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The public sector pay premium, compensating differentials and unions: propensity score matching evidence from Australia, Canada, Great Britain and the United States

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Abstract

Propensity score matching is used to estimate the size of the public sector pay premium in four countries. Three sets of matching covariates are used; worker characteristics only, then including job attributes and finally adding union membership. When worker characteristics and job attributes are controlled for, the public sector pay premium ranges from 30% in Canada to 19-20% in Australia and Great Britain and only 6% in the United States. Differences in job attributes between private sector and public sector workers make almost no difference to the estimated pay premium. But once differences in union membership across sectors are controlled for, the estimated public sector pay premium is reduced in all countries and disappears in Canada. This finding favors the hypothesis that the pay premium partially reflects rents accruing to public sector workers, obtained most probably with assistance from the actions of their labor unions.

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1. Introduction

Workers in the public sector typically are paid more than those in the private sector. Some of this pay gap reflects differences in characteristics, such as higher average education, but non-competitive rents also may be accruing to public sector workers (Gregory and Borland, 1999). Such rents are possible because the public sector is now much more heavily unionized than the private sector (Visser, 2006) and public sector employers may not face the same hard budget constraint and competitive conditions that inhibit wage rises in many private firms. But it is also claimed that different job attributes account for much of the pay gap, at least in the U.K. (Bender and Elliott, 2002). This claim is contentious because many studies find better job attributes, such as greater job security and more generous pension schemes, for public sector workers (e.g., Poterba and Rueben, 1998). Compensating differentials should imply lower public sector pay to offset these more favorable job conditions.

To help inform this debate, this note reports propensity score matching (PSM) estimates of the public sector pay premium for Australia, Canada, Great Britain and the United States. Attention is restricted to these four English-speaking countries because previous evidence on wage differentials and sorting by worker quality across establishment sizes finds that sorting is mainly evident in English-speaking countries (Gibson and Stillman, 2009). Since establishments in the public sector are typically larger than those in the private sector (Belman and Heywood, 1990) worker quality differences across sectors are also likely, especially in the English-speaking countries, making these countries the most interesting ones for testing explanations about public sector pay.

The data come from the 2005 ISSP Work Orientations survey. This survey is uniquely suited to this task since in addition to recording standard worker characteristics like age, education and gender it also records job attributes such as whether the job is secure, interesting, improves skills, causes stress or exhaustion, is dangerous, and so on. For employed respondents the survey also records whether they are currently members of a union. While the survey does not record establishment size, job attributes such as whether the worker feels that they can work independently may proxy for size of the workforce.

The strategy in this note is to report three sets of PSM estimates, first matching just on worker characteristics,[1] then also including these new measures of job attributes in the matching, and finally including union membership as a matching covariate. If the estimated public sector pay premium persists when comparing earnings of observationally similar public sector and private sector workers who are doing jobs with observationally similar attributes, it would tend to count against the claim that much of the pay gap is due to differences in job attributes. If the estimated premium then falls once union membership is added to the worker characteristics and job attributes in the set of matching covariates, it would support the claim that the pay premium partially reflects rents accruing to public sector workers, in this case through the actions of their labor unions.

2. Data Description

ISSP surveys are carried out each year in approximately 30 countries, with a common core of questions asked of a probability-based, nationwide sample of 1000-2000 adults. An additional

[1] Previous research on the public sector pay premium using ISSP data on worker characteristics is reported by Gregory and Borland (1999). Those estimates come from OLS regressions which is more prone to self-selection bias than PSM, at least if selection is based on observables.

rotating module of the survey changes topic each year, with the work orientations module that records job attributes last asked in 2005. The job attributes are measured using five-point Likert scales. Previous international comparisons of the public sector pay premium have used ISSP data (Gregory and Borland, 1999) but only from the core questionnaire that records worker characteristics but not job attributes.

A description of the variables measuring worker characteristics, job attributes and union membership is reported in Table 1. The sample is restricted to those respondents that were working for pay at the time of the 2005 survey. There are a number of significant differences between public sector workers and private sector workers in the reported variables. In all four countries the public sector workers have significantly more years of education and are less likely to be male. Public sector workers are also older but only significantly so in Australia.

