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PUBLIC-PRIVATE SECTOR PAY DIFFERENTIALS IN A DEVOLVED SCOTLAND

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The public-private sector wage gap in Scotland in 2000 is analysed using the extension sample of the British Household Panel Study (BHPS). Employing a switching regression model, and testing for double sample selection from the participation decision and sector choice, the wage gap is shown to be 10 % for males and 24 % for females. For males this is mainly due to differences in productive characteristics and selectivity, while for females the picture is more ambiguous. Findings also suggest that there exists a male private sector wage premium. While there is no evidence of a sample selection bias for females, the sector choice of males is systematically correlated with unobservables. Furthermore, the structural switching regression indicates that expected wage differentials between sectors are an important driving force for sectoral assignment.

JEL-Classification: J71, J31, C24

Keywords: wage differentials, switching model, double sample selection, decomposition

I. Introduction

Devolution has brought partial political and economic independence to some regions in the United Kingdom. Prominently, Scotland is among those regions and has elected its own parliament in 1999. However, devolution has had surprisingly little impact on public sector pay setting arrangements across the country while it may have increased the relative size of the sector in devolved regions. The aim of this paper is therefore to establish whether a Scottish public sector earnings premium exists at the outset of the devolution and how, if at all, results differ compared to

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other UK studies. Differences between the market rate and earnings in the public sector will also raise the question whether it is in the interest of Scotland to remove any wage premium if granted the necessary autonomy. This is of particular importance given that Scotland finds it ever more challenging to compete for skilled employees in a buoyant British labour market and circumvent out-migration to other parts of the UK.

In general, understanding potential wage differentials between the public and private sector presumes an understanding of the pay determination. The economic literature on public-private sector wage differentials offers various theoretical explications for the existence of wage premiums both in the public and private sector, reviewed comprehensively in Bender (1998) and Gregory and Borland (1999).

The fundamental and most widely used explanation has been that wage determination in the public sector is subject to an ultimate political constraint whereas the private sector is characterised by a profit constraint. For example, public sector employees do not only produce goods and services but also engage in vote-producing activities, which may justify higher pay (Gunderson 1979).

Furthermore, trade unions may exhibit more freedom to bargain as public sector services are essential and labour demand is, therefore, rather inelastic. Unsurprisingly, union membership density across many developed countries is much higher in the public sector compared with the private sector (Gregory and Borland 1999). This is also the case, for instance, in the UK, where unions in the private sector lost ground in the 1980s but remained relatively strong in the public sector.

On the other hand, it is not clear a priori why public sector employees should enjoy higher wages despite the above explanations. As Gregory (1990) argues, employees in the public sector may enjoy fringe benefits such as longer holidays, greater job satisfaction or superior pension schemes compared with private sector employees. Hence, wages for similar employees in comparable jobs should be lower in the public sector. Since these fringe benefits are rarely observed in empirical studies they may lead to an observed private sector wage premium which in fact is just a compensation for the lack of fringe benefits.

The validity of theories, however, depends very much on the economic, institutional and political environment. Elliott et al. (1999) list several possible dimensions through which wage setting in both public and private sector may be affected, such as changes in the product market environment, pressure to contain production costs, new production technologies, changes in the role of unions and political pressure to decrease public spending.

In the United Kingdom, wage setting has been characterised by the principle of comparability between public and private sector pay for the last 100 years or so. A paramount aim of governments has been to guarantee equal pay across sectors. This commitment played a particularly crucial role in the late 1940s when the public sector was expanded significantly due to the nationalisation of former private industries, as Bender (2003) reports. In the second half of the 1980s, however, many of these nationalised industries were re-privatised and the role of trade unions diminished. At the same time the principle of comparability of wages was replaced by the comparability of the growth rates of the average wage. Pay Review Boards, first introduced in the early 1970s to review wages and pass recommendation to the government, were extended to cover further occupations. Wage bargaining was further decentralised in the early 1990s.

However, devolution has not intensified this trend. While some public sector occupations in Scotland are negotiated at a Scottish rather than a UK-wide level (e.g. central government, local authorities, teachers, and prison officers) for still a large number regional flexibility in pay setting is limited (e.g. police and fire-services, universities, and UK government departments with a presence in Scotland). Additionally, hybrid systems are operated in areas such as the NHS.

The principle of pay comparability should assure that the public-private sector wage differential in the UK is rather small. However, the ambiguous existence of pay differentials found in several empirical studies may indicate a lack of enforcement of these policies. Studies for the UK on sectoral wage differentials are scarce.

A number of empirical studies based on micro-level data have been published in the last two decades or so drawing on mainly four different data sources: the New Earnings Survey (NES) (Elliott and Murphy 1987, Gregory 1990, and Elliott and Duffus 1996), the General Household Survey (GHS) (Rees and Shah 1995 and Disney and Gosling 1998), the British Household Panel Study (BHPS) (Bender and Elliott 1999) and more recently the British SCELI survey (Bender 2003). These studies vary significantly in methodology, scope, and findings.

At the same time there is a noticeable lack of studies on earnings differences across UK regions. In the only UK-wide study Henley and Thomas (2001) show that private and public sector employment are weakly positively correlated across British regions using BHPS data. However, they also find evidence of significant regional differences in public-private sector pay gaps. In a second study which has evolved in parallel to this paper Elliott et al. (2004) analyse pre- and post-devolution earnings differentials for Scotland using data from the Labour Force

Survey (LFS). They find a decline in the average public sector premium in the post-devolution period but also show that the pay gap varies along the earnings distribution. The authors argue that bringing public sector pay in line with private sector wages may therefore increase efficiency but may also increase earnings inequality.

In this paper, new data from the BHPS is used to analyse wage differentials in the public and private sector in Scotland for the post-devolution period. In contrast to other UK-wide and regional studies the paper accounts explicitly for possible sample selection bias from both sector choice and labour participation and explores its impact on the pay gap.[1]

The paper is structured as follows. The next section will describe the data set and provide some descriptive statistics. Section III outlines the econometric framework and discusses identification issues. Finally, Section IV reports the results on wage equations and decomposition and the final Section discusses policy implications.

II. Data

The data is drawn from the BHPS. Since its introduction in 1991, over 5,000 households made up of roughly 10,000 individuals have been interviewed annually. While it has always been a nationally representative sample, only since 1999 extension samples for Scotland and Wales have been launched, aiming to increase the relatively small sample size - approximately 500 households in each country - to 1,500 households. The main objective has been to enable independent analysis of the two countries on a representative level.

The sample of employees contains only individuals who are, at the date of the interview, full-time employees, aged 16 to 64 (16 to 59 for women), not self-employed and not working in either agriculture, non-profit organisations or for the armed forces. Since the econometric framework requires identification variables which are only available for some waves, the paper makes use of data from wave 10 in 2000.

