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ABSTRACT

Public-Private Sector Wage Differentials in Scotland: An Endogenous Switching Model*

The public-private sector wage gap in Scotland in 2000 is analysed using the extension sample of the British Household Panel Survey (BHPS). Employing an endogenous switching model, and testing for double sample selection from the participation decision and sector choice, the unadjusted wage gap is shown to be 10 per cent for males and 24 per cent for females. For males this is mainly due to differences in productive characteristics and selectivity. For females the picture is more ambiguous. Findings also suggest that there exists a substantial wage premium for male private sector employees. 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

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1 Introduction

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 [4] and Gregory and Borland [16].

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 [17]).

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 [16]). 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 [15] 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.

Bellante and Link [1] find that public sector employees are more risk averse than their private sector counterparts. Thus, given the common assumption that markets are dominated by risk averse individuals rather than risk seekers there will be an excess supply of labour to the public sector and wages should adjust to equalise demand and supply.[1]

The validity of theories, however, depends very much on the economic, institutional and political environment. Elliott *et al.* [11] 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

[1] Note, however, that causality in Bellante and Link's [1] study is ambiguous and Gregory and Borland provide an alternative interpretation by arguing that it might well be possible that public sector employers prefer to hire risk averse individuals rather than risk averters who seek employment in the public sector.

when the public sector was expanded significantly due to the nationalisation of former private industries, as Bender [5] 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. Additionally, wage bargaining was further decentralised in the early 1990s, a trend that has gained even more importance in recent years in the aftermath of the devolution of Wales and Scotland. Today, with the exception of the large number of NHS staff and university lecturers, almost all other public occupations in Scotland are negotiated at a Scottish rather than a UK-wide level.

The principle of pay comparability should, therefore, 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. Empirical studies for the UK on sectoral wage differentials are also scarce.[2] Only a small number of studies based on micro-level data have been published in the last two decades or so, of which three have been reviewed in Bender [5] and will only be briefly re-stated here in chronological order.[3]

Employing data from the General Household Survey (GHS) for the years 1983, 1985, and 1987, Rees and Shah [29] estimate the average wage gap for males, using a single equation model with a sector dummy, to be between 9.8 and 11.4 per cent. The difference is even larger for females and lies between 22.3 and 26.3 per cent. A simple Oaxaca [28] decomposition into explained (differences in characteristics) and unexplained (returns to characteristics) parts shows that most of the male gap is due to differences in characteristics between public and private sector employees. Yet, for females their results suggest that the substantial earnings differential is solely due to a positive wage premium.

Disney and Gosling [8] use data from both the GHS and the British Household Panel Survey (BHPS) for the years 1983 and the early 1990s. Simple OLS estimation, with a dummy for public sector employment and conditional on age and education, reveals that the pay differential for males has fallen from 5 to 1 per cent and increased from 11 to 14 per cent for women from the mid- 1980s to the early 1990s. Furthermore, applying quantile regressions on the wage distribution, the authors find an increase in the income dispersion. Interestingly, this increase is not confined to the public sector but occurs in both sectors. However, as one would expect, income inequality in the public sector is less pronounced than in the private.

Using BHPS data and controlling for various personal and job related char-

[2] For a comprehensive review of empirical studies across countries see Bender [4] and Gregory and Borland [16].

[3] There are, however, early studies predominantly based on data from the New Earnings Survey (NES) which do not allow a detailed control for personal and job related characteristics. In particular these are Elliott and Murphy [12], Gregory [15], and Elliott and Duffus [10].

acteristics, Bender and Elliott [3] found that between 1991 and 1994 the overall wage differential between public and private sector increased from 23.2 to 29.4 per cent. In the vast majority of model specifications the gap is explained by differences in characteristics. The study does not distinguish between gender due to the small sample size as a result of the detailed disaggregation of occupations. However, in contrast to the previous studies, Bender and Elliott run separate regressions for public and private sector wages.

Finally, in the most recent study, Bender[5] estimates separate wage equations for public and private sector distinguishing males and females. Furthermore, the author controls for possible sample selection due to the sector choice of employees. In addition to the usual decomposition of the average wages between the sectors, a decomposition technique based on Belman and Heywood [2] is applied that takes differences in the wage distribution into account. Applying data from the British SCELI survey in 1986, Bender finds that males at the low end of the public sector pay distribution are better off than their private counterparts whilst high-paid private sector males earn more than high-paid public employees. Yet, there is no evidence for this "double imbalance" in the female wage distribution. Furthermore, the results of the simple Oaxaca decomposition show that much of the difference in average wages is due to differences in returns rather than characteristics. This stands in contrast to what to Rees and Shah [29] find.[4]

In the following, new data from the BHPS is used to analyse wage differentials in the public and private sector in Scotland. This is believed to be of importance as the devolution and subsequent election of the Scottish Parliament has increased the share of government employment and brought more independence to public sector wage bargaining.

The paper is structured as follows. The next section will describe the data set and provide some descriptive statistics. Section 3 outlines the econometric framework and discusses identification issues. Finally, section 4 reports the results on wage equations and decomposition and discusses their implications.

2 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 recently 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

[4] Even though Bender[5] 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 [18]).

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 1054 males and 1230 females of which 61 per cent of males and 42 per cent of females are participating in the labour market. Of these, roughly 22 per cent of men and 41 per cent of women are employed in the public sector. Table 1 shows the occupational composition within the public sector. 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.

Table 2 reports the development in the Scottish public sector since 1996.[5] While 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 reflected in the female figures although it is less pronounced.

Table 3 shows the age distribution among public and private employees by sex. While both the male and female age distributions in the private sector are positively skewed, the majority of employees in the public sector are relatively older. For example, the cumulative percentage of employees aged 54 and under is 94 per cent in the private but only 89 per cent in the public sector for males.

Furthermore, public sector workers are selected differently into occupations (Table 4). 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 per cent are employed in this occupation compared to only 31 per cent 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.* [6]) wages are expressed in log hourly gross pay.[6] As figure 1 shows, the male pay distribution in the public sector is less spread compared to the

[5]Note that table 2 compares pre-extension with post-extension sample figures and that prior to 1999 the Scottish sample was not representative. Furthermore, these are unweighted numbers which causes minor differences compared to other tables which use cross-sectional weights.

[6]The log hourly gross wage rate is defined as $w = \ln(Paygu/(30/7)(Hs + \alpha Hot))$, 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.

private sector distribution. Furthermore, the mean wage in the former seems to be higher than in the latter. The respective mean log wages in the public and private sector are 2.2 and 2.0 (table 5). 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 per cent significance level. In contrast, the variation in the female public sector pay distribution is smaller compared to men. Once again, the difference in means between public and private sector is positive (0.39) and significant at the 1 per cent level. Hence, on average there is a substantial and significant pay gap for both males and females.

