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PRIVATE- AND PUBLIC-SECTOR EARNINGS STRUCTURES IN FINLAND*

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ABSTRACT: The present paper focuses on displaying similarities and dissimilarities in the earnings structures of the private and public sectors in Finland using Labour Force Survey data for 1987. Special attention is thereby paid to the returns on investment in human capital received by male and female employees in the two sectors, with a further distinction made between males and females employed in, respectively, non-manual and manual jobs. The earnings equations are estimated using general selection techniques in order to account for potential selection bias arising from the individuals' decisions on labour force participation and their choice of sectoral and occupational status.
The empirical evidence suggests that the estimation of separate earnings equations for the private and the public sector may produce quite a puzzling picture of the effects of human capital on the earnings structure of the two sectors. This picture changes, however, dramatically when a further distinction is made between employees in non-manual and manual jobs in each sector. In particular, the estimation results display a high degree of similarity across sectoral returns on investment in human capital within the broad categories of non-manual and manual workers, implying that the differences in human capital returns tend to be larger between occupational categories than across sectors.

KEY WORDS: earnings differentials, human capital, private/public sector

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TIIVISTELMÄ: Tutkimuksessa tarkastellaan julkisen ja yksityisen sektorin palkkarakennetta Suomessa käyttämällä Tilastokeskuksen työvoimatiedustelua vuodelta 1987. Erityistä huomiota kiinnitetään mies- ja naispalkansaajien koulutuksesta sekä muusta inhimillisestä pääomasta saamiin tuottoihin ja niissä esiintyviin eroihin. Tarkastelun pääkohteena on missä määrin koulutuksen, työkokemuksen yms. vaikutus palkkoihin vaihtelee toisaalta sektoreiden välillä ja toisaalta sektoreiden sisällä sukupuolten ja eri työntekijäryhmien välillä. Tätä tutkitaan estimoimalla inhimillisen pääoman teoriaan perustuvia laajennettuja palkkayhtälöitä yleistettyjä valikoituvuusmenetelmiä käyttäen. Näillä menetelmillä pyritään korjaamaan estimointituloksia mahdollisesta valikoituvuusharhasta, joka voi syntyä yksilöiden päätöksistä osallistua tai olla osallistumatta työelämään sekä heidän päätöksistään hakeutua tietyn sektorin palvelukseen ja tiettyyn ammattiryhmään.
Inhimilliseen pääomaan ja etenkin työkokemukseen liittyvien palkkavaikutusten osalta palkkayhtälöiden estimointi erikseen julkiselle ja yksityiselle sektorille tuottaa varsin yllättäviä ja osin myös ennakko-odotuksista poikkeavia tutkimustuloksia. Kuva kuitenkin muuttuu merkittävästi, kun otetaan huomioon sektoreiden erilaiset ammattirakenteet jakamalla mies- ja naispalkansaajat edelleen toimihenkilöihin ja työntekijöihin. Näille palkansaajaryhmille estimoidut palkkayhtälöt osoittavat, että inhimillisen pääoman vaikutus toimihenkilöiden ja työntekijöiden palkanmuodostukseen vaihtelee varsin vähän sektoreittain. Tästä voi päätellä, että inhimillisen pääoman palkkavaikutuksissa esiintyvät erot ovat huomattavasti suuremmat ammattiryhmien välillä kuin julkisen ja yksityisen sektorin välillä.

AVAINSANOJA: inhimillinen pääoma, palkkaeroja, julkinen/yksityinen sektori

1. INTRODUCTION

In the past few decades, the public sector has experienced a rapid increase both in employment and in relative earnings (e.g. Oxley & Martin, 1991). Simultaneously the traditional differences between the private- and public-sector labour markets have become more vague or vanished. In view of this it is hardly surprising that earnings differentials between the private and public sectors have received much attention in the international literature.

The original stimulus for this development was provided by Smith (1976,1977) who analysed the wage setting processes in the government and private sectors in the United States by estimating separate earnings equations for the two sectors from a large micro data set. The methodology introduced by Smith has been frequently repeated by other authors. However, it has also been criticized for producing biased estimates since it assumes a random distribution of individuals between sectors (Belman & Heywood, 1989).

