994).

Appendix

Table A1: Summary statistics of the samples

	Male						Female					
	Public		Private		All		Public		Private		All	
	Mean	Std	Mean	Std	Mean	Std	Mean	Std	Mean	Std	Mean	Std
Hourly wages	23.816	10.628	21.576	17.234	22.153	15.830	21.229	10.121	17.831	10.642	19.035	10.585
Married	0.820	0.384	0.742	0.438	0.762	0.426	0.718	0.450	0.714	0.452	0.715	0.451
Degree	0.431	0.495	0.221	0.415	0.275	0.447	0.495	0.500	0.237	0.425	0.328	0.470
Other post-school qualification	0.357	0.479	0.427	0.495	0.409	0.492	0.236	0.424	0.277	0.448	0.262	0.440
Year 12	0.088	0.284	0.118	0.323	0.111	0.314	0.098	0.297	0.165	0.371	0.141	0.348
Life-time employment	24.257	10.170	21.812	10.874	22.441	10.750	20.409	8.933	18.721	9.552	19.319	9.372
Life-time employment square	691.826	507.308	593.999	531.429	619.176	527.048	496.310	393.958	441.706	409.431	461.064	404.843
Indigenous	0.015	0.123	0.009	0.093	0.010	0.101	0.023	0.148	0.009	0.095	0.014	0.117
Immigrants from Eng-speaking country	0.115	0.319	0.118	0.323	0.118	0.322	0.094	0.292	0.113	0.317	0.106	0.308
Immigrants from non-Eng speaking country	0.087	0.282	0.128	0.335	0.118	0.322	0.094	0.292	0.132	0.339	0.119	0.323
NSW/ACT	0.371	0.483	0.304	0.460	0.321	0.467	0.339	0.473	0.320	0.466	0.327	0.469
VIC	0.226	0.418	0.267	0.442	0.256	0.437	0.234	0.424	0.266	0.442	0.255	0.436
QLD	0.200	0.400	0.208	0.406	0.206	0.404	0.200	0.400	0.209	0.407	0.206	0.404
SA	0.081	0.272	0.085	0.279	0.084	0.277	0.085	0.279	0.084	0.278	0.085	0.279
WA/NT	0.100	0.300	0.111	0.314	0.108	0.310	0.099	0.299	0.093	0.290	0.095	0.293
TAS	0.022	0.147	0.026	0.159	0.025	0.156	0.042	0.201	0.027	0.163	0.033	0.178
Capital city	0.595	0.491	0.650	0.477	0.636	0.481	0.586	0.493	0.646	0.478	0.625	0.484
Part-time	0.081	0.273	0.081	0.273	0.081	0.273	0.376	0.484	0.448	0.497	0.422	0.494
Casual	0.066	0.247	0.155	0.362	0.132	0.339	0.123	0.328	0.293	0.455	0.233	0.423
Part-time & casual	0.040	0.196	0.059	0.236	0.054	0.226	0.107	0.309	0.237	0.425	0.191	0.393
White collar workers	0.466	0.499	0.275	0.447	0.324	0.468	0.564	0.496	0.252	0.434	0.363	0.481
Other white collar workers	0.357	0.479	0.282	0.450	0.301	0.459	0.404	0.491	0.615	0.487	0.540	0.498
Blue collar workers	0.177	0.382	0.442	0.497	0.374	0.484	0.032	0.177	0.133	0.340	0.097	0.297

Having health conditions	0.176	0.381	0.149	0.356	0.156	0.363	0.151	0.358	0.140	0.347	0.144	0.351
Firm size <20	0.188	0.391	0.367	0.482	0.321	0.467	0.176	0.381	0.452	0.498	0.354	0.478
firm size 20-99	0.306	0.461	0.311	0.463	0.310	0.462	0.372	0.484	0.279	0.448	0.312	0.463
Firm size 100-199	0.139	0.346	0.109	0.312	0.117	0.321	0.101	0.301	0.099	0.299	0.100	0.300
Firm size 200-499	0.130	0.336	0.106	0.307	0.112	0.315	0.108	0.311	0.089	0.285	0.096	0.294
Firm size 500+	0.230	0.421	0.101	0.302	0.134	0.341	0.228	0.420	0.070	0.255	0.126	0.332
Firm size unknown	0.007	0.086	0.006	0.079	0.007	0.081	0.014	0.117	0.011	0.104	0.012	0.109
Union workers	0.601	0.490	0.264	0.441	0.351	0.477	0.540	0.498	0.205	0.404	0.324	0.468
Wave 2	0.170	0.376	0.169	0.375	0.170	0.375	0.164	0.370	0.166	0.372	0.165	0.372
Wave 3	0.168	0.374	0.166	0.372	0.166	0.372	0.169	0.374	0.162	0.369	0.165	0.371
Wave 4	0.164	0.370	0.159	0.365	0.160	0.367	0.159	0.365	0.158	0.365	0.158	0.365
Wave 5	0.163	0.370	0.163	0.370	0.163	0.370	0.166	0.372	0.165	0.371	0.165	0.372
Wave 6	0.160	0.367	0.166	0.372	0.165	0.371	0.170	0.375	0.171	0.376	0.170	0.376
Selection index	0.369	0.102	0.377	0.095	0.375	0.097	0.306	0.101	0.336	0.087	0.326	0.093
Selection index square	0.146	0.064	0.151	0.062	0.150	0.063	0.104	0.055	0.121	0.054	0.115	0.055
No. of observations	3663		10570		14233		5058		9209		14267	