3 Econometric framework

3.1 Wage Equations

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. A popular approach is to control in the wage equation for public sector employment by including a dummy variable (see e.g. Rees and Shah [29]). However, this only captures the level effect of the sector choice on wages and restricts the coefficients to being the same across sectors. Hence, in the following an endogenous switching model is adopted (Lee [23]) that estimates separate wage equations for the public and private sectors.[7] 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 [21]). 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 [25] and Nelson [26]). The recent literature on public-private sector earning differentials has mainly accounted for the latter and widely ignored

[7] Estimating separate wage equations is equivalent to a single equation where every regressor is interacted with the sector dummy.

the former.[8] However, controlling for one type of selection in earnings equations only and ignoring non-labour force participants may still lead to biased estimates (Co *et al.* [7]).

In order to test, and potentially account for, both types of selection, a double sample selection model is fitted following Tunali [32]. 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 u_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 (a) the expected earnings differential is positive and/or (b) personal preferences for public sector employment are strong. These preferences may be correlated with individual characteristics and captured in B.[9]

Since neither latent variable is observable, two index functions are defined. In the case of participation this is

$$\begin{aligned} P_i &= 1 \;\; if \;\; P_i^* > 0 \\ P_i &= 0 \;\; if \;\; P_i^* \leq 0 \end{aligned}$$

where $P_i = 1$ and $P_i = 0$ indicate labour market participation and non-participation respectively. Similarly, for the sector choice

$$\begin{aligned} S_i &= 1 \;\; if \;\; S_i^* > 0 \\ S_i &= 0 \;\; if \;\; S_i^* \leq 0 \end{aligned}$$

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). Yet, the number of parameters to be estimated is rather large. Alternatively, a simple two-step Heckman [21] approach with extended correction terms may be adopted (see e.g. Lee [24], Ham [19], Fishe *et al.* [14] and Tunali [32]). In the first step equations (3) and (4) are estimated and sample selection correction terms are constructed. In the second step (1) and (2) are then estimated via simple OLS including the correction terms as additional regressors.

Two cases can be distinguished. First, assume the participation decision and the sector choice are independent, i.e. $\rho_{uv} = 0$, where ρ is the error correlation

[8] The only exception the author is aware of is Stillman [31] in a study on the Russian labour market.

[9] For a more rigorous theoretical derivation see e.g. van der Gaag and Vijverberg [33].

term between equation (3) and (8). Then, similar to the single selection case, four separate correction terms are constructed based on two probit models for participation and sector choice. In particular, the estimated correction terms for the public sector wage equation are

$$\hat{\lambda}_{i,p1} = \phi(Z_i^{'}\hat{\gamma})/\Phi(Z_i^{'}\hat{\gamma}) \tag{5}$$

$$\hat{\lambda}_{i,s1} = \phi(B_i^{'}\hat{\mu})/\Phi(B_i^{'}\hat{\mu}) \tag{6}$$

and the estimated correction terms for the private sector wage equation are

$$\hat{\lambda}_{i,p2} = \phi(Z_i^{'}\hat{\gamma})/\Phi(Z_i^{'}\hat{\gamma}) \tag{7}$$

$$\hat{\lambda}_{i,s2} = -\phi(B_i^{'}\hat{\mu})/\Phi(-B_i^{'}\hat{\mu}) \tag{8}$$

where $\hat{\lambda}_{i,kj}$ is the correction term for participation and sector choice $(k = p, s)$ respectively for both public and private employment $(j = 1, 2)$ and $\hat{\lambda}_{i,p1} = \hat{\lambda}_{i,p2}$ since public or private wages are only observed for employed individuals; ϕ and Φ are the univariate standard normal density and distribution functions respectively.

In general, the two selection processes (3) and (4) may not be independent. Hence, an alternative approach is to estimate them using a bivariate probit (Ham [19], Tunali, [32]) assuming the error correlation term to be $\rho_{uv} \neq 0$. In that case (ε_j, u, v) are jointly normally distributed with mean zero and covariance matrix

$$\Sigma_j = \begin{bmatrix} \sigma_{jj}^2 & \sigma_{jjv} & \sigma_{jju} \\ & \sigma_v^2 & \sigma_{vu} \\ & & \sigma_u^2 \end{bmatrix}$$

where σ_v^2 and σ_u^2 are normalised to unity for identification purposes following standard practice.[10] Maximum likelihood of the bivariate probit leads to four sample selection correction terms

[10] 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.* [7] for details; see Koop and Poirier [22] for a discussion on these identification issues).

$$\hat{\lambda}_{i,p1} = \phi(Z_i^{'}\hat{\gamma})\Phi\left[\frac{B_i^{'}\hat{\mu} - \rho Z_i^{'}\hat{\gamma}}{(1-\rho^2)^{1/2}}\right] \times F(B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, \rho)^{-1} \quad (9)$$

$$\hat{\lambda}_{i,s1} = \phi(B_i^{'}\hat{\mu})\Phi\left[\frac{Z_i^{'}\hat{\gamma} - \rho B_i^{'}\hat{\mu}}{(1-\rho^2)^{1/2})}\right] \times F(B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, \rho)^{-1} \quad (10)$$

$$\hat{\lambda}_{i,p2} = \phi(Z_i^{'}\hat{\gamma})\Phi\left[-\frac{(B_i^{'}\hat{\mu} - \rho Z_i^{'}\hat{\gamma})}{(1-\rho^2)^{1/2})}\right] \times F(-B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, -\rho)^{-1} \quad (11)$$

$$\hat{\lambda}_{i,s2} = -\phi(B_i^{'}\hat{\mu})\Phi\left[\frac{Z_i^{'}\hat{\gamma} - \rho B_i^{'}\hat{\mu}}{(1-\rho^2)^{1/2}}\right] \times F(-B_i^{'}\hat{\mu}, Z_i^{'}\hat{\gamma}, -\rho)^{-1} \quad (12)$$

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.

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

$$E\left(\ln w_{1,i}|X_i^{'}, P_i = 1, S_i = 1\right) = X_{1,i}^{'}\beta_1 + \sigma_{11}\rho_{1u}\hat{\lambda}_{i,p1} + \sigma_{11}\rho_{1v}\hat{\lambda}_{i,s1} \quad (13)$$

$$E\left(\ln w_{2,i}|X_i^{'}, P_i = 1, S_i = 0\right) = X_{2,i}^{'}\beta_2 + \sigma_{22}\rho_{2u}\hat{\lambda}_{i,p2} + \sigma_{22}\rho_{2v}\hat{\lambda}_{i,s2} \quad (14)$$

Depending on the assumption on ρ_{uv} the correction terms are either as in equations (5)-(8) if $\rho_{uv} = 0$ or (9)-(12) if $\rho_{uv} \neq 0$.