The resulting overall cross-section sample consists of 1,054 males and 1,230 females of which 61% of males and 42% of females are participating in the labour

[1] Even though Bender (2003) controls for sample selection from the sector choice, the selection term is not separately stated in the decomposition and it is also unclear whether that component is part of the explained or unexplained part which might have a significant impact on the results (Gyourko and Tracy 1988).

market. Of these, roughly 22% of men and 41% of women are employed in the public sector. Appropriate cross-section weights have been applied to enable separate analysis for Scotland only.

The total number of females employed in the public sector follows a u-shape, the number for men fell since 1996, peaked in 1999 before dropping in the following year and has risen again since. As one would expect, the latest increase in male public sector employment is mainly driven by increases in civil service and local government jobs in the aftermath of devolution and the instalment of the Scottish Parliament. This development is also present among females although it is less pronounced. Unsurprisingly, the percentage of women working in health care and higher education is far higher compared to men. On the other hand, men dominate jobs in the central governments as well as in the civil service. However, the largest share of public sector employment for both men and women stems from local governments.

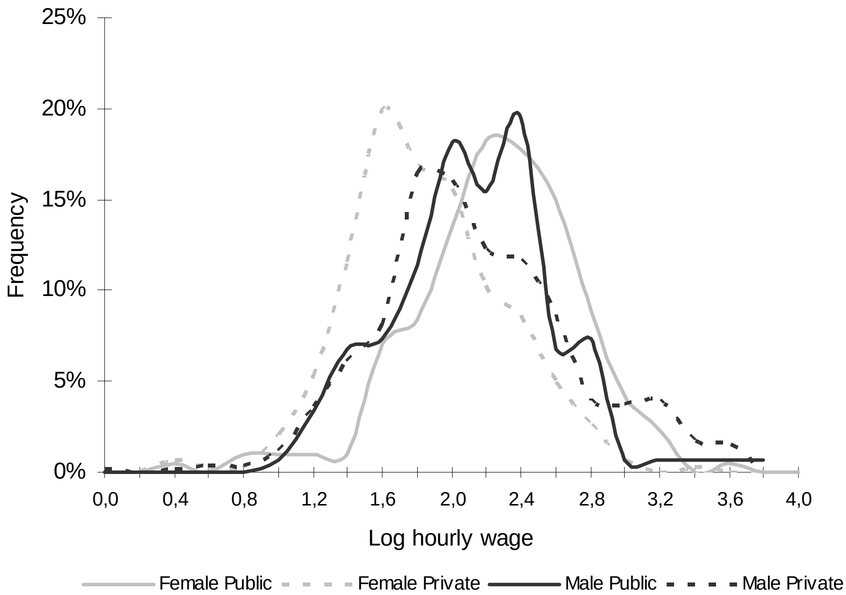
Furthermore, public sector workers are selected differently into occupations. For both men and women, public employees are more likely to be found in managerial occupations compared to their private counterparts. This is especially pronounced for females where 63% are employed in this occupation compared to only 31% in the private sector. Additionally, public workers are more likely to be in professional rather than in unskilled occupations. However, whether the above patterns are supply or demand driven is a priori not clear.

Figure 1 depicts the unconditional log wage distribution in the public and private sector by gender. Following the usual convention in the literature (e.g. Booth at al. 2003) wages are expressed in log hourly gross pay.[2] As Figure 1 shows, the male pay distribution in the public sector is less spread compared to the private sector distribution though there is a significant overlap. Furthermore, the mean wage in the former is higher than in the latter. The respective mean log wages in the public and private sector are 2.2 and 2.0. A simple "difference in means" test reveals that the pay in the public sector is indeed significantly higher than in the private sector even on the 1% significance level. In contrast, the overlap in the female public and private earnings distribution is much smaller compared to males. Public pay is systematically higher in the upper end and much flatter at the lower end of the distribution compared to the private one. Once again, the difference in means between public (2.15) and private (1.74) sector is positive and significant

[2] The log hourly gross wage rate is defined as $w = \ln[\ Paygu\ /\ ((30/7)\ (Hs + \alpha \mathrm{Hot}))\ ]$, where *Paygu* is the monthly gross pay in the current job, *Hs* is standard weekly hours, *Hot* is paid overtime hours per week and α is the overtime premium set to 1.5.

at the 1% level. Hence, on average, there is a substantial and significant pay gap for both males and females. A Wilcoxon rank-sum test for the equality of the public and private distribution is clearly rejected for both males and females.

Figure 1. Log hourly wages in the public and private sector by gender in 2000.



III. Econometric framework

The paramount aim of the paper is to estimate wages in the public and private sector in order to study the causes of the unconditional earning differentials and derive a *structural* switching equation that determines the sector choice. Hence, in the following a switching regression model is adopted (Lee 1978) that estimates separate wage equations for the public and private sectors. Let $w_{1,i}$ and $w_{2,i}$ be the hourly wages in the public and private sector, respectively. Thus, the two log wage equations to be estimated are:

$$\ln w_{1,i} = X_{1,i}'\beta_1 + \varepsilon_{1,i} \tag{1}$$

$$\ln w_{2,i} = X_{2,i}{}' \beta_2 + \varepsilon_{2,i} \tag{2}$$

where X_i is a matrix of explanatory variables, β the vector of corresponding coefficients to be estimated and ε the error term. Henceforth, the index $j = 1,2$ refers to public and private sector, respectively.

In general, simple OLS of equations (1) and (2) may lead to inconsistent estimates. First, OLS estimates are prone to suffer from sample selection bias due to the exclusion of non-participants in the labour force. If the participation decision is systematic the pool of employees is non-random. This problem is commonly addressed by including an additional regressor which corrects for the bias in the participation decision (Heckman 1979). Second, given the participation decision, individuals have to decide in which sector to work. Again, if the assignment to public or private employment is non-random, OLS estimates are biased and a further correction term for this type of self-selection is required (Maddala 1983, Maddala and Nelson 1975). The recent literature on public-private sector earning differentials has mainly accounted for the latter and widely ignored the former.[3] However, controlling for one type of selection in earnings equations only and ignoring non-labour force participants may still lead to biased estimates.

Hence, this section will closely follow an approach by Co et al. (1999) used to study self-employment decisions, adjusting for multiple selection types. In order to test, and potentially account for, both types of selection, a double sample selection model can be fitted (Tunali 1986). Let the reduced form participation and sector choice equations be determined by

$$P_i^* = Z_i' \gamma + u_i \tag{3}$$

$$S_i^* = B_i' \mu + v_i \tag{4}$$

where P^* and S^* are latent variables, Z and B the vectors of characteristics, γ and μ the coefficients to be estimated and μ_i and v_i the error terms for participation and sector respectively.

An individual will participate in the labour market if the utility of participation exceeds the gain of non-participation. Similarly, individuals will choose the public sector if the expected earnings differential is positive and/or personal preferences for public sector employment are strong. These preferences may be correlated

[3] The only exception the author is aware of is Stillman (2000) in a study on the Russian labour market.

with individual characteristics and captured in B.[4] Note that the expected earnings difference is not observable prior to the estimation of the wage equations (1) and (2). Hence, we will first estimate the *reduced* sector choice equation to gain consistent wage estimates and only then the *structural* switching regression.