Table A2: F-statistics on the joint significance of the interaction variables between public status dummy and other independent variables

	10%	20%	30%	40%	50%	60%	70%	80%	90%	OLS
Males	5.85***	5.11***	7.57***	10.61***	10.55***	13.02***	15.94***	14.61***	10.28***	5.30***
Females	1.24	2.72***	3.19***	4.09***	5.84***	6.45***	5.02***	5.99***	5.31***	2.54***

Note: *** denotes significant at the 1% level; ** 5% level and * 10% level.

Table A3: Probit and Semi-parametric estimations of the selection equation

	Males				Females			
	Probit		Semi-parametric		Probit		Semi-parametric	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Age	-0.6432	0.0772	-0.7918	0.2827	-0.5758	0.0533	-0.8132	0.2646
Age square	-0.2461	0.0211	-0.2520	0.0885	-0.0577	0.0162	-0.0666	0.0301
Married	0.3626	0.0321	0.3632	0.1204	0.0234	0.0233	0.0239	0.0339
Degree	1.0252	0.0418	1.1271	0.3813	0.7464	0.0282	0.9506	0.2882
Other post-school qualification	0.2911	0.0313	0.2671	0.0963	0.2816	0.0257	0.3727	0.1188
Year 12	0.4506	0.0478	0.4383	0.1565	0.2627	0.0313	0.3414	0.1128
Life-time employment	1.167	0.0724	1.3155	0.4528	1.5187	0.0425	2.0306	0.6211
Life-time employment square	0.032	0.014	0.0323	0.0191	-0.1808	0.0099	-0.2391	0.0739
Immigrants from Eng-speaking country	0.0066	0.0412	0.0432	0.0478	-0.1138	0.0334	-0.1477	0.0653
Immigrants from non-Eng speaking country	-0.1688	0.0396	-0.1367	0.0623	-0.3042	0.0298	-0.3762	0.1185
VIC	-0.0101	0.0357	-0.0496	0.0429	0.0094	0.0269	0.0144	0.0386
QLD	-0.0819	0.0374	-0.0634	0.0468	-0.0918	0.0284	-0.1167	0.0525
SA	-0.1369	0.0483	-0.1836	0.0812	-0.1013	0.0374	-0.1383	0.0669
WA/NT	0.0488	0.0482	0.0519	0.0577	-0.1729	0.0358	-0.2324	0.0868
TAS	-0.1569	0.0748	-0.1526	0.0940	0.1063	0.0601	0.1551	0.0993
Capital city	0.2273	0.0288	0.2668	0.0946	0.0287	0.0218	0.0363	0.0322
Having health conditions	-0.8843	0.0277	-0.8258	0.2810	-0.6761	0.0243	-0.8818	0.2693
Have children under 5 years	-0.0408	0.0528	-0.0433	0.0642	-0.9941	0.0324	-1.2968	0.3991
Have children 5-14 years	0.096	0.0536	0.1019	0.0712	-0.0103	0.033	-0.0280	0.0484
Number of children	-0.0868	0.0246	-0.1024	0.0438	-0.1352	0.0159	-0.1712	0.0571
Aged 55 or over	-0.3283	0.0658	-0.2948	0.1232	-0.2795	0.0539	-0.3697	0.1378
Wave 2	0.0662	0.0422	0.0643	0.0507	-0.0105	0.0332	0.0093	0.0477
Wave 3	0.2091	0.0431	0.2247	0.0887	0.0582	0.0335	0.0865	0.0541
Wave 4	0.2644	0.0443	0.2954	0.1103	0.065	0.034	0.0894	0.0549
Wave 5	0.3516	0.0449	0.3710	0.1332	0.1563	0.0342	0.2124	0.0787
Wave 6	0.3657	0.0451	0.3984	0.1412	0.1976	0.0343	0.2616	0.0907
Constant	-0.199	0.0688	.	.	-0.001	0.0522	.	.
Non-labour income	-1	.	-1	.	-1	.	-1	.

Note: For identification the constant term cannot be included in the semi-parametric estimation and the coefficient on one of the continuous variables has to be normalised to be one. Here we normalised the coefficient on non-labour income to be -1. For comparison, we also normalised the coefficient of non-labour income to be -1 in the probit model.

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