3.2 Decomposition

Once wages are consistently estimated, differences in public and private sector pay can be decomposed into several components. The wage gap can be split into three terms (Neuman and Oaxaca [27]) such that

$$\begin{aligned} \ln \bar{w}_1 - \ln \bar{w}_2 &= (X_{1,i}^{'}\hat{\beta}_1 + \sigma_{11}\rho_{1u}\hat{\lambda}_{i,p1} + \sigma_{11}\rho_{1v}\hat{\lambda}_{i,s1}) \\ &\quad -(X_{2,i}^{'}\hat{\beta}_2 + \sigma_{22}\rho_{2u}\hat{\lambda}_{i,p2} - \sigma_{22}\rho_{2v}\hat{\lambda}_{i,s2}) \\ &= \hat{\beta}_1(\bar{X}_1 - \bar{X}_2) + \bar{X}_2^{'}(\hat{\beta}_1 - \hat{\beta}_2) \\ &\quad +(\sigma_{11}\rho_{1u}\hat{\lambda}_{i,p1} + \sigma_{11}\rho_{1v}\hat{\lambda}_{i,s1}) \\ &\quad -(\sigma_{22}\rho_{2u}\hat{\lambda}_{i,p2} + \sigma_{22}\rho_{2v}\hat{\lambda}_{i,s2}) \end{aligned} \quad (15)$$

where $\ln \bar{w}$ is the predicted mean log wage, $\bar{X}$ 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 [13]).

Similar to the simple Oaxaca decomposition (Oaxaca [28]) the term $\hat{\beta}_1(\bar{X}_1 - \bar{X}_2)$ is the explained and $\bar{X}_2^{'}(\hat{\beta}_1 - \hat{\beta}_2)$ the unexplained part of the predicted mean wage gap. However, it is *a priori* unclear how to tread the selection terms in equation (15). 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 [30]).[11]

4 Estimation Results

4.1 Identification and Variable Choice

Estimating the above two step model requires some identification assumptions on the coefficients and the covariance parameters (Tunali [32]). First, as already stated, σ_v^2 and σ_u^2 are normalised to 1. Depending on whether the selection 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 [5] and Hartog and Oosterbeek [20]).

Even though the BHPS contains questions on e.g. 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 [16]). 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.[12] Since union

[11] Equation (15) 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.

[12] In particular, individuals are asked whether strong trade unions are needed to protect the

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.

Tables 6 and 7 report the mean characteristics of the sample separately by gender and for public sector workers. Several features are worth mentioning. 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.

4.2 Estimation Results

The results of the bivariate probit alongside their univariate counterparts by gender are reported in tables 8 and 9. Both for males and females the results do not support a simultaneous estimation of the participation and sector choice decision. 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.* [14] and Tunali [32]).

In particular, 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 [4] and van der Gaag and Vijverberg [33]).

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

working conditions and wages of employees. Four possible answers can be given which range from strong agreement to disagreement.

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. 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.[13] Education and job tenure on the other hand do not impact significantly on either decision with some exceptions in the univariate probit case.[14]

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 without offspring. 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 10 and 11 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 [9] 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.[15] For each sample all coefficients are re-estimated and then used to derive standard errors and confidence intervals.

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

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

[14] 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.

[15] In principal it is not clear whether to keep the sub-sample sizes (proportion of public/private and participant/non-participant) constant or draw from the entire sample N. See Appendix **??** for some Monte Carlo studies on this issue. In the following the latter method has been used.

confidence intervals are roughly identical. Hence, all three intervals will be very similar for an approximately normally distributed bootstrap statistic and a small bias.

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 effect on male public sector wages but 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 (13) and (14)), 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.[16]

4.3 Wage Gap and Switching Regression

Given the above results, predicted log wages in the public and private sector can be estimated consistently. Tables 12 to 14 report predicted log wages, differences in predicted log wages and the decomposition of predicted log wages into various parts according to equation (15) by gender and for three different model specifications. Furthermore, results are shown for two different weighting specifications, $\beta^* = \beta_1 = \beta^{pub}$ and $\beta^* = \beta_2 = \beta^{pri}$. 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 per cent for males and 24 per cent for females, which lies well in line with Rees and Shah [29] but is slightly higher than Bender's [5] 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

[16] The reported correlation terms and the standard deviations are constructed following Tunali [32].

reduce the unadjusted gap, though the term is statistically insignificant. That is, differences in the productive and job-related characteristics of public sector employees more than counterbalance the lower public sector compensation.[17].

Results are more sensitive against the weighting scheme for women. Using the pubilc wage structure the results are fairly similar to the male ones. However, if the private sector wage structure is used as weight, only between 46 and 66 per cent of the wage difference is explained by differences in characteristics.

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 component decreases substantially, suggesting a male wage premium though, not for the public but for private sector. Assuming a well-functioning labour market, the premium may be a - though not perfect - measure of the fringe benefits in the public sector or the magnitude of the excess supply to public sector jobs.

In the OLS model this premium was counterbalanced by unobservable characteristics, 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 13 and 14 also 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, there is no evidence that suggests a wage premium for public sector employees both for men and women.

Eventually, given the predicted wages, a *structural* switching regression

$$S_i^* = \delta(\ln \hat{w}_{1,i} - \ln \hat{w}_{2,i}) + B_i^{'}\mu + v_i^* \qquad (16)$$

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, [25], van der Gaag and Vijverberg, [33] and Hartog and Oesterbeek, [20]). For public sector workers the private sector wage is a counterfactual and vice versa for private sector employees.

[17] Note that Rees and Shah [29] find the same for males but not for females while Bender [5] 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 15 reports the marginal effects of both the reduced form and the *structural* switching regression 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. Predicted wages are based on the single selection correction specification for sector choice and simple OLS for males and females respectively.[18]

This specification requires some remarks. Equation (16) contains two estimated regressors. Lee ([24]) showed that the resulting coefficients δ and μ are consistent. However, the standard errors are incorrect and need to be adjusted (Maddala, [25]). 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 in tables 8 and 9. Yet, explicitly including the estimated wage gab 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 continues variable is includes both as 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.