A consistent pattern in all four countries is that public sector workers consider their jobs more secure and also consider them as more helpful to others and more useful to society (in each case as evaluated by the worker themselves). Except for Canada, the public sector workers are significantly more likely to consider their job interesting and that it helps to improve their skills. In both Australia and the United States, public sector workers are significantly less likely to report that their job involves hard physical work. Offsetting these positive attributes of jobs, public sector workers outside of Great Britain report that they are significantly less able to work independently, and in Great Britain and the United States that they are significantly more likely to find their job is stressful.

The union membership rate is much higher in the public sector. In the United States only 9% of the private sector workers in the ISSP sample were current union members, and in the three other countries the private sector unionization rate ranged from 17-19%. In contrast, just over one-third of the ISSP employees in the public sector in the United States were current union members and approximately two-thirds were union members in the other three countries. The public sector unionization rate appears especially high in Canada, at 70.3%.

The ISSP also asks respondents to report pre-tax yearly personal income from all sources. While there is no question on earnings, for the respondents who are currently working most of their annual personal income should come from labor earnings. The logarithm of annual income is therefore used as the proxy measure of pay in this study. This same proxy is used in other international comparative studies using ISSP data (e.g., Gregory and Borland, 1999). According to this proxy, the raw pay gap for working in the public sector averages 24.7% across the four countries, ranging from 32.5% in Canada to (a statistically insignificant) 8.3% in Great Britain.[2]

3. Estimation Method and Specification

Let Y_{1i} and Y_{0i} be the log annual earnings of a randomly chosen worker in two counterfactual situations of treatment $(T_i = 1)$ and non-treatment $(T_i = 0)$, where the treatment in this case is working in the public sector. For a set of observable covariates, X_i let $P(X_i)$ be the propensity score, defined as: $P(X_i) \equiv \Pr(T_i = 1 | X_i)$; in this case the conditional probability of working in the public sector. The average treatment effect on the treated (ATT) can be estimated as:

[2] This is calculated from the difference in the logarithm of pre-tax annual incomes. For example, for Canada this is 10.8284-10.5472=0.2812. The percentage difference is then: 100×[exp (0.2812) – 1] = 32.5%.

$$
\begin{aligned}
ATT &\equiv E\left\{E\left[Y_{1i} - Y_{0i} \middle| T_i = 1, P(X_i)\right]\right\} \\
&= E\left\{E\left[Y_{1i} \middle| T_i = 1, P(X_i)\right] - E\left[Y_{0i} \middle| T_i = 0, P(X_i)\right] \middle| T_i = 1\right\}
\end{aligned} \tag{1}
$$

which is interpreted as the expected gain in log earnings for public sector workers (treatment group) compared with what they would have earned had they been in the private sector (non-treatment group). Conceptually the ATT requires a mean for the unobservable counterfactual, $E\left[Y_{0i} \middle| T_i = 1\right]$ so for the observable quantities in equation (1) to identify the ATT relies on three key conditions introduced into the literature by Rosenbaum and Rubin (1983).

The first condition is that of "unconfoundedness" $\left(Y_{0i}, Y_{1i} \perp T_i \middle| X_i\right)$ where $\perp$ denotes independence. According to this, potential outcomes are independent of treatment, conditional on the observable covariates, X_i. For example, in the case of the public sector pay premium, unconfoundedness implies that the odds of working in the public sector are unrelated to what earnings would be if the worker was not in the public sector, conditional on the observable covariates, X_i.