Since neither latent variable is observable, two index functions are defined. In the case of participation this is $P_i = 1$ *if* $P^* > 0$ and $P_i = 0$ *if* $P^* \leq 0$, where $P_i = 1$ and $P_i = 0$ indicate labour market participation and non-participation respectively. Similarly, for the reduced sector choice equation $S_i = 1$ *if* $S^* > 0$ and $S_i = 0$ *if* $S^* \leq 0$, where $S_i = 1$ and $S_i = 0$ indicate public or private sector employment, respectively. Clearly, $S_i = 1$ and $S_i = 0$ are only observed for $P_i = 1$.

Given the above structure, consistent estimates can be achieved by Maximum Likelihood Estimation (MLE) (Co et al. 1999). Yet, the number of parameters to be estimated is rather large. Alternatively, a simple two-step Heckman (1979) approach with extended correction terms may be adopted (see, e.g., Lee 1979, Ham 1982, Fishe et al. 1981 and Tunali 1986). In the first, step equations (3) and (4) are estimated and sample selection correction terms are constructed. In the second step, equations (1) and (2) are estimated via simple OLS including the correction terms as additional regressors.

Two cases can be distinguished, $\rho_{uv} = 0$ and $\rho_{uv} \neq 0$ where ρ is the error correlation term between equation (3) and (4) and the former is a special case of the latter.[5] For $\rho_{uv} \neq 0$ the approach is to estimate (3) and (4) using a bivariate probit (Ham 1982, Tunali 1986). In that case (ε_j, u, v) are jointly normally distributed with mean zero and covariance matrix

$$\sum{}_j = \begin{bmatrix} \sigma^2_{\varepsilon j} & \sigma_{\varepsilon jv} & \sigma_{\varepsilon ju} \\ & \sigma^2_v & \sigma_{vu} \\ & & \sigma^2_u \end{bmatrix}$$

where σ^2_v and σ^2_u are normalised to unity for identification purposes following

[4] For a more rigorous theoretical derivation see, e.g., van der Gaag and Vijverberg (1988).

[5] In the extreme case where the participation and sector choice depend solely on the difference between reservation wage and either public or private sector wage the decision to participate and the sector choice are indeed simultaneous. Yet, the two decisions are likely to depend on the expected utility which is impacted on by more than simply wage differences. The bivariate probit is a way to check whether the decisions are correlated or not.

standard practice.[6] Maximum likelihood of the bivariate probit leads to four sample selection correction terms

$$\hat{\lambda}_{i,p1} = \phi(Z_i^{'}\hat{\gamma})\Phi\left[\frac{B_i^{'}\hat{\mu} - \rho Z_i^{'}\hat{\gamma}}{(1-\rho^2)^{1/2}}\right] \times F(B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, \rho)^{-1} \tag{5}$$

$$\hat{\lambda}_{i,s1} = \phi(B_i^{'}\hat{\mu})\Phi\left[\frac{Z_i^{'}\hat{\gamma} - \rho B_i^{'}\hat{\mu}}{(1-\rho^2)^{1/2}}\right] \times F(B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, \rho)^{-1} \tag{6}$$

$$\hat{\lambda}_{i,p2} = \phi(Z_i^{'}\hat{\gamma})\Phi\left[-\frac{B_i^{'}\hat{\mu} - \rho Z_i^{'}\hat{\gamma}}{(1-\rho^2)^{1/2}}\right] \times F(-B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, -\rho)^{-1} \tag{7}$$

$$\hat{\lambda}_{i,s2} = -\phi(B_i^{'}\hat{\mu})\Phi\left[\frac{Z_i^{'}\hat{\gamma} - \rho B_i^{'}\hat{\mu}}{(1-\rho^2)^{1/2}}\right] \times F(-B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, -\rho)^{-1} \tag{8}$$

where $\rho = \rho_{uv}$. Again, ϕ and Φ are the univariate standard normal density and distribution functions respectively, and F the bivariate standard normal distribution function. Equations (5) to (8) collapse to the usual univariate Heckman expressions for $\rho_{uv} = 0$.

Now, the wage equations (1) and (2) can be re-written as

$$E(\ln w_{1,i} \mid X_i^{'}, P_i = 1, S_i = 1) = X_{1,i}^{'}\beta_1 + \sigma_{11}\rho_{1u}\hat{\lambda}_{i,p1} + \sigma_{11}\rho_{1v}\hat{\lambda}_{i,s1} \tag{9}$$

$$E(\ln w_{2,i} \mid X_i^{'}, P_i = 1, S_i = 0) = X_{2,i}^{'}\beta_2 + \sigma_{22}\rho_{2u}\hat{\lambda}_{i,p2} + \sigma_{22}\rho_{2v}\hat{\lambda}_{i,s2} \tag{10}$$

where the correction terms are according to (5) to (8).

[6] Since the covariance of ε_1 and ε_2 is not identifiable in this model, the covariance matrix has been split into two matrices (see Co et al. 1999 for details; see Koop and Poirier 1997 for a discussion on these identification issues).

IV. Decomposition

Once wages are consistently estimated, differences in public and private sector pay can be decomposed into several components. In the following a modified decomposition methodology is applied suggested by Neuman and Oaxaca (2002). According to this the wage gap is split into three terms such that

$$\ln \overline{w}_1 / \overline{w}_2 = (\overline{X}_1 - \overline{X}_2)\hat{\beta}_1 + \overline{X}_2(\hat{\beta}_1 - \hat{\beta}_2) + \tag{11}$$

$$[\hat{\sigma}_{11}(\hat{\rho}_{1u}\hat{\lambda}_{i,p1} + \hat{\rho}_{1v}\hat{\lambda}_{i,s1}) - \hat{\sigma}_{22}(\hat{\rho}_{2u}\hat{\lambda}_{i,p2} + \hat{\rho}_{2v}\hat{\lambda}_{i,s2})],$$

where $\ln \overline{w}$ is the predicted mean log wage, $\overline{X}$ the mean vector of characteristics, $\hat{\beta}$ the estimated vector of coefficients, and $\hat{\lambda}$ the estimated mean correction term. Yet, $\hat{\lambda}$ is a non-linear function in $Z_i^{'}\hat{\gamma}$ and $B_i^{'}\hat{\mu}$ and the central tendency is estimated as $\hat{\lambda} = \sum_{i=1}^{N_j}\hat{\lambda}_i / N_j$, where $\hat{\lambda}_i$ is the estimated correction term from the first step in equation (3) and (4) and N_j refers to the respective set of observations in each sector (Even and Macpherson 1990).