The predicted probabilities of public sector employment are 22 and 41 per cent for males and females respectively, which is identical to the predicted probabilities for the univariate case in tables 8 and 9 and the share of public sector workers in the data set. Yet, there are changes compared to tables 8 and 9. Most pronounced are the differences in marginal effects for the occupational variables. While many occupational variables in the reduced form regression exhibit a significant impact, this is not the case in the structural form specification. Other variables seem to be almost unaffected by the inclusion of the wage difference such as the perception of unions.

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 ([20]) find in their study for the Netherlands using a MLE approach and Lee ([23]) 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

[18] 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.

change in the predicted wage differential increases the probability of choosing public sector employment by roughly 1.3 and 2.9 per cent for men and women respectively.[19] This is a rather small effect compared to other variables such as firm size or union perception.

Interestingly, some characteristics which have been significant in the reduced form equation but insignificant in the structural equation have also been significant in the wage equation such as age. However, the opposite holds true as well for variables such as job tenure, which is insignificant in both the wage and the structural equation but significant in the reduced form regression. This is, not controlling for the earnings difference obscures other non-pecuniary effects in the sector choice equation without affecting the overall predicted probability especially if the earnings effect is strong.

5 Conclusion and Implications

A long-standing interest of labour economists has been earnings differentials between public and private sectors. While there are empirical studies on the UK-wide level with conflicting evidence of a wage premium in the public sector, thus far no such study has been conducted on a regional level. Employing data for 2000 from the newly available extension sample of the BHPS, potential earnings differentials for Scotland are analysed. This is believed to be insightful for at least two reasons. First, in the aftermath of devolution and the establishment of the Scottish Parliament in 1999, the occupational structure of the public sector has changed. Second, more power over wage setting has been granted to local authorities. Both can be thought to have an impact on the wage structure in the public sector.

An endogenous switching model is estimated rather than a single wage equation with a sector dummy. Additionally, in order to correct for potential sample selection bias two additional models are estimated alongside the simple OLS specification, controlling for both sample selection from the participation decision and the sector choice.

Surprisingly, sample selection from the labour force participation decision for both males and females is not an issue. Furthermore, there is also no indication that the sector choice and the participation decision are jointly correlated. However, there is strong evidence that the former impacts significantly on the male wage equation in the private and public sector and, hence, on pay differentials.

Controlling for various personal and job related characteristics, the estimated wage differential between the public and private sector is 10 per cent for men and 24 per cent for women, which coincides with the unconditional wage gap. Yet, given this pay differential the question arises whether it should be ascribed to differences in observed and unobserved productivity and job related characteristics or returns to these characteristics.

[19] Neither Lee ([23]) nor van der Gaag and Vijverberg ([33]) put a meaning on the coefficient of the predicted wage differential. On the other hand, Hartog and Oesterbeek ([20]) interpret the coefficient as selection elasticity.

Decomposing the wage gap into various components depending on the model specification (selection correct or uncorrected) shows that higher male public sector wages can be explained by differences in characteristics both observed and unobserved and are not due to a wage premium; quite the contrary, a private sector wage premium is found, which is, however, more than compensated for by superior productivity and job related characteristics and unobservables. The results emphasise the importance of a separate treatment of sample selection terms since they are otherwise likely to obscure the unexplained component in the decomposition. In contrast, for females there is some evidence of a wage premium in the public sector.

Finally, using the expected counterfactual wage differential to estimate the structural switching regression of the sector choice shows that for both men and women earnings differentials play a prominent role in the sector selection.

The policy implications are two-fold. First, and most importantly, for men there is no evidence for a positive earnings discrimination in the Scottish public sector. Differences in pay are solely due to differences in personal and job related characteristics and the way employees select themselves into sectors.

If one believes that the Scottish labour market achieves an efficient allocation of employment, 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 the hourly wages used in the analysis. Thus higher wages in the private sector may reflect a flexible earnings structure rather than premiums and in contrast to what Bender [5] concludes for the whole of the UK, the principal of pay comparability seems to have been successful.

However, public sector spending has been on the decrease since the early 1990s and only recently picked up again. Hence, given the time series nature of the data used in this analysis, lower conditional public sector wages may simply reflect this decrease in spending rather than a private sector wage premium. Second, there is ambiguous evidence of a wage premium for females in the public sector which cannot be explained by differences in characteristics.

Hence, the political aim to achieve an efficient labour allocation and comparability in pay in the public and private sector has partly been met. However, further reforms may be required especially in those occupations predominantly occupied by women.

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Public sector composition (%)	Males	Females
Civil servant/central govt	21.48	11.21
Local govt/town hall	56.38	51.87
NHS or higher education	17.45	35.98
Nationalised Industry	4.70	0.93

Table 1: Composition of public sector employment by gender in 2000. Cross-sectional weights applied. (BHPS)

Public sector composition (%)						
Males	1996*	1997*	1998*	1999	2000	2001
Civil servant/central govt	18.00	17.07	32.35	21.19	21.94	22.70
Local govt/town hall	42.00	51.22	41.18	49.67	55.48	49.08
NHS or higher education	24.00	21.95	23.53	22.52	17.42	22.09
Nationalised Industry	16.00	9.76	2.94	6.62	5.16	6.13
Females						
Civil servant/central govt	13.43	13.43	10.53	9.77	11.06	10.83
Local govt/town hall	52.24	46.27	50.88	51.16	52.21	51.25
NHS or higher education	34.33	40.30	38.60	39.07	35.84	37.08
Nationalised Industry	-	-	-	-	0.88	0.83
Total						
Males	26.46	18.64	17.99	21.75	20.23	21.68
Females	41.61	39.18	33.73	37.85	38.63	41.17

Table 2: Composition of public sector employment by gender 1996-2001. *Numbers prior to the Scottish extension sample. Unweighted figures (BHPS)

Age group	Freq.	Percent	Freq.	Percent
Males	Private		Public	
16-19	32	6.06	2	1.34
20-24	58	10.98	9	6.04
25-29	70	13.26	18	12.08
30-34	76	14.39	17	11.41
35-39	93	17.61	26	17.45
40-44	78	14.77	20	13.42
45-49	44	8.33	21	14.09
50-54	47	8.90	20	13.42
55-59	22	4.17	11	7.38
60-64	8	1.52	5	3.36
Total	528	100.00	149	100.00
Females	Private		Public	
16-19	26	8.61	4	1.87
20-24	44	14.57	15	7.01
25-29	56	18.54	17	7.94
30-34	32	10.60	36	16.82
35-39	33	10.93	35	16.36
40-44	42	13.91	34	15.89
45-49	37	12.25	43	20.09
50-54	22	7.28	17	7.94
55-59	10	3.31	13	6.07
Total	302	100.00	214	100.00

Table 3: Age structure of public sector employment by gender in 2000. Cross-section weights applied. (BHPS)