Unconfoundedness is often expressed as a conditional independence assumption; given observable covariates, assignment to the treatment group is random. Hence, systematic differences in actual outcomes between treated and non-treated individuals with the same value of the covariates can be attributed to the treatment. The second condition is that of "common support" or overlap $\left(0 < prob(T_i = 1 \middle| X_i) < 1\right)$, so that for each X_i a sufficiently large sample will yield observations for which we observe a Y_{0i} and other observations for which we observe a Y_{1i}. In other words, all treated individuals have a counterpart in the non-treatment group for each X_i for which we seek to make a comparison.[3]

The overlap condition seems to create dimensionality problems when many covariates are matched on; for example, if X_i contains *k* covariates which are all dichotomous the number of possible matches will be 2^k. But the propensity score reduces the dimensionality of the matching problem because it is possible to match on $P(X_i)$ which is scalar, rather than on the vector of observable variables X_i. Formally, the unconfoundedness and overlap conditions can be re-expressed as $\left(Y_{0i}, Y_{1i} \perp T_i \middle| P(X_i)\right)$ and $\left(0 < prob(T_i = 1 \middle| P(X_i)) < 1\right)$, so long as the "balancing" property $\left(prob\left(X_i \middle| T_i = 1, P(X_i) = p\right) = prob\left(X_i \middle| T_i = 0, P(X_i) = p\right)\right)$ holds (Rosenbaum and Rubin, 1983). In other words, conditional on the propensity score, the means of the covariates should be identical across the treatment and control groups if the balancing property holds.

Since the propensity score is a continuous variable it is unlikely that there are two observations with exactly the same value of $P(X_i)$, so further refinement is needed for equation (1) to be estimated. Therefore a "kernel matching" procedure is used, following Dehejia and Wahba (2002). With kernel matching, all treated individuals are matched with a weighted average of non-treated individuals, where the weights are inversely proportional to the distance

[3] If there are regions where the support of X_i does not overlap for the treated and non-treated groups, matching can only be performed over the common support region.

between propensity scores of the treated and non-treated. In addition, to show the robustness of the results a nearest-neighbor matching procedure is also used.

In summary, to implement the PSM estimates of the public sector pay premium, probit equations were estimated separately for each country in the sample. In each case the dependent variable was an indicator variable for whether the worker was employed in the public sector and the characteristics of the worker, job attributes and union membership are used as covariates (the full list is in Table 1). After the matching a "balancing" test is carried out. The treatment effects are then estimated over the region of common support using the method of kernel matching, with nearest-neighbor matching also used to provide some corroboration. Since the ISSP samples are small, males and females are pooled together and a gender dummy variable is included in the matching covariates.[4]

4. Results

Table 2 contains the results of estimating equation (1) for each country, using three sets of covariates; just worker characteristics, then adding job attributes, and finally adding union membership.[5] In the first block of the table, where just worker characteristics are controlled for, the ATT of the expected gain in log earnings for a public sector worker compared to what they would have earned in the private sector ranges from 0.119 in Britain to 0.269 in Canada. The ATT estimates are statistically significant except in Britain, according to standard errors from 500 bootstrap replications. Thus, matching on age, education, gender, marital status and location, the premium for working in the public sector is 30.9% in Canada, 25.2% in Australia and 16.6% in the United States. If nearest-neighbor matching is used the patterns are similar, although with slightly higher point estimates for Canada and the United States.

Expanding the matching so that job attributes are also included causes no change in the estimated premium in Canada, a slight fall in Australia and a higher but still statistically insignificant premium in Britain. Only in the United States does matching on job attributes cause the public sector pay premium to fall significantly.[6] That job attributes have so little impact does not reflect some lack of relevance since the attributes are jointly significant at the $p<0.001$ level in the probits used for the propensity scores. Rather it is that many attributes of public sector jobs are in some sense `better' (e.g. being more secure, less of a dead-end, less physically demanding, etc) so matching on these does not reveal any compensating differentials in public sector pay. Indeed, in contrast to Bender and Elliot (2002) the addition of job attributes makes the unexplained public sector pay premium in Great Britain appear even larger (although surrounded by wide standard errors).

While job attributes do not have much impact on the estimated public sector pay premium, union membership does. The average treatment effects in the final block of Table 2 are lower for

[4] Gregory and Borland (1999) also use pooled male and female sub-samples of ISSP to estimate earnings equations.