Similar to the simple Oaxaca decomposition (Oaxaca 1973) the term $(\overline{X}_1 - \overline{X}_2)\hat{\beta}_1$ is the explained and $\overline{X}_2^{'}(\hat{\beta}_1 - \hat{\beta}_2)$ the unexplained part of the predicted mean wage gap. However, it is a priori unclear how to tread the selection terms in equation (11). One way of dealing with them is by subtracting the terms from the left hand side which leaves one with the familiar Oaxaca decomposition where the left hand side is now the selectivity corrected wage differential as opposed to the observed differential (e.g. Reimers 1983).[7]

V. Estimation results

A. Identification and variable choice

Estimating the above two step model requires some identification assumptions on the coefficients and the covariance parameters (Tunali 1986). First, as already stated, σ_v^2 and σ_u^2 are normalised to 1. Depending on whether the selection

[7] Equation (11) can be decomposed differently by using the private sector wage structure $\hat{\beta}_2$ as weight rather than the public wage structure $\hat{\beta}_1$. Since results may vary, both methods are reported.

equations are independent, further identification assumptions are necessary. In case $\rho_{uv} = 0$, the matrices $Z_i^{'}$ and $B_i^{'}$ are required to contain at least one element that is not part of $X_{i,j}^{'}$ (j=1,2). In the second case, where $\rho_{uv} \neq 0$, an additional identification is necessary in order to estimate σ_{vu}. Hence, at least one element in $Z_i^{'}$ must not be contained in $B_i^{'}$ and vice versa. Additionally, these variables must not be part of $X_{i,j}^{'}$.

In the case of the participation equation, identification has often been achieved by controlling for the number of children. Hence, data on the number of children in two age groups, 0-11 and 12-18 are included in the participation equation and not in either the sector choice or wage equations.

More difficult, however, is the identification of the sector choice equation (4). It has been argued that social background characteristics do not impact on the wage but on the sector decision. Various variables such as father's and mother's education and occupation and the number of siblings have been used for identification (see, e.g., Bender 2003 and Hartog and Oosterbeek 1993).

Even though the BHPS contains questions on, for example, the occupation of the respondent's father, the number of observations is very small and does not allow the use of these variables. Yet, there is evidence to suggest that public sector employees are far more likely to be unionised compared to private sector workers (see e.g. Gregory and Borland 1999). Hence, one alternative identification measure is union status. On the other hand, union status is usually controlled for in the wage equation as well and using it in the sector choice equation as an identification variable would render this unfeasible. The BHPS, however, asks individuals about their perception and the importance of unions.[8] Since union status and union perception are positively correlated but well below unity, which makes it possible to treat them as two distinct variables, union perception has been used as identification in the sector choice equation. Unfortunately, this question is only included in wave 10 and not in the latest wave.

Besides the identification variables, $Z_i^{'}$ and $B_i^{'}$ contain information on personal characteristics such as age, marital status and education. In addition, the sector choice equation also controls for occupation, firm size and job tenure. The wage equations contain the same set of regressors except for the identification variables. Additionally, appropriate sample correction terms are included depending on the model estimated.

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[8] In particular, individuals are asked whether strong trade unions are needed to protect the working conditions and wages of employees. Four possible answers can be given which range from strong agreement to disagreement.

Several features are worth mentioning in the sample.[9] First, as already shown, women are far more likely to be employed in the public sector. Women in the public sector are also significantly less likely to have children aged 0-11 but slightly more likely to have older children compared to men. This also holds for the private sector.

Second, union coverage is significantly higher in the public sector for both men and women. This also carries over into the perception of unions. Employees of both sexes in the public sector perceive unions as an important institution.

Furthermore, public sector workers are more likely to be married and exhibit far greater job attachment. Yet, there are only minor differences in terms of educational levels. On the other hand, as already mentioned, both men and women in the public sector are more likely to be found in managerial positions compared to their private sector counterparts.

B. Estimation results

Univariate and bivariate results for both males and females do not support a simultaneous estimation of the participation and sector choice decision.[10] The correlation coefficient ρ_{uv} is not significantly different from zero. Furthermore, the assumption of correlation does not change the estimated coefficients substantially, which is not uncommon in this literature (e.g., Fishe et al. 1981 and Tunali 1986).

For male employees the following patterns with respect to participation and sector choice arise. First, young people (aged 16-20) are significantly less likely to both select themselves into the public sector and participation compared to the base category (employees aged 31 and older). Many studies have ascribed the age effect to queuing for public sector jobs (see, e.g., Bender 1998 and van der Gaag and Vijverberg 1988).

Second, married men are significantly more likely to participate in the labour market compared to unmarried males. Yet, the marital status has no impact on the sector choice. Similarly, men with children (aged 0-11) are more likely to be found working compared to males without children. This holds also for men with older children, however, the effect is not statistically different from zero in the univariate probit case.

Third, the identification variables on trade union perception perform well.

[9] A summary table with sample characteristics can be obtained upon request.

[10] Full tables with univariate and bivariate results can be obtained from the author.

Employees who regard trade unions as important for the protection of working conditions and wages select themselves into public sector jobs compared to individuals who do not adhere to this perception.

Unsurprisingly, individuals employed in small or medium sized firms have a higher probability of working in the private sector.[11] Education and job tenure on the other hand do not impact significantly on either decision with some exceptions in the univariate probit case.[12] Finally, occupation matters for the sectoral decision. Individuals are significantly more likely to be employed in the public sector if they work in managerial or non-manual occupations.

The results for women are very similar, with some exceptions. First and somewhat surprisingly, females with children have a lower participation probability compared to women with very young children. Furthermore, obtaining a higher degree significantly increases the likelihood of labour market participation. Similarly, professional occupation increases the probability of public sector employment, as does job tenure.

In the second step, wage equations for public and private employment have been estimated using the probit results to construct selection correction terms. Tables 1 and 2 report the results for men and women respectively. Alongside OLS results, selection corrected wage equations are estimated for both the univariate and bivariate case. Standard errors for models 3 to 6 are based on a simple re-sampling bootstrap method (see Efron and Tibshirani 1993 for details) as the calculation of the corrected variance-covariance matrix is cumbersome. Thus, 1000 samples of size N are drawn from the original sample (parent sample) with replacement. For each sample all coefficients are re-estimated and then used to derive standard errors and confidence intervals.[13]

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[11] The majority of public sector workers working in small establishments is employed with local governments or work in town halls.

[12] Yet, once one does not control for occupation, education exhibits a significant impact on the sector decision. Hence, the main effect of education is on occupation and occupation then affects the sector choice.

[13] Three different types of intervals have been calculated, the normal (N), the percentile (P) and the bias correct (BC). If the bootstrap statistics are roughly normally distributed, the normal and percentile intervals will be fairly similar. However, if there are significant differences, percentile intervals are usually preferred. Furthermore, the point estimate of the original sample and the average statistic of the bootstrap do not necessarily agree and their difference is referred to as bias. Then, the bias corrected confidence interval takes these possible discrepancies into account. If the bias is small, percentile and bias corrected confidence intervals are roughly identical. Hence, all three intervals will be very similar for an approximately normally distributed bootstrap statistic and a small bias.