Occupation	Freq.	Percent	Freq.	Percent
Males	Private		Public	
Professional	36	6.82	13	8.72
Managerial	131	24.81	58	38.93
Skilled non-manual	80	15.15	33	22.15
Skilled manual	197	37.31	24	16.11
Unskilled	84	15.91	21	14.09
Total	528	100.00	149	100.00
Females	Private		Public	
Professional	8	2.65	9	4.21
Managerial	85	28.15	135	63.08
Skilled non-manual	120	39.74	48	22.43
Skilled manual	38	12.58	9	4.21
Unskilled	51	16.89	13	6.07
Total	302	100.00	214	100.00

Table 4: Occupational composition of public sector employment by gender in 2000. Cross-section weights applied. (BHPS)

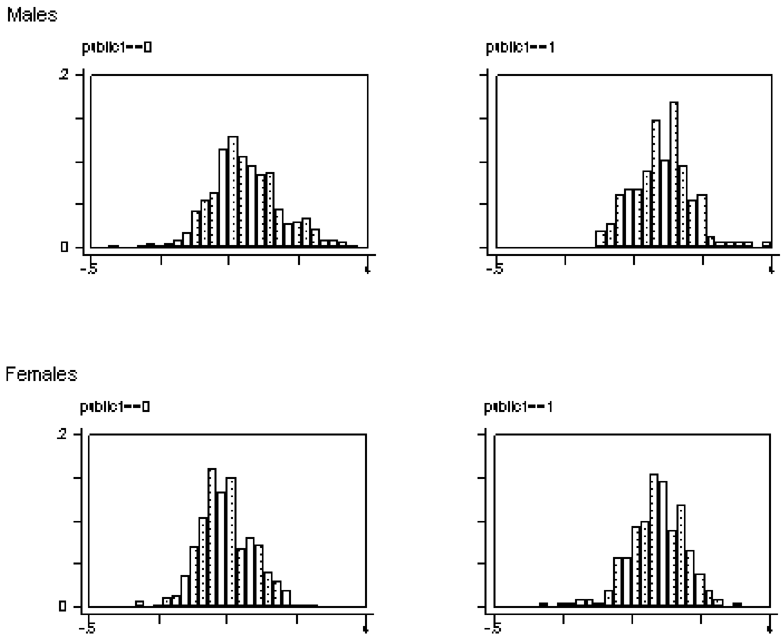

Figure 1: Log hourly wages in the public and private sector by gender in 2000.

	Male		Female	
	Public	Private	Public	Private
Mean log hourly wages	2.24	2.03	2.15	1.74
Standard deviation	(0.039)	(0.025)	(0.032)	(0.025)
Difference in means (t-test)	0.194 (3.75)***		0.415 (10.16)***	

Table 5: Mean test (t-test) for log hourly wages in the public and private sector by gender 2000.

Variable	All		Public	
	Mean	Std. Dev.	Mean	Std. Dev.
Age	37.179	11.092	40.503	10.949
Age 16-20	0.069	0.254	0.020	0.141
Age 21-30	0.210	0.407	0.174	0.381
Married	0.560	0.497	0.638	0.482
Children aged 0-11	0.514	0.833	0.490	0.819
Children aged 12-18	0.229	0.512	0.195	0.529
Higher degree	0.047	0.212	0.060	0.239
A-level	0.273	0.446	0.295	0.458
O-level	0.232	0.422	0.168	0.375
Job tenure	5.893	6.740	7.797	7.348
Member of union	0.318	0.466	0.725	0.448
Opinion on union (very important)	0.191	0.393	0.235	0.425
Opinion on union (important)	0.458	0.499	0.591	0.493
Opinion on union (neutral)	0.176	0.381	0.074	0.262
Professional	0.072	0.259	0.087	0.283
Managerial	0.279	0.449	0.389	0.489
Skilled non-manual	0.167	0.373	0.221	0.417
Skilled manual	0.326	0.469	0.161	0.369
Log hourly wage	2.079	0.563	2.230	0.480
Small	0.412	0.493	0.282	0.451
Medium	0.375	0.485	0.423	0.496
Public	0.220	0.415		
N	677		149	

Table 6: Mean characteristics of variables used in the analysis for all employed and separately for employees in the public sector (males) in 2000. Cross-section weights applied.

Variable	All Mean	Std. Dev.	Public Mean	Std. Dev.
Age	36.430	10.926	39.266	9.841
Age 16-20	0.074	0.261	0.019	0.136
Age 21-30	0.240	0.428	0.150	0.357
Married	0.498	0.500	0.542	0.499
Children aged 0-11	0.302	0.599	0.322	0.660
Children aged 12-18	0.254	0.539	0.248	0.530
Higher degree	0.043	0.202	0.056	0.231
A-level	0.250	0.433	0.206	0.405
O-level	0.244	0.430	0.206	0.405
Job tenure	4.800	5.529	6.245	6.348
Member of union	0.403	0.491	0.715	0.452
Opinion on union (very important)	0.163	0.370	0.168	0.375
Opinion on union (important)	0.494	0.500	0.561	0.497
Opinion on union (neutral)	0.178	0.383	0.150	0.357
Professional	0.033	0.179	0.042	0.201
Managerial	0.426	0.495	0.631	0.484
Skilled non-manual	0.326	0.469	0.224	0.418
Skilled manual	0.091	0.288	0.042	0.201
Log hourly wage	1.921	0.500	2.164	0.468
Small	0.455	0.498	0.402	0.491
Medium	0.357	0.479	0.332	0.472
Public	0.415	0.493		
N	516		214	

Table 7: Mean characteristics of variables used in the analysis for all employed and separately for employees in the public sector (females) in 2000. Cross-section weights applied.