In more recent years, empirical evidence on private/public-sector wage structures and differentials has been published also for the Nordic countries: for Denmark by Pedersen et al. (1990), for Norway by Barth & Mastekaasa (1990), and for Sweden by Johansson & Selén (1988), Zetterberg (1988), Arai (1991), and Kazamaki & D'Agostino (1992). The estimation results reported in these studies are, however, hard to generalize because of notable differences in underlying data, included variables and estimation methods used. Moreover, some studies control for the sectoral status of employees by adding indicator variables to a single earnings equation, while other studies estimate separate earnings equations for the private and the public sector, occasionally with a further distinction made in each sector between male and female employees or white- and blue-collar workers.

The empirical evidence available for Finland is restricted to a few estimates on private/public-sector earnings differentials as measured by sector controls introduced in a single earnings equation. Estimation results reported by Brunila (1990) suggest that in both 1975 and 1985, the average earnings level of upper-level non-manuals in the public sector was lower for men and higher for women as compared to the private sector. Among lower-level non-manuals, on the other hand, the average earnings level of both genders was found to be higher in the public sector in 1975 but lower in 1985. Estimates reported by Eriksson (1992), also based on Population Census data, point to a significantly positive wage premium on the part of both local and central government employees in the early 1970s. This wage differential had turned negative by 1985, but significantly so for the local government level only. Estimates based on Labour Force Survey data for 1987 point to a significant negative public-sector wage differential for men and to a negligible wage differential across sectors among women (Asplund et al., 1993).

A negative wage differential for public-sector employment has generally been obtained also for the other Nordic countries. A common interpretation of this negative public-sector differential is that it is the price public-sector employees pay for better job security and/or other non-pecuniary rewards. In the United States, in turn, the general finding is a positive public-sector wage differential (e.g. Ehrenberg & Schwarz (1986), Moulton (1990)), which is found to increase further if the wage rates are adjusted for fringe benefits (Bellante & Long, 1981). This positive wage gap is commonly interpreted as an economic rent.

However, controlling for the sectoral status of employees by introducing indicator variables into a single earnings equation can be criticized for restricting the earnings effects of various background factors to be the same across sectors. Since the private and public sectors act in different market environments, the wage determination process may be expected to differ between the two sectors. Therefore, in the subsequent empirical analysis of earnings determination in the private and public sectors in Finland, separate earnings equations of the Mincer (1974) type are estimated for each sector and gender. In doing this, the individuals are assumed to exercise some choice over their sectoral status; that is, the observed allocation is thought to be the outcome of a non-random distribution of individuals on sectors, reflecting different preferences over, inter alia, working conditions.

The Swedish labour market has been found to show tendencies of producing higher mean wages in the private sector for high-paid groups and higher mean wages in the public sector for a number of low-paid groups (e.g. Wadensjö (1986), Johansson & Selén (1988), and Zetterberg (1988)). Hence, of perhaps even more interest than the overall sectoral earnings structures are the differences between specific categories of private- and public-sector employees.

In a second step, therefore, both the private/public-sector status and the occupational social status of the employees are treated as endogenously determined through a selection process. More formally, this multiple choice is dealt with in the estimations by using a general selection model proposed by Lee (1983) which, in the present analysis, allows the employees to select across four labour markets: private- and public-sector non-manual/manual jobs. This model has previously been applied by Trost & Lee (1984) to the estimation of returns to technical schooling, by Reilly (1991) and myself (Asplund, 1992a) to the estimation of occupational earnings equations, and by Gyourko & Tracy (1988) to the analysis of private- and public-sector union/non-union wage structures.

The rest of the paper is organized as follows. Section 2 presents the empirical models, the estimation methods employed and the data underlying the estimation results reported in the next two sections. Section 3 focuses on the role of primarily acquired human capital but also of other relevant personal and job-related characteristics in the determination of male and female earnings in the private and the public sector. In Section 4, the empirical analysis is extended to the gender-specific earnings structures of non-manual and manual workers employed in the two sectors. In particular, separate earnings equations are estimated by gender and sector for employees in, respectively, non-manual and manual jobs. Section 5 summarizes the reported empirical evidence with reference to evidence obtained for the other Nordic countries.