Table 1. Estimation results for males and three different model specifications

	OLS (no correction)		Univariate probit correction		Bivariate probit correction	
	(1)	(2)	(3)	(4)	(5)	(6)
	Public	Private	Public	Private	Public	Private
Constant	1.715	1.931	1.020	2.256	0.810	2.124
	(0.154)***	(0.077)***	(0.402)**	(0.180)***	(0.459)*	(0.137)***
Age 16-20	-0.347	-0.568	-0.493	-0.586	-0.617	-0.637
	(0.263)	(0.081)***	(0.407)	(0.164)*** (N)(BC)	(0.434)	(0.103)***
Age 21-30	-0.220	-0.151	-0.207	-0.186	-0.196	-0.173
	(0.104)**	(0.055)***	(0.132)	(0.065)***	(0.130)	(0.064)***
Married	0.162	0.161	0.193	0.124	0.256	0.159
	(0.072)**	(0.048)***	(0.141)	(0.071)	(0.134)**	(0.070)***
Member of union	0.125	0.060	0.255	0.080	0.246	0.076
	(0.070)*	(0.052)	(0.093)***	(0.054)	(0.086)***	(0.055)
Professional	0.741	0.769	0.747	0.810	0.704	0.784
	(0.156)***	(0.084)***	(0.198)***	(0.102)***	(0.190)***	(0.100)***
Managerial	0.643	0.763	0.774	0.811	0.780	0.810
	(0.107)***	(0.070)***	(0.134)***	(0.082)***	(0.140)***	(0.079)***
Skilled	0.305	0.183	0.433	0.242	0.489	0.269
non-manual	(0.090)***	(0.065)***	(0.133)***	(0.078)***	(0.143)***	(0.077)***
Skilled manual	0.052	0.199	-0.135	0.134	-0.106	0.147
	(0.102)	(0.061)***	(0.143)	(0.074)**	(0.138)	(0.071)**
Higher degree	-0.025	0.272	-0.046	0.225	0.003	0.256
	(0.222)	(0.098)***	(0.294)	(0.132)**	(0.290)	(0.121)**
A-level	-0.050	0.058	-0.020	0.045	0.007	0.063
	(0.082)	(0.048)	(0.109)	(0.062)	(0.105)	(0.059)
O-level	-0.100	-0.118	-0.156	-0.147	-0.165	-0.140
	(0.093)	(0.054)**	(0.124)	(0.067)***	(0.122)	(0.061)***
Job tenure	-0.006	0.001	0.013	0.015	0.007	0.011
	(0.014)	(0.008)	(0.019)	(0.010)	(0.019)	(0.010)
Job tenure sq	0.000	-0.000	-0.000	-0.001	-0.000	-0.001
	(0.001)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)*
Small	0.062	-0.341	-0.073	-0.431	-0.031	-0.404
	(0.104)	(0.049)***	(0.140)	(0.062)***	(0.130)	(0.058)***
Medium	0.047	-0.267	-0.010	-0.313	0.024	-0.291
	(0.093)	(0.054)***	(0.115)	(0.065)***	(0.109)	(0.060)***

Table 1. (Continued) Estimation results for males and three different model specifications

	OLS (no correction)		Univariate probit correction		Bivariate probit correction	
	(1)	(2)	(3)	(4)	(5)	(6)
	Public	Private	Public	Private	Public	Private
λ_p (Participation)			-0.164	-0.234	0.058	-0.079
			(0.416)	(0.257)	(0.822)	(0.236)
λ_s (Sector)			0.518	0.486	0.518	0.407
			(0.178)***	(0.162)***	(0.230)**	(0.163)***
ρ_{ju}			-0.283	-0.384	0.112	-0.160
			(0.537)	(0.360)	(0.787)	(0.380)
ρ_{jv}			0.892	0.799	1.012	0.826
			(0.186)***	(0.799)***	(0.288)**	(0.274)**
σ_{jj}			0.580	0.608	0.538	0.493
			(0.171)**	(0.124)**	(0.436)**	(0.982)**
					(P)(BC)	(P)(BC)
Observations	149	528	149	528	149	528
R-squared	0.46	0.54	0.52	0.56	0.51	0.56

Notes: OLS, sample selection terms based on separate probits, and sample correction terms based on bivariate probit. Dependent variable is wage in public and private sector. Cross-section weights applied. *,**, and *** denote significance at 10%, 5% and 1% respectively. Standard errors in parentheses, where OLS standard errors are robust and the S.E. for the remaining models are bootstrapped. (P) refers to percentile, (BC) to bias correct.

Clearly, significant coefficients are very similar across model specifications, while significance levels vary. This is particularly pronounced for females. In general, there is a tendency for standard errors in the OLS model to be slightly smaller than estimated using bootstrapping. However, in the majority of cases this only marginally affects significance levels.

Besides age, occupation has the most pronounced impact on wages in both the public and private sector. Unsurprisingly, professional, managerial and non-manual occupations yield significantly higher wages than unskilled occupations.

As expected, union membership significantly increases wages in the public sector yet unions do not affect the private sector wage. In contrast, firm size does impact on the latter, while it has hardly any affect on the former. Education has little

Table 2. Estimation results for females and three different model specifications

	OLS (no correction)		Univariate probit correction		Bivariate probit correction	
	(1)	(2)	(3)	(4)	(5)	(6)
	Public	Private	Public	Private	Public	Private
Constant	1.448	1.504 1	.188	1.567 1.	218	1.563
	(0.159)***	(0.098)***	(0.287)***	(0.198)***	(0.267)***	(0.216)***
Age 16-20	-0.500	-0.364	-0.628	-0.458	-0.560	-0.453
	(0.119)***	(0.098)***	(0.219)***	(0.134)***	(0.202)***	(0.122)***
Age 21-30	-0.201	0.005	-0.269	-0.052	-0.262	-0.047
	(0.099)**	(0.056)	(0.146)*	(0.087)	(0.139)*	(0.087)
Married	0.020	0.0900	0.008	0.064	0.003	0.057
	(0.051)	(0.053)*	(0.055)	(0.058)	(0.055)	(0.061)
Member	0.163	0.001	0.172	0.007	0.171	0.006
of union	(0.065)**	(0.050)	(0.066)***	(0.052)	(0.067)***	(0.052)
Professional	0.723	0.779	0.860	0.900	0.840	0.883
	(0.197)***	(0.129)***	(0.253)***	(0.181)***	(0.228)***	(0.174)***
Managerial	0.786	0.530	0.930	0.660	0.922	0.653
	(0.151)***	(0.081)***	(0.215)***	(0.135)***	(0.192)***	(0.130)***
Skilled	0.500	0.250	0.544	0.282	0.538	0.278
non-manual	(0.157)***	(0.066)***	(0.178)***	(0.078)***	(0.163)***	(0.075)***
Skilled manual	0.243	0.218	0.248	0.235	0.260	0.243
	(0.195)	(0.080)***	(0.218)	(0.093)***	(0.219)	(0.093)***
Higher degree	0.251	0.321	0.244	0.360	0.265	0.389
	(0.104)**	(0.106)***	(0.122)**	(0.144)***	(0.129)**	(0.147)***
A-level	-0.125	0.110	-0.151	0.090	-0.151	0.088
	(0.072)*	(0.059)*	(0.084)*	(0.069)	(0.083)**	(0.068)
					(N)	
O-level	-0.292	-0.005	-0.305	-0.024	-0.311	-0.031
	(0.097)***	(0.055)	(0.101)***	(0.062)	(0.105)***	(0.062)
Job tenure	0.002	0.007	0.010	0.015	0.007	0.013
	(0.012)	(0.012)	(0.015)	(0.013)	(0.014)	(0.013)
Job tenure sq	-0.000	-0.001	-0.000	-0.001	-0.000	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Small	0.031	-0.194	-0.031	-0.249	-0.024	-0.244
	(0.068)	(0.068)***	(0.094)	(0.090)***	(0.092)	(0.089)***
Medium	0.039	-0.049	-0.017	-0.107	-0.020	-0.113
	(0.077)	(0.070)	(0.097)	(0.091)	(0.100)	(0.095)
λ_p (Participation)			0.055	0.094	0.058	0.112
			(0.102)	(0.138)	(0.126)	(0.155)