[5] Probits for the propensity scores and tests of the balancing property (holding in all cases) are available on request. Also available are the results of an indirect test of the unconfoundedness assumption, where the ATT is estimated for an outcome (family size) that is not expected to be affected by working in the public service, and this outcome is shown to be independent between the PSM treatment and control groups that are formed conditional on the covariates.

[6] There are also fewer control group observations in the common support, with up to 30% not used in the matching. This suggests that some private sector jobs are quite dissimilar to public sector jobs.

all four countries and for both matching estimators, compared with the middle block where the same covariates are used except for union membership. For example, the pay premium for public sector workers in Canada falls from 30.1% when the matching is on worker characteristics and job attributes to just 8.4% when the matching also includes current union membership. It is only in Australia that there remains a statistically significant public sector pay premium once differences in union membership across the public and private sector are controlled for.

5. Conclusions

When worker characteristics and job attributes are controlled for, the public sector pay premium ranges from 30% in Canada to 19-20% in Australia and Great Britain and only 6% in the United States. Ignoring differences in job attributes between private sector and public sector workers makes almost no difference to the estimated pay premium. That job attributes have so little impact suggests that higher pay in the public sector is not because of any compensating differentials.

However the estimated public sector pay premium is reduced in all countries, and disappears in Canada, once differences in union membership across sectors are controlled for. This finding favors the hypothesis that the pay premium partially reflects rents accruing to public sector workers, obtained most probably with assistance from the actions of their labor unions.

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Table 1: Average Characteristics of the Public Sector and Private Sector Workers, Job Attributes and Union Status in the Four Country Samples

	Australia		Canada		Great Britain		United States	
	Private	Public	Private	Public	Private	Public	Private	Public
Log annual earnings	10.5347	10.7743**	10.5472	10.8284**	9.7361	9.8156	10.1868	10.4577**
	(.7147)	(.5221)	(.7311)	(.5098)	(.7729)	(.6942)	(.9725)	(.6600)
Years of schooling	14.5078	15.0673*	14.4972	15.2797**	12.8357	13.6441*	13.3295	14.9873**
	(5.1802)	(3.6042)	(3.4122)	(3.5659)	(5.5552)	(3.2358)	(2.899)	(3.0053)
Age	42.2521	44.5625**	44.1583	45.5932	41.3811	42.7288	42.3067	42.8797
	(12.7366)	(10.3563)	(11.6877)	(10.6941)	(12.1293)	(10.573)	(12.2441)	(11.3302)
Male	.5308	.3846**	.5417	.4322**	.5140	.3559**	.5595	.4051**
	(.4994)	(.4877)	(.499)	(.4975)	(.5007)	(.4808)	(.4968)	(.4925)
Married	.6731	.7115	.6861	.6610	.6748	.6864	.5057	.5949**
	(.4694)	(.4541)	(.4647)	(.4754)	(.4693)	(.4659)	(.5004)	(.4925)
Job attributes: (5-point scale where 1=agree or always and 5 = disagree or never)								
Job is secure	2.4487	2.0577**	2.4028	2.1610**	2.3706	2.1186**	2.2577	1.9241**
	(1.0661)	(1.0151)	(1.1278)	(1.0699)	(.9925)	(.9352)	(1.0628)	(.981)
Job is interesting	2.2618	2.0673**	2.0028	1.8983	2.2692	1.9237**	2.0294	1.7405**
	(.9196)	(.8599)	(.8972)	(.7995)	(.9519)	(.8076)	(.9178)	(.9455)
I can work independently	2.0265	2.1731**	1.7583	2.0508**	2.0594	2.1102	2.0359	2.1709*
	(.7844)	(.8563)	(.7899)	(.8458)	(.8584)	(.9677)	(1.022)	(1.0356)
My job can help other people	2.2111	1.8846**	1.9306	1.7458**	2.2552	1.5932**	1.876	1.5063**
	(.9273)	(.7716)	(.8945)	(.7418)	(.9335)	(.6437)	(.8427)	(.6934)
My job is useful to society	2.3353	1.7788**	2.1361	1.6949**	2.4720	1.678**	2.0408	1.4873**
	(.9908)	(.735)	(.9387)	(.7791)	(1.0247)	(.8047)	(.9463)	(.7115)
Job helps improve my skills	2.2400	1.9856**	2.1000	2.0339	2.3671	2.1017**	2.0277	1.7911**
	(.9333)	(.8596)	(.9025)	(.8666)	(.907)	(.9991)	(.9398)	(.9447)
Job has hard physical work	3.4885	3.9327**	3.6056	3.7966	3.5559	3.6695	3.3328	3.9494**
	(1.2233)	(1.0473)	(1.2173)	(1.1213)	(1.2434)	(1.1698)	(1.4108)	(1.1554)
Job is stressful	2.7358	2.7452	2.6917	2.5593	2.7762	2.6017*	2.7259	2.557*
	(.88)	(.8209)	(.957)	(.9387)	(.8072)	(.8784)	(1.1143)	(.9936)
Job is dangerous	4.0410	4.0048	3.9944	3.8305	4.1958	4.0254	3.9347	3.8861
	(1.1167)	(1.1229)	(1.1121)	(1.157)	(.9859)	(1.1433)	(1.2455)	(1.3161)
Member of a union in my job	0.1894	0.5673**	0.1833	0.7034**	0.1678	0.6272**	0.0897	0.3481**
	(0.3921)	(0.4966)	(0.3875)	(0.4587)	(0.3743)	(0.4856)	(0.2860)	(.04779)
Sample size	829	208	360	118	286	118	613	158