2. MODEL SPECIFICATIONS AND DATA

Separate private-sector and public-sector earnings equations of the Mincer (1974) type are estimated for each gender.[1] Specifically, the log earnings of the i^{th} male/female employed in respective sector j (j=1,2) are explained in terms of a broad set of personal and job-related characteristics. The gender-specific earnings equations for the private sector ($\ln EARN_{i1}^{g}$) and the public sector ($\ln EARN_{i2}^{g}$) may be written in the general form

$$(1)\quad \ln EARN_{i1}^{g} = X_{i1}^{g}\alpha_{1}^{g} + \varepsilon_{i1}^{g}, \qquad \varepsilon_{ij}^{g} \sim N(0,\sigma_{j}^{2})$$

$$(2)\quad \ln EARN_{i2}^{g} = X_{i2}^{g}\alpha_{2}^{g} + \varepsilon_{i2}^{g}, \qquad i = 1,\ldots,N$$

where the X_{ij}s are vectors of explanatory variables, the α_js are vectors of parameters to be estimated, and the ϵ_{ij}s are disturbance terms. For convenience, the superscript g denoting gender is suppressed in the following.

Estimation of the sector-specific earnings equations in (1) and (2) using ordinary least squares (OLS) techniques may involve problems of sample selectivity bias and endogeneity of explanatory variables. First, in the survey data used in the analysis, the sample individuals recorded as being in employment represent persons who were employed during the week of the questionnaire, excluding all individuals who, for some reason, were not in employment at that particular time. Second, the allocation of employees into the private and public sectors may not be the outcome of a random drawing, allowing sector employment to be treated as exogenously given. Instead it can be expected to be the outcome of individual choice over employment in the two sectors. Given that these potential sources of selection bias have a non-negligible influence on the estimation results, OLS-estimation of the sectoral earnings equations will result in inconsistent parameter estimates.

Adjustment for potential selection bias influencing the estimation results is done by estimating the earnings functions in (1) and (2) in combination with a sequential selection model of the bivariate probit type explaining the probability of the i^{th} sample individual being employed and, moreover, in the given sector. In other words, there are two criterion functions: the selection of being in employment (W^*_{i1}), and the selection of private versus public status (W^*_{i2}). These two criteria for selectivity may be written

$$(3)\quad W^*_{i1} = Y_{i1}\beta_1 + \mu_{i1}, \qquad \mu_{i1}, \mu_{i2} \sim N(0,1)$$

$$(4)\quad W^*_{i2} = Y_{i2}\beta_2 + \mu_{i2}, \qquad Cov(\mu_{i1}, \mu_{i2}) = \rho_{\mu_1\mu_2}$$

where the Y_{ik}s are vectors of explanatory variables, the β_ks are vectors of unknown parameters, and the μ_{ik}s are disturbance terms with a bivariate standard normal distribution and correlation ρ. Hence, no restrictions are imposed a priori on the independence or dependence of the two decisions.

The dependent variables (W^*_{ik}) in the bivariate probit model are unobservable, but both have a dichotomous observable realization W_{i1} (employed or not) and W_{i2} (employment in given sector) which is related to, respectively, W^*_{i1} and W^*_{i2} as follows:

$$W_{i1} = 1 \quad iff \quad W^*_{i1} > 0, \qquad W_{i1} = 0 \quad otherwise$$

$$W_{i2} = 1 \quad iff \quad W^*_{i2} > 0, \qquad W_{i2} = 0 \quad otherwise$$

Data on W_{i2} are, however, not observed unless $W_{i1} = 1$; that is, employment in the private or the public sector is observed only for the subset of working individuals, implying that the data on W_{i2} are nonrandomly selected from the entire sample population. Further, the private-sector earnings equation in (1) is observed only if $W_{i1} = 1$ and $W_{i2} = 1$, while the public-sector earnings equation in (2) is observed only if $W_{i1} = 1$ and $W_{i2} = 0$. The two sets with $W_{i1} = 0$ will logically be empty.