Table 2. (Continued) Estimation results for females and three different model specifications

	OLS (no correction)		Univariate probit correction		Bivariate probit correction	
	(1) Public	(2) Private	(3) Public	(4) Private	(5) Public	(6) Private
λ_s (Sector)			0.192	0.249	0.180	0.227
			(0.188)	(0.202)	(0.167)	(0.191)
ρ_{ju}			0.143	0.223	0.495	0.317
			(0.253)	(0.308)	(0.409)	(0.397)
ρ_{jv}			0.499	0.594	0.158	0.643
			(0.406)	(0.396)	(0.306)	(0.564)
σ_{jj}			0.384	0.419	0.364	0.353
			(0.072)	(0.076)	(0.052)**	(0.060)**
Observations	214	302	214	302	214	302
R-squared	0.47	0.38	0.48	0.39	0.48	0.39

Notes: OLS, sample selection terms based on separate probits, and sample correction terms based on bivariate probit. Dependent variable is wage in public and private sector. Cross-section weights applied. *significant at 10%, **significant at 5 %, ***significant at 1%. Standard errors in parentheses, where OLS standard errors are robust and the s.e. for the remaining models are bootstrapped. (N) refers to normal.

effect on male public sector wages. The picture for females is more diverse, e.g., having a higher degree increases wages in the public sector; in contrast, A- and O-level holders perform worse in all three models in the public sector.

For males, sector selection has a significant impact on wages. Given conditional expected wages (equations (9) and (10)), employees working in the public or private sector perform better than a random individual would have done as the positive and significant coefficients on the correlation terms show. However, there is no indication for a participation bias in the univariate model and only weak evidence in the bivariate specification. For women neither selection coefficient is significant regardless of the model specification. This is surprising since the labour force participation for females in the sample is much lower compared to men.[14]

[14] The reported correlation terms and the standard deviations are constructed following Tunali (1986).

C. Wage gap and switching regression

Given the above results, predicted log wages in the public and private sector can be estimated consistently. Tables 3 to 5 report predicted log wages, differences in predicted log wages and the decomposition of predicted log wages into various parts according to equation (11) by gender and for three different model specifications. Furthermore, results are shown for two different weighting specifications with different base categories, $\beta^* = \beta_2 = \beta^{pri}$ and $\beta^* = \beta_1 = \beta^{pub}$, as results are usually different. Confidence intervals for significance tests have been bootstrapped and refer to 1300 replications.

Regardless of the specification, the unadjusted wage gap between public and private sector is statistically significant and around 10 % for males and 24 % for females, which lies well in line with Rees and Shah (1995) but is slightly higher than Bender's (2003) findings for the whole of the UK. For men in all three models and for either weighting scheme, the gap is more than accounted for by differences in characteristics. In contrast, differences in returns reduce the unadjusted gap, though the term is statistically insignificant. Hence, while returns to productive and job-related characteristics are lower for those in the public sector, differences in these characteristics of public sector employees more than counterbalance this effect leading to higher wages in the public sector.[15]

Results are more sensitive against the weighting scheme for women. Using the public wage structure the absolute figures are fairly similar to the male ones. However, if the private sector wage structure is used as weight, only between 46 and 67 % of the wage difference is explained by differences in characteristics suggesting a substantial public sector premium.

The model specifications with sample selection correction have an additional term besides the explained and unexplained component. In the presence of sample selection, unobserved productivity related characteristics will be captured in the unexplained component in the simple OLS specification. The existence (or non-existence) of wage premiums may, therefore, be simply due to these unobserved characteristics rather than discrimination.

As the results show, the selection term is positive for both males and females and statistically different from zero for the former. At the same time the unexplained

[15] Note that Rees and Shah (1995) find the same for males but not for females while Bender (2003) reports mainly positive contributions from both the explained and unexplained components. Yet, since both the data and methodology are very different, comparisons have to be drawn with caution.

Table 3. Pay gap between public and private sector: OLS results

	Males		Females	
	$\beta^{*}=\beta^{pri}$	$\beta^{*}=\beta^{pub}$	$\beta^{*}=\beta^{pri}$	$\beta^{*}=\beta^{pub}$
Predicted log wages	$\ln\overline{w}_{pub}=2.2380$	$\ln\overline{w}_{pri}=2.0239$	$\ln\overline{w}_{pub}=2.1496$	$\ln\overline{w}_{pri}=1.7358$
$\ln\overline{w}_{pub}-\ln\overline{w}_{pri}$		0.2132 **		0.4140 **
$(\overline{X}^{pub}-\overline{X}^{pri})\beta^{*}$	0.2270 **	0.2215 **	0.1892 **	0.3802 **
	(106.0 %)	(104.0 %)	(46.0 %)	(92.0 %)
$\overline{X}^{pri}(\beta^{*}-\beta^{pri})$	-	-0.0082	-	0.0338
		(-4.0%)		(8.0 %)
$\overline{X}^{pub}(\beta^{pub}-\beta^{*})$	-0.0138	-	0.2248 **	
	(-6.0 %)		(54.0 %)	-

Note: Cross-section weights applied. *,**, and *** denote significance at 10%, 5% and 1% respectively. Standard errors in parentheses, where OLS standard errors are robust and the s.e. for the remaining models are bootstrapped. (N) refers to normal, (P) to percentile, and (BC) to bias corrected confidence intervals (1000 replications).

Table 4. Pay gap between public and private sector: Simple probit model.