	Single Probit Public	Participation	Bivariate Probit Public	Participation
Constant	-1.5134	0.1184	-1.7763	0.1681
	(0.2811)***	(0.0892)	(0.3069)***	(0.0719)**
Age 16-20	-0.9022	-0.4720	-1.0514	-0.4534
	(0.3459)***	(0.1440)***	(0.3499)***	(0.1329)***
Age 21-30	-0.1463	0.1299	-0.1042	0.1216
	(0.1645)	(0.1143)	(0.1922)	(0.0905)
Married	0.1464	0.2593	0.2585	0.1967
	(0.1365)	(0.0980)***	(0.1690)	(0.0745)***
Higher degree	0.1086	0.4603	0.1547	0.3401
	(0.2521)	(0.2305)**	(0.2708)	(0.2430)
A-level	0.0494	0.2116	0.0807	0.1232
	(0.1419)	(0.0994)**	(0.1459)	(0.0767)
O-level	-0.0864	0.1507	-0.1383	0.0305
	(0.1575)	(0.1033)	(0.1635)	(0.0780)
Children aged 0-11		0.2264		0.2560
		(0.0599)***		(0.0379)***
Children aged 12-18		0.0828		0.1695
		(0.0820)		(0.0616)***
Small	-0.5207		-0.3718	
	(0.1519)***		(0.1623)**	
Medium	-0.2308		-0.1249	
	(0.1490)		(0.1573)	
Opinion on union (very important)	0.8275		0.6788	
	(0.2023)***		(0.2215)***	
Opinion on union (important)	0.8850		0.8033	
	(0.1765)***		(0.2195)***	
Opinion on union (neutral)	-0.0511		0.0028	
	(0.2306)		(0.2322)	
Job tenure	0.0574		0.0367	
	(0.0244)**		(0.0263)	
Job tenure sq	-0.0014		-0.0007	
	(0.0010)		(0.0010)	
Professional	0.2096		0.0558	
	(0.2599)		(0.2548)	
Managerial	0.4913		0.5720	
	(0.2015)**		(0.2170)***	
Skilled non-manual	0.4861		0.4343	
	(0.1877)***		(0.1985)**	
Skilled manual	-0.4132		-0.3332	
	(0.1953)**		(0.1979)*	
ρ			0.5377	
			(0.5596)	
Observations	677	1054	1054	

Table 8: Estimation results of the bivariate probit and single probits for males. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Robust standard errors in parantheses.

	Single Probit Public	Participation	Bivariate Probit Public	Participation
Constant	-0.9889	-0.0497	-0.8074	-0.0660
	(0.3171)***	(0.0785)	(0.4196)*	(0.0732)
Age 16-20	-0.7519	-0.4572	-0.6495	-0.4019
	(0.3319)**	(0.1465)***	(0.3559)*	(0.1432)***
Age 21-30	-0.7644	0.2805	-0.7737	0.2514
	(0.1592)***	(0.0970)***	(0.1820)***	(0.0900)***
Married	-0.1406	0.0706	-0.2214	0.0882
	(0.1344)	(0.0824)	(0.1428)	(0.0761)
Higher degree	-0.2226	0.7922	-0.1225	0.9450
	(0.3140)	(0.2556)***	(0.3449)	(0.1919)***
A-level	-0.2549	0.2190	-0.2495	0.1297
	(0.1593)	(0.0970)**	(0.1780)	(0.0903)
O-level	-0.2099	0.0827	-0.2803	0.0581
	(0.1554)	(0.0922)	(0.1692)*	(0.0868)
Children aged 0-11		-0.5143		-0.4536
		(0.0503)***		(0.0497)***
Children aged 12-18		-0.0012		0.0204
		(0.0689)		(0.0649)
Small	-0.6144		-0.6124	
	(0.1662)***		(0.1839)***	
Medium	-0.5335		-0.6236	
	(0.1740)***		(0.1852)***	
Opinion on union (very important)	0.5716		0.5627	
	(0.2335)**		(0.2482)**	
Opinion on union (important)	0.7642		0.8462	
	(0.1940)***		(0.2092)***	
Opinion on union (neutral)	0.4524		0.5333	
	(0.2313)*		(0.2499)**	
Job tenure	0.0758		0.0544	
	(0.0286)***		(0.0298)*	
Job tenure sq	-0.0015		-0.0008	
	(0.0012)		(0.0013)	
Professional	1.2367		1.1403	
	(0.3910)***		(0.4066)***	
Managerial	0.4364		0.4069	
	(0.2343)*		(0.2403)*	
Skilled non-manual	1.3375		1.3575	
	(0.2300)***		(0.2392)***	
Skilled manual	0.3203		0.4227	
	(0.3103)		(0.3291)	
ρ			-0.0197	
			(0.3021)	
Observations	516	1230	1230	

Table 9: Estimation results of the bivariate probit and single probits for females. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Robust standard errors in parantheses.

	OLS (no correction)		Univariate Probit Correction		Bivariate Probit Correction	
	(1) Public	(2) Private	(3) Public	(4) Private	(5) Public	(6) Private
Constant	1.7151	1.9310	1.0195	2.2558	0.8102	2.1240
	(0.1542)***	(0.0765)***	(0.4024)**	(0.1795)***	(0.4590)*	(0.1372)***
Age 16-20	-0.3468	-0.5680	-0.4928	-0.5856	-0.6171	-0.6374
	(0.2632)	(0.0806)***	(0.4071)	(0.1644)*** (N) (BC)	(0.4340)	(0.1027)***
Age 21-30	-0.2201	-0.1509	-0.2073	-0.1857	-0.1964	-0.1728
	(0.1042)**	(0.0550)***	(0.1317)	(0.0645)***	(0.1300)	(0.0638)***
Married	0.1619	0.1608	0.1930	0.1235	0.2555	0.1587
	(0.0719)**	(0.0475)***	(0.1413)	(0.0712)	(0.1342)**	(0.0695)***
Member of union	0.1248	0.0596	0.2554	0.0799	0.2455	0.0757
	(0.0696)*	(0.0520)	(0.0929)***	(0.0539)	(0.0863)***	(0.0549)
Professional	0.7409	0.7687	0.7467	0.8102	0.7041	0.7835
	(0.1557)***	(0.0844)***	(0.1976)***	(0.1021)***	(0.1898)***	(0.0995)***
Managerial	0.6433	0.7633	0.7739	0.8107	0.7801	0.8096
	(0.1065)***	(0.0698)***	(0.1335)***	(0.0823)***	(0.1398)***	(0.0787)***
Skilled non-manual	0.3053	0.1833	0.4326	0.2421	0.4885	0.2686
	(0.0899)***	(0.0648)***	(0.1334)***	(0.0780)***	(0.1426)***	(0.0769)***
Skilled manual	0.0521	0.1989	-0.1348	0.1340	-0.1057	0.1472
	(0.1016)	(0.0614)***	(0.1425)	(0.0737)**	(0.1384)	(0.0713)**
Higher degree	-0.0254	0.2721	-0.0456	0.2250	0.0029	0.2557
	(0.2223)	(0.0981)***	(0.2935)	(0.1317)**	(0.2895)	(0.1205)**
A-level	-0.0503	0.0581	-0.0201	0.0447	0.0073	0.0628
	(0.0824)	(0.0478)	(0.1089)	(0.0621)	(0.1046)	(0.0593)
O-level	-0.0999	-0.1181	-0.1561	-0.1465	-0.1649	-0.1403
	(0.0934)	(0.0540)**	(0.1241)	(0.0672)***	(0.1215)	(0.0615)***
Job tenure	-0.0057	0.0008	0.0128	0.0148	0.0074	0.0106
	(0.0140)	(0.0084)	(0.0186)	(0.0099)	(0.0189)	(0.0099)
Job tenure sq	0.0002	-0.0003	-0.0002	-0.0007	-0.0000	-0.0005
	(0.0006)	(0.0003)	(0.0007)	(0.0003)	(0.0007)	(0.0003)*
Small	0.0617	-0.3409	-0.0729	-0.4312	-0.0312	-0.4042
	(0.1043)	(0.0488)***	(0.1404)	(0.0616)***	(0.1295)	(0.0577)***
Medium	0.0474	-0.2673	-0.0101	-0.3134	0.0237	-0.2910
	(0.0931)	(0.0541)***	(0.1148)	(0.0645)***	(0.1090)	(0.0600)***
λ_p (Participation)			-0.1640	-0.2337	0.0576	-0.0790
			(0.4156)	(0.2570)	(0.8223)	(0.2356)
λ_s (Sector)			0.5175	0.4862	0.5182	0.4070
			(0.1783)***	(0.1615)***	(0.2303)**	(0.1632)***
ρ_{ju}			-0.2826	-0.384	0.1124	-0.1602
			(0.5369)	(0.3598)	(0.7867)	(0.3803)
ρ_{jv}			0.8915	0.7989	1.0124	0.8256
			(0.1864)***	(0.7989)***	(0.2875)**	(0.2735)**
σ_{jj}			0.5804	0.6084	0.5377	0.4929
			(0.1710)**	(0.1242)**	(0.4356)**(P) (BC)	(0.9817)** (P
Observations	149	528	149	528	149	528
R-squared	0.46	0.54	0.52	0.56	0.51	0.56