The information obtained from estimating the bivariate sequential-decision model in (3) and (4), i.e.

$$(5)\ Prob(W_{i1}=1, W_{i2}=1) = Prob(\mu_{i1} < Y_{i1}\beta_1, \mu_{i2} < Y_{i2}\beta_2) = F(Y_{i1}\beta_1, Y_{i2}\beta_2, \rho_{\mu_1\mu_2})$$

applying bivariate probit analysis is then used to correct the sector-specific earnings equations in (1) and (2) for the potential presence of selectivity bias arising from the decision whether or not to enter the labour force and, if so, whether to work in the private sector or the public sector. By allowing the two decisions to be correlated, i.e. $Cov(\mu_{i1}, \mu_{i2}) = \rho_{\mu_1\mu_2}$, the expressions for the selectivity bias correction become considerably more complicated compared to those of the standard Heckman (1979) two-stage estimation procedure, which would require the two decisions to be independent ($Cov(\mu_{i1}, \mu_{i2}) = 0$).

Following Fishe et al. (1981) and Maddala (1983), the conditional expectation of, say, the private-sector earnings equation in (1), when assuming dependence in the underlying decisions, may be written

$$(6)\ E(\ln EARN_{i1} | W_{i1}=1, W_{i2}=1) = X_{i1}\alpha_1 + E(\varepsilon_{i1} | \mu_{i1} < Y_{i1}\beta_1, \mu_{i2} < Y_{i2}\beta_2) = X_{i1}\alpha_1 + \lambda_{11}M_{12} + \lambda_{12}M_{21}$$

where ϵ_{i1}, μ_{i1} and μ_{i2} are assumed to follow a trivariate normal distribution and where

$$(7)\ \lambda_{jk} = Cov(\varepsilon_j, \mu_k), \qquad j = 1,2 \quad k = 1,2$$

$$(8)\ M_{jk} = (1 - \rho^2_{\mu_1\mu_2})^{-1}(P_j - \rho_{\mu_1\mu_2}P_k)$$

$$(9)\ P_k = \frac{\int_{-\infty}^{Y_{i1}\beta_1}\int_{-\infty}^{Y_{i2}\beta_2} \mu_{ik} f(\mu_{i1}, \mu_{i2})\, d\mu_{i2}\, d\mu_{i1}}{F(Y_{i1}\beta_1, Y_{i2}\beta_2)}$$

After having used bivariate probit methods to estimate β_1, β_2, and $\rho_{\mu_1\mu_2}$, the second stage of the estimation procedure thus involves regression of individual private-sector earnings ($\ln EARN_{i1}$) on X_{i1} and the constructed variables M_{12} and M_{21} in order to obtain consistent estimates of α_1, λ_{11}, and λ_{12}. The public-sector earnings equation in (2) is corrected for potential selection bias in an analogous way.

Various empirical specifications of the sectoral earnings equations in (1) and (2) are estimated within the LIMDEP framework, whereby correction for the potential presence of the two sources of selectivity bias is done using the bivariate probit sample selection procedure outlined above. More formally, the applied estimation method allows the two decisions underlying employment in a given sector to be correlated, and accounts for sample selection both in the bivariate probit model and in the earnings model.

However, because of the different occupational structure of the private and public sectors there is reason to expect some degree of selectivity also when it comes to the occupational status of the employee. Therefore, in the following model specification both the sectoral status and the occupational status of employees are treated as endogenous variables. A distinction is made between four labour markets: private-sector non-manuals, private-sector manuals, public-sector non-manuals, and public-sector manuals. The potential sample selectivity bias arising from the decision whether or not to join the labour force is not accounted for in this context.[2]

Following Lee (1983), the adopted approach, involving censored dependent variables in combination with multiple choice, can be formulated in terms of a polychotomous choice model with four mutually exclusive labour markets (LM_{im}) and four earnings equations ($EARN_{im}$):

$$(10)\quad \ln EARN_{im} = Z_{im}\gamma_m + \zeta_{im}, \qquad \zeta_{im} \sim N(0,\sigma_m^2)$$

$$(11)\quad LM_{im}^{*} = V_i\theta_m + \eta_{im}, \qquad i = 1,\ldots,N \qquad m = 1,2,3,4$$

where Z_{im} and V_i are vectors of explanatory variables, γ_m and θ_m are vectors of unknown parameters, and ζ_{im} and η_{im} are disturbance terms. The potential earnings of the i^{th} sample employee in the m^{th} labour market given by the earnings function in (10) will be affected by selectivity bias if the disturbances in (10) and (11) are correlated.