	Males		Females	
	$\beta^* = \beta^{pri}$	$\beta^* = \beta^{pub}$	$\beta^* = \beta^{pri}$	$\beta^* = \beta^{pub}$
Predicted log wages	$\ln \overline{w}_{pub} = 2.2392$	$\ln \overline{w}_{priv} = 2.0263$	$\ln \overline{w}_{pub} = 2.1516$	$\ln \overline{w}_{priv} = 1.7369$
$\ln \overline{w}_{pub} - \ln \overline{w}_{pri}$		0.2128 **		0.4147 **
$(\overline{X}^{pub} - \overline{X}^{pri})\beta^*$	0.2950 **	0.4198 **	0.2763 **	0.4732 **
	(139.0 %)	(197.0 %)	(67.0 %)	(114.0 %)
$\overline{X}^{pri}(\beta^* - \beta^{pri})$	-	-0.9866 **	-	-0.2906
		(-463.0 %)	-	(-70.0 %)
$\overline{X}^{pub}(\beta^{pub} - \beta^*)$	-0.8619 **	-	-0.0936	
	(-405.0 %)	-	(-23.0 %)	-
Selection (due to differences in		0.7797 **		0.2321
the selection terms)		(366.0 %)		(56.0 %)
Sector choice		0.7389 **		0.2633 ** (BC)
Participation		-0.0408		-0.0312
Selectivity corrected wage gap		-0.5668		0.1826

Note: Cross-section weights applied. *,**, and *** denote significance at 10%, 5% and 1% respectively. Standard errors in parentheses, where OLS standard errors are robust and the s.e. for the remaining models are bootstrapped. (N) refers to normal, (P) to percentile, and (BC) to bias corrected confidence intervals (1000 replications).

Table 5. Pay gap between public and private sector: Bivariate probit model

	Males		Females	
	$\beta^{*} = \beta^{pri}$	$\beta^{*} = \beta^{pub}$	$\beta^{*} = \beta^{pri}$	$\beta^{*} = \beta^{pub}$
Predicted log wages	$\ln \overline{w}_{pub} = 2.2400$	$\ln \overline{w}_{pri} = 2.0258$	$\ln \overline{w}_{pub} = 2.1518$	$\ln \overline{w}_{pri} = 1.7365$
$\ln \overline{w}_{pub} - \ln \overline{w}_{pri}$		0.2142 **		0.4152 **
$(\overline{X}^{pub} - \overline{X}^{pri})\beta^{*}$	0.2913 **	0.4194 **	0.2686 **	0.4629 **
	(136.0 %)	(196.0 %)	(65.0 %)	(112.0 %)
$\overline{X}^{pri}(\beta^{*} - \beta^{pri})$	-	-1.033 **	-	-0.2501
		(-482.0 %)		(-60.0 %)
$\overline{X}^{pub}(\beta^{pub} - \beta^{*})$	-0.9103 **	-	-0.0559	-
	(-424.0 %)		(-13.0 %)	
Selection (due to differences		0.8332 **		0.2025
in the selection terms)		(388.0 %)		(48.0 %)
Selection choice		0.7740 **		0.2466 ** (BC)
Participation		0.0599		-0.0441
Selectivity corrected wage gap		-0.6189 ** (BC)		0.2127

Note: Cross-section weights applied. *,**, and *** denote significance at 10%, 5% and 1% respectively. Standard errors in parentheses, where OLS standard errors are robust and the s.e. for the remaining models are bootstrapped. (N) refers to normal, (P) to percentile, and (BC) to bias corrected confidence intervals (1000 replications).

component decreases substantially, suggesting a male wage premium, though not for the public but for private sector. In the OLS model this premium was counterbalanced by unobserved heterogeneity, which highlights the importance of a separate treatment of the selection component. Note that the selection effect is driven by the sector choice rather than the participation choice. Furthermore, it becomes apparent that even if observable characteristics and compensation were identical in the two sectors, male public employees would still earn more due to differences in unobservable productivity related characteristics.

Tables 4 and 5 report the *selectivity corrected* wage differentials in the bottom row. Interestingly, only the male differential is significant and indicates that correcting for non-random assignments into labour force and sector, the overall wage gap becomes negative. For females, however, higher productivity characteristics of public sector workers are fully outweighed by lower compensation differences.

Hence, not accounting for selection bias may explain a large part of the male differences compared to other studies for the UK or Scotland. For example, Bender (2003) controls for sector selection but the unexplained part in the decomposition does not separate these effects out which may lead to an overestimate of the public sector premium. Elliott et al. (2004) do not account for selection in the first place and results may suffer from the same problem.

Figure 1 has shown that there are not only differences in average wages but also in the distribution of earnings. This is particularly pronounced for women. The above decomposition, though appealing in its simplicity, fails to capture the distribution of the unexplained earnings gap. However, employing a technique developed by Jenkins (1994) shows that the results (with few exceptions) also hold along the pay distribution.[16]

Bringing average wages in the public sector more in line with market rates therefore might disproportionately affect individuals at the lower end of the earnings distribution. However, the direction of change is very different by gender. While men will actually benefit from such a policy, women will be affected negatively. As a result, policy makers face a trade-off between efficiency and equity similar to the one identified by Elliott et al. (2004).

[16] Results are not reported here but can be obtained from the author. In summary, unexplained earnings differentials are small along the lower part of the male distribution suggesting a minor public sector premium. Yet, at the top end there is indication of a private sector premium that seems to dominate the average wage analysis. The reverse holds for females and at the lower end of the distribution there is indication of a small private sector premium. This is very much in line with the actual distributions shown in Figure 1.

Eventually, given the predicted wages, a structural switching regression

$$S^* = \delta(\ln \hat{w}_{1,i} - \ln \hat{w}_{2,i}) + B_i^{'}\mu + v_i^* \tag{12}$$

can be estimated, where $v_i^* = \delta(\varepsilon_{1,i} - \varepsilon_{2,i}) - v_i$ and $\ln \hat{w}_{1,i}$ and $\ln \hat{w}_{2,i}$ are the predicted wages for the public and private sector respectively (Maddala 1983, van der Gaag and Vijverberg 1988 and Hartog and Oesterbeek 1993). For public sector workers the private sector wage is a counterfactual and vice versa for private sector employees.

Table 6 reports the marginal effects of both the reduced form (4) and the structural switching regression (12) which takes expected earnings differences in the sector choice explicitly into consideration. The vector of regressors is identical to equation (4) except that the difference in expected predicted wages in the public and private sector for an employee have been included as an instrumental variable. Again, predicted wages are based on the single selection correction specification for sector choice and simple OLS for males and females respectively.[17]

This specification requires some remarks. Equation (12) contains two estimated regressors. Lee (1979) showed that the resulting coefficients δ and μ are consistent. However, the standard errors are incorrect and need to be adjusted (Maddala 1983). Rather than recovering the corrected variance-covariance matrix, bootstrapping is applied once more as described above.

Furthermore, recall that the set of characteristics for wage equation and sector choice are identical except for the two variables on union membership and union perception. Hence, the potential wage differential is already implicitly included in the reduced form probit estimations. Yet, explicitly including the estimated wage gap will have two effects. First, since the earnings difference is a non-linear construct it will substitute for higher order terms. However, since all but one regressor are dummy variables and the only continuous variable is tenure which is included in both as a first and second order term, this should be of lesser concern. Secondly, it will net-out the earnings effect from the coefficients and the remaining effect can be seen as everything that impacts on the sector choice except earnings.