Notes: Source is the 2005 ISSP Work Orientations survey. Samples in each country are those respondents that were working for pay at time of the survey. Standard deviations in parentheses.

** (*) = mean for public and private sector workers is significantly different at 5% (10%) level within that country.

Table 2: Propensity Score Matching Estimates of the Public Sector Pay Premium in Four Countries

Country	Average treatment effect on treated - kernel (ATT*k*)	Bootstrap standard error for ATT*k*	*t* statistic on ATT*k*	Percent pay premium	% of control group sample in common support	ATT (nearest neighbor) ATT*n*	Stand-ard error ATT*n*
	Only worker characteristics used to form propensity scores						
Australia	0.225	(0.046)	4.87	25.2%	92.7	0.229	(0.044)
Canada	0.269	(0.064)	4.17	30.9%	99.2	0.306	(0.063)
Great Britain	0.119	(0.086)	1.39	12.6%	96.9	0.101	(0.081)
United States	0.154	(0.069)	2.24	16.6%	99.5	0.194	(0.105)
	Worker characteristics + job attributes used to form propensity scores						
Australia	0.186	(0.050)	3.71	20.4%	86.4	0.212	(0.079)
Canada	0.263	(0.077)	3.43	30.1%	75.6	0.257	(0.115)
Great Britain	0.173	(0.115)	1.50	18.9%	70.6	0.207	(0.147)
United States	0.061	(0.080)	0.76	6.3%	90.2	0.140	(0.119)
	Worker characteristics + job attributes + union membership used to form propensity scores						
Australia	0.172	(0.055)	3.15	18.8%	89.0	0.171	(0.079)
Canada	0.081	(0.073)	1.11	8.4%	78.3	0.096	(0.130)
Great Britain	0.056	(0.138)	0.41	5.8%	62.6	0.118	(0.184)
United States	0.001	(0.079)	0.00	0.1%	85.8	-0.025	(0.171)

Note: The average treatment effect on the treated (ATT) is based on log annual income of ISSP-WO respondents working for pay at the time of the survey and only using observations that satisfy the common support assumption, with either kernel matching (ATT*k*) or nearest neighbor (ATT*n*) matching used. The percent pay premium is estimated as 100×[exp(ATT)-1]. The standard errors are obtained via 500 bootstrap replications. The covariates included in the sets of worker characteristics and job attributes are listed in Table 1 and also include dummy variables for region of residence within each country.