Table 10: Estimation results for males and three different model specifications: OLS, simple selection terms based on separate probits, and sample correction terms based on a bivariate probite. Dependent variable is wage in public and private sector. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Standard errors in parantheses, where OLS standard errors are robust and the s.e. for the remaining two models are bootstrapped. (N) refers to Normal, (P) to Percentile and (BC) to Bias corrected confidence intervals (1000 replications).

	OLS (no correction)		Univariate Probit Correction		Bivariate Probit Correction	
	(1)	(2)	(3)	(4)	(5)	(6)
	Public	Private	Public	Private	Public	Private
Constant	1.4481	1.5040	1.1882	1.5665	1.2181	1.5626
	(0.1590)***	(0.0983)***	(0.2866)***	(0.1977)***	(0.2674)***	(0.2159)***
Age 16-20	-0.5000	-0.3640	-0.6282	-0.4575	-0.5998	-0.4526
	(0.1188)***	(0.0975)***	(0.2187)***	(0.1343)***	(0.2023)***	(0.1218)***
Age 21-30	-0.2013	0.0054	-0.2689	-0.0515	-0.2615	-0.0469
	(0.0990)**	(0.0556)	(0.1460)*	(0.0874)	(0.1392)*	(0.0870)
Married	0.0201	0.0897	0.0078	0.0636	0.0031	0.0565
	(0.0507)	(0.0525)*	(0.0554)	(0.0582)	(0.0553)	(0.0608)
Member of union	0.1632	0.0011	0.1720	0.0067	0.1713	0.0063
	(0.0649)**	(0.0503)	(0.0664)***	(0.0522)	(0.0666)***	(0.0516)
Professional	0.7229	0.7788	0.8601	0.9002	0.8397	0.8831
	(0.1965)***	(0.1288)***	(0.2527)***	(0.1807)***	(0.2276)***	(0.1742)***
Managerial	0.7860	0.5297	0.9295	0.6595	0.9215	0.6530
	(0.1513)***	(0.0806)***	(0.2153)***	(0.1346)***	(0.1920)***	(0.1303)***
Skilled non-manual	0.5002	0.2504	0.5441	0.2815	0.5382	0.2776
	(0.1566)***	(0.0658)***	(0.1776)***	(0.0784)***	(0.1632)***	(0.0748)***
Skilled manual	0.2427	0.2180	0.2481	0.2353	0.2595	0.2434
	(0.1951)	(0.0802)***	(0.2180)	(0.0931)***	(0.2188)	(0.0925)***
Higher degree	0.2509	0.3211	0.2443	0.3599	0.2645	0.3893
	(0.1039)**	(0.1059)***	(0.1217)**	(0.1444)***	(0.1293)**	(0.1474)***
A-level	-0.1250	0.1104	-0.1506	0.0896	-0.1513	0.0882
	(0.0717)*	(0.0591)*	(0.0841)*	(0.0692)	(0.0831)**(N)	(0.0678)
O-level	-0.2915	-0.0045	-0.3051	-0.0243	-0.3105	-0.0306
	(0.0972)***	(0.0553)	(0.1008)***	(0.0620)	(0.1048)***	(0.0624)
Job tenure	0.0020	0.0065	0.0095	0.0154	0.0068	0.0125
	(0.0120)	(0.0122)	(0.0149)	(0.0131)	(0.0142)	(0.0132)
Job tenure sq	-0.0000	-0.0005	-0.0002	-0.0006	-0.0001	-0.0006
	(0.0005)	(0.0006)	(0.0006)	(0.0006)	(0.0006)	(0.0006)
Small	0.0308	-0.1943	-0.0310	-0.2493	-0.0237	-0.2443
	(0.0682)	(0.0674)***	(0.0944)	(0.0897)***	(0.0918)	(0.0887)***
Medium	0.0385	-0.0486	-0.0165	-0.1065	-0.0204	-0.1125
	(0.0766)	(0.0702)	(0.0971)	(0.0906)	(0.0997)	(0.0954)
λ_p (Participation)			0.0550	0.0935	0.0575	0.1121
			(0.1019)	(0.1380)	(0.1260)	(0.1547)
λ_s (Sector)			0.1916	0.2488	0.1803	0.2271
			(0.1879)	(0.2022)	(0.1673)	(0.1912)
ρ_{ju}			0.1433	0.2232	0.4950	0.3174
			(0.2527)	(0.3079)	(0.4092)	(0.3969)
ρ_{jv}			0.4992	0.5939	0.1577	0.6431
			(0.4061)	(0.3955)	(0.3057)	(0.5636)
σ_{jj}			0.3838	0.4188	0.3642	0.3531
			(0.0721)	(0.0755)	(0.0520)**	(0.0598)**
Observations	214	302	214	302	214	302
R-squared	0.47	0.38	0.48	0.39	0.48	0.39

Table 11: Estimation results for females and three different model specifications: OLS, simple selection terms based on separate probits, and sample correction terms based on a bivariate probite. Dependent variable is wage in public and private sector. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Standard errors in parantheses, where OLS standard errors are robust and the s.e. for the remaining two models are bootstrapped. (N) refers to Normal, (P) to Percentile and (BC) to Bias corrected confidence intervals (1000 replications).