[17] Note, however, that the main results are not sensitive against the model specification used to predict the expected wages. For example, the double sample selection specification for males yields almost exactly the same results as the single sample selection one.

Table 6. Marginal effects of the reduced form and switching regression for males and females

	Males		Females	
	Reduced form	Structural form	Reduced form	Structural form
Age 16-20	-0.149	-0.216	-0.246	0.229
	(0.032)***	(1.824)	(0.084)***	(1.684)
Age 21-30	-0.035	0.009	-0.267	0.354
	(0.038)	(0.663)	(0.049)***	(1.239)
Married	0.036	0.011	-0.054	0.167
	(0.034)	(0.478)	(0.051)	(0.648)
Higher degree	0.029	0.595	-0.082	0.204
	(0.069)	(1.709)	(0.111)	(1.397)
A-level	0.013	0.112	-0.095	0.564
	(0.036)	(0.563)	(0.058)	(1.054)
O-level	-0.021	-0.038	-0.079	0.683
	(0.038)	(0.601)	(0.057)	(1.358)*** (P) (BC)
Small	-0.125	-0.388	-0.230	-0.431
	(0.035)***	(1.463)*** (BC)	(0.060)***	(1.071)***
Medium	-0.056	-0.555	-0.197	-0.710
	(0.035)	(1.192)*** (BC)	(0.061)***	(0.902)
Opinion on union	0.253	0.380	0.225	0.047
(very important)	(0.069)***	(0.579)***	(0.091)**	(0.277)
Opinion on union	0.228	0.381	0.286	0.171
(important)	(0.046)***	(0.604)***	(0.069)***	(0.223)*** (P) (BC)
Opinion on union	-0.013	-0.074	0.177	0.103
(neutral)	(0.056)	(0.611)	(0.091)*	(0.264)
Job tenure	0.014	0.027	0.029	0.025
	(0.006)**	(0.090)	(0.011)***	(0.155)
Job tenure sq	-0.000	-0.001	-0.001	-0.002
	(0.000)	(0.004)	(0.001)	(0.007)
Professional	0.057	0.229	0.447	0.497
	(0.076)	(1.087)	(0.106)***	(2.274)
Managerial	0.134	0.312	0.489	-0.304
	(0.056)**	(0.772)	(0.074)***	(1.693)

Table 6. (Continued) Marginal effects of the reduced form and switching regression for males and females

	Males		Females	
	Reduced form	Structural form	Reduced form	Structural form
Skilled non-manual	0.141	-0.022	0.169	-0.496
	(0.065) **	(0.725)	(0.091) *	(1.658)
Skilled manual	-0.096	0.111	0.126	-0.006
	(0.042) **	(0.828)	(0.123)	(2.013)
Wage differential		1.322		2.964
		(2.97) *** (P) (BC)		(3.029) ***
Observations		677		516

Notes: The male model is based on the univariate probit specification with sample selection correction for sector choice only and the female model on simple OLS. *,**, and *** denote significance at 10%, 5% and 1% respectively.

The most interesting variable is certainly the predicted wage differential between the public and private sector. Clearly, there is evidence that wage differentials impact significantly on the sector choice. Both the male and female effects on the expected wage differential between the public and private sector are positive and highly significant. The marginal effect seems to be by far the largest compared to other regressors. This is similar to what Hartog and Oesterbeek (1993) find in their study for the Netherlands using a MLE approach and Lee (1978) for the union/non-union decision. However, in both cases the coefficient rather than the marginal effect is reported.

The interpretation of the marginal effect is cumbersome. Given that the predicted wage differential is the difference in two log terms, a log percentage change in the predicted wage differential increases the probability of choosing public sector employment by roughly 1.3 and 2.9 % for men and women respectively.[18] This is a

[18] Neither Lee (1978) nor van der Gaag and Vijverberg (1988) put a meaning on the coefficient of the predicted wage differential. On the other hand, Hartog and Oesterbeek (1993) interpret the coefficient as selection elasticity.

rather small effect compared to other variables such as firm size or union perception.[19]

V. Implications and conclusions

Post-devolution public-private wage differentials in Scotland are studied using newly available data from the BHPS to establish whether earnings differentials exist and differ from the rest of the UK. In contrast to other UK-wide and regional studies this paper controls for both sample selection from the participation decision and the sector choice.

The wage gap is shown to be 10 % for males and 24 % for females. For males this is mainly due to differences in productive characteristics and sector selection. For females, the picture is more ambiguous. In contrast to other UK studies, there is evidence of a male private sector wage premium that is mainly due to sector selection emphasising the need to control for section bias. It is shown that wage differentials between sectors are an important driving force for sectoral assignment. Furthermore, results vary along the pay distribution. Evidence shows that private and public sector premiums for men and women respectively are mainly an issue in the upper parts of the earnings distributions.

While sector selection is expected to be responsible for a large part of the differences in results compared with other UK studies, there may be further explanations. Firstly, despite the devolution, wage setting in the public sector does not match up with the new institutional arrangements and has not significantly changed since 1999. Hence, the Scottish Executive is only partly responsible for any pay differential. Secondly, Scotland faces fierce competition with the rest of the country over skilled labour in both the public and private sector. While earnings in the private sector can adjust to a tight labour market, the public sector often lacks regional flexibility. Therefore, the distinct absence of a male wage public sector premium in Scotland may also be partly the result of adjustments in the private rather than public sector.

This is different for women where findings are more in line with UK wide results. However, the high proportion of women in the public sector in general and in health and education in particular combined with a relative shortage of staff in

[19] However, it has been remarked that the effect at the mean is still likely to be large given the large raw difference in public-private sector wages.

these sectors may go some way in explaining a higher female public sector pay nationwide. In addition, there is relatively more regional independence in wage bargaining for traditional female occupations such as local governments and education.

Alternatively, if one is prepared to believe that the Scottish labour market achieves an efficient allocation of labour, part of the male wage premium found in the private sector may be ascribed to the existence of fringe benefits in the public sector which are not captured in average hourly wages such as higher job security, holidays and pension entitlements. Again, this seems to be different for women and also implies a different fringe benefit culture north of the border. Theoretically, the latter does not seem unreasonable as non-pecuniary benefits may well be a means to increase the attractiveness of Scottish public sector jobs in the absence of full pay autonomy. Yet, even some of these are still determined by Westminster such as the Civil Service pensions scheme, making this explanation less compelling.

However, even if the Scottish Executive were granted the necessary independence in public sector wage setting it is conceivable that it would only ever want to pay market rates in occupations where it is not competing with other UK labour markets. Hence, incentives to set public sector earnings in line with private sector pay may be very small. A pay reform that grants more autonomy to regions may therefore not necessarily remove public sector premiums but runs the risk of introducing new trade-offs between regional efficiency, equity, and national competitiveness. However, given the results in this paper, public sector pay in Scotland may after all not be as far off market rates as suggested by other studies.

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