The dependent variable ($\ln EARN_{im}$) in the labour market earnings equation is observed only if the employee chooses to work in labour market m. This choice is assumed to be the outcome of an optimization process where the individual compares the maximum lifetime expected utility attainable from participating in respective labour market and selects that alternative which provides the highest present value of net benefits. The utility maximization process is thought to be captured by the labour market indicator function

$$(12)\quad LM_i = m \qquad iff \quad LM^*_{im} > \max LM^*_{ik}\,, \qquad k = 1,2,3,4 \quad k \neq m$$

Following Lee (1983), the i^{th} individual's choice of the m^{th} labour market as expressed in (12) can be reformulated as a binary decision, i.e.

$$(13)\quad LM_i = m \qquad iff \quad V_i\theta_m > \psi_{im}$$

where ψ_{im} is the residual for each individual and labour market and is defined as

$$(14)\quad \psi_{im} = \max LM^*_{ik} - \eta_{im}\,, \qquad k = 1,2,3,4 \quad k \neq m$$

Assuming that the residuals (η_{im}) of the utility function in (11) are independently and identically distributed with the Type 1 extreme value distribution[3], the probability that the labour market m will be chosen can be represented by a multinomial logit model[4]

$$(15)\quad Prob(\psi_{im} < V_i\theta_m) = Prob(LM_i = m) = \frac{\exp(V_i\theta_m)}{1 + \sum_{k=1}^{4}\exp(V_i\theta_k)}$$

Only the parameters of three of the four labour markets investigated can be identified, which requires a normalization $\Sigma\theta_m = 0$ (m = 1,2,3,4) to be imposed in the estimations.[5] The earnings equation conditional on labour market m being chosen may then be written

$$(16)\quad E(\ln EARN_{im}|LM_i = m) = Z_{im}\gamma_m + E(\zeta_{im}|LM_i = m)$$
$$= Z_{im}\gamma_m + E(\zeta_{im}|\psi_{im} < V_i\theta_m)$$

Given that ζ_{im} and ψ_{im} follow a bivariate normal distribution, a two-step estimation procedure similar to that postulated by Heckman (1979) can be used in order to correct the labour market earnings function in (10) for the potential effects of selectivity bias arising from the employees' choice of labour markets. Following Lee (1983), the ψ_{im}s are transformed into standard normal random variables and a modified earnings equation conditional on labour market m being chosen is derived

$$(17)\quad E(\ln EARN_{im}|LM_i = m) = Z_{im}\gamma_m - \rho_m\sigma_m\frac{\phi[F_m(V_i\theta_m)]}{\Phi[F_m(V_i\theta_m)]} = Z_{im}\gamma_m + \rho_m\sigma_m\lambda_m$$

where ρ_m is the correlation coefficient between the ζ_{im}s and the transformed ψ_{im}s, σ_m is the standard deviation of the disturbance term in the earnings equation, $\phi(\cdot)$ and $\Phi(\cdot)$ are, respectively, the density function and the distribution function of the standard normal distribution, and $F(\cdot)$ denotes the probability distribution function.

Various empirical specifications of the labour market earnings equation in (17) are estimated within the LIMDEP framework using the multinomial logit-OLS two-stage estimator of Lee (1983). More exactly, the multinomial probability function in (15) is estimated by maximum likelihood and the obtained information is used to compute λ_m, i.e. the term controlling for the potential effects of selectivity bias. Consistent estimates of γ_m and $\Omega_m=\rho_m\sigma_m$ are then obtained by ordinary least squares regression of $\ln EARN_{im}$ on Z_{im} and $\hat{\lambda}_m$:

$$(18)\ \ln EARN_{im} = Z_{im}\gamma_m + \Omega_m\hat{\lambda}_m + \tau_{im}$$

where $E(\tau_{im}|LM_i=m) = 0$. The standard errors are corrected using the heteroscedasticity-consistent estimator suggested by White (1980).

The three criterion functions appearing in the two selection models outlined above are specified as follows. The selectivity criterion in (3) explaining the probability of the i^{th} sample individual being employed includes a set of personal characteristics containing age and indicators for educational level, marital status, family size, and location of residence.