	Males		Females	
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**
$\beta^*(\overline{X}^{pub} - \overline{X}^{pri})$	0.2270** (106.0 %)	0.2215** (104.0 %)	0.1892** (46.0 %)	0.3802** (92.0 %)
$\overline{X}^{pri}(\beta^* - \beta^{pri})$	0	-0.0082 (-4.0%)	0	0.0338 (8.0 %)
$\overline{X}^{pub}(\beta^{pub} - \beta^*)$	-0.0138 (-6.0 %)	0	0.2248** (54.0 %)	0

Table 12: Pay gap between public and private sector by gender and for two different weighting schemes. OLS results. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Significance levels are bootstraped, where (N) refers to Normal, (P) to Percentile and (BC) to Bias corrected confidence intervals (1300 replications).

	Males		Females	
Predicted log wages	$\ln \overline{w}_{pub} = 2.2392$	$\ln \overline{w}_{priv} = 2.0263$	$\ln \overline{w}_{pub} = 2.1516$	$\ln \overline{w}_{priv} = 1.7369$
$\ln \overline{w}_{pub} - \ln \overline{w}_{pri}$		0.2130**		0.4148**
$\beta^*(\overline{X}^{pub} - \overline{X}^{pri})$	0.2957** (138.0 %)	0.4131** (194.0 %)	0.2759** (66.0 %)	0.4729** (114.0 %)
$\overline{X}^{pri}(\beta^* - \beta^{pri})$	0	-0.8973** (N) (BC) (-424.0 %)	0	-0.2994 (-72.0 %)
$\overline{X}^{pub}(\beta^{pub} - \beta^*)$	-0.7798** (N) (BC) (-365.0 %)	0	-0.1023 (-24.0 %)	0
Selection (due to differences in the selection terms)		0.6972** (N) (327.0 %)		0.2412 (58.0 %)
Sector choice		0.7366**		0.2622** (BC)
Participation		-0.0393		-0.0209
Selectivity correct wage gap		-0.4841		0.1735

Table 13: Pay gap between public and private sector by gender and for two different weighting schemes. Simple probit model. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Significance levels are bootstraped, where (N) refers to Normal, (P) to Percentile and (BC) to Bias corrected confidence intervals (1300 replications).

	Males		Females	
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**
$\beta^*(\overline{X}^{pub} - \overline{X}^{pri})$	0.2913** (136.0 %)	0.4194** (196.0 %)	0.2686** (65.0 %)	0.4629** (112.0 %)
$\overline{X}^{pri}(\beta^* - \beta^{pri})$	0	-1.033** (-482.0 %)	0	-0.2501 (-60.0 %)
$\overline{X}^{pub}(\beta^{pub} - \beta^*)$	-0.9103** (-424.0 %)	0	-0.0559 (-13.0 %)	0
Selection (due to differences in the selection terms)		0.8332** (388.0 %)		0.2025 (48.0 %)
Selection choice		0.7740**		0.2466** (BC)
Participation		0.0599		-0.0441
Selectivity correct wage gap		-0.6189** (BC)		0.2127

Table 14: Pay gap between public and private sector by gender and for two different weighting schemes. Bivariate probit model. Cross-section weights applied. *significant at 10 per cent **significant at 5 per cent ***significant at 1 per cent. Significance levels are bootstraped, where (N) refers to Normal, (P) to Percentile and (BC) to Bias corrected confidence intervals (1300 replications).

	Males		Females	
	Marginal Effects			
	Reduced Form	Structural Form	Reduced Form	Structural Form
Constant				
Age 16-20	-0.1494	-0.2160	-0.2455	0.2285
	(0.0320)***	(0.0188)***	(0.0840)***	(0.1491)
Age 21-30	-0.0351	0.0090	-0.2666	0.3544
	(0.0378)	(0.0428)	(0.0485)***	(0.0938)***
Married	0.0363	0.0105	-0.0537	0.1670
	(0.0336)	(0.0355)	(0.0513)	(0.0586)***
Higher degree	0.0285	0.5945	-0.0821	0.2041
	(0.0690)	(0.1007)***	(0.1108)	(0.1363)
A-level	0.0125	0.1121	-0.0954	0.5642
	(0.0363)	(0.0465)**	(0.0580)	(0.0748)***
O-level	-0.0211	-0.0376	-0.0788	0.6834
	(0.0376)	(0.0374)	(0.0572)	(0.0658)***
Small	-0.1247	-0.3884	-0.2302	-0.4314
	(0.0345)***	(0.0455)***	(0.0599)***	(0.0620)***
Medium	-0.0561	-0.5554	-0.1969	-0.7104
	(0.0352)	(0.0538)***	(0.0611)***	(0.0557)***
Opinion on union (very important)	0.2531	0.3800	0.2239	0.0474
	(0.0694)***	(0.0747)***	(0.0906)**	(0.1010)
Opinion on union (important)	0.2278	0.3814	0.2862	0.1711
	(0.0461)***	(0.0523)***	(0.0690)***	(0.0790)**
Opinion on union (netural)	-0.0126	-0.0736	0.1772	0.1033
	(0.0558)	(0.0463)	(0.0909)*	(0.0961)
Job tenure	0.0144	0.0268	0.0290	0.0251
	(0.0062)**	(0.0068)***	(0.0110)***	(0.0120)**
Job tenure sq	-0.0004	-0.0012	-0.0006	-0.0015
	(0.0002)	(0.0003)***	(0.0005)	(0.0006)***
Professional	0.0569	0.2294	0.4469	0.4966
	(0.0761)	(0.1049)**	(0.1056)***	(0.0962)***
Managerial	0.1340	0.3119	0.4890	-0.3040
	(0.0562)**	(0.0686)***	(0.0735)***	(0.1196)**
Skilled non-manual	0.1413	-0.0221	0.1688	-0.4961
	(0.0646)**	(0.0525)	(0.0906)*	(0.0832)***
Skilled manual	-0.0962	0.1106	0.1257	-0.0055
	(0.0415)**	(0.0611)*	(0.1232)	(0.1332)
Wage differential		1.3218		2.9644
		(0.1586)***		(0.3366)***
Observations	677		516	

Table 15: Marginal effects of the reduced form and switching regression for males and females. The male wage difference is based on the univariate probit correction specification with sample selection for sector choice only and the female one on simple OLS.

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