The probability of private/public-sector employment (eq. (4)), in turn, is taken to depend on the individual's accumulated human capital, marital status, preferences over job characteristics, and on variations across regional labour markets. The allocation of employees in the four labour markets of private/public-sector non-manual/manual workers (eq. (15)) is assumed to depend on the same broad set of characteristics. The current age of the individual is not included as an explanatory variable in the sectoral and labour market criterion functions (eqs. (4) and (15)), the underlying assumption being that there is no systematic movement of employees between labour markets as they grow older (cf. Gyourko & Tracy, 1988).

Finally, the observed earnings variance among male and female employees in the private and public sectors is assumed to be dependent on the employees' formal education, labour market experience and training, family responsibilities, location of residence, employment and working conditions, union membership, and industry affiliation. Apart from these explanatory variables, the sector-specific earnings models in (1) and (2) are also supplemented with a set of occupation indicators in order to examine whether the interaction effects of the individual's position in the occupational hierarchy differ between the two sectors. The observed earnings differentials between the four labour markets (eq. (10)) are explained in terms of the same set of explanatory variables, except for occupation controls.

It should be pointed out in this context that at least in the United States, the inclusion of occupation controls in earnings equations and especially in sectoral earnings equations has generated some controversy. In particular, it has occasionally been claimed that occupation indicators should be excluded from the regressions as certain jobs may be classified differently in the private and public sectors (cf. Belman & Heywood, 1989). It is argued that the addition of occupation controls will, as a consequence, distort the comparability of estimates across sectors. Estimation of sectoral earnings functions both inclusive and exclusive of occupation controls may therefore shed some further light on this topic.

The earnings models outlined above are estimated using cross-sectional micro data from the Labour Force Survey for 1987 conducted by Statistics Finland. The survey covers a random sample of some 9000 persons, representing the entire population aged 15-64 years as stratified according to sex, age and region. When the data are restricted to employed wage and salary earners at the age 16 to 64 and sorted out with respect to missing and incomplete information on crucial variables, the sample of employees retained in the actual estimating data shrinks to covering a total of 3895 individuals.

The dependent variable is chosen to be average before-tax hourly earnings in order to account for interpersonal differences in months and weekly hours worked, and to make the earnings of full-time and part-time employees comparable. The earnings data used comprise most types of compensation, including overtime and vacation pay and fringe benefits[6]. The available register data on formal schooling merely show the highest single education completed by each sample individual, not their actual schooling years. There is a total of eight levels of education, which are represented in the estimations by both linear and non-linear schooling variables. A notable advantage of the employed data set

Table 1. Sample statistics for selected variables (standard deviations are in parentheses)

Variable	MALE EMPLOYEES Private sector	Public sector	FEMALE EMPLOYEES Private sector	Public sector
Hourly earnings	49.12 (23.57)	50.84 (25.59)	39.65 (24.31)	42.86 (22.71)
Log hourly earnings	3.81 (0.38)	3.85 (0.37)	3.60 (0.36)	3.68 (0.37)
Hourly earnings$^{\text{no fringes}}$	48.43 (22.64)	50.67 (25.31)	39.27 (23.52)	42.69 (22.65)
Log hourly earnings$^{\text{no fringes}}$	3.80 (0.37)	3.85 (0.37)	3.59 (0.36)	3.67 (0.37)
Schooling (S+9 years)	10.77 (1.76)	11.70 (2.39)	10.54 (1.65)	11.55 (2.11)
Experience	17.15 (10.98)	18.34 (10.39)	16.27 (9.95)	15.98 (9.65)
Seniority (tenure)	8.58 (8.36)	11.23 (8.93)	8.42 (8.04)	8.82 (8.14)
Share of trained employees	0.33	0.46	0.32	0.46
Number of obs.	1423	485	1099	888
Share in				
- sample,%	36.5	12.5	28.2	22.8
- whole economy,%	40.7	10.5	30.6	18.3
Share of females in resp. sector				
- sample,%			43.6	64.7
- whole economy,%			42.9	63.6

Source: Labour Force Survey for 1987 and Table B of Appendix.

is that it provides (self-reported) information on each person's total years of labour market experience as well as on his or her years with the current employer, i.e. seniority (tenure). Hence, the estimation results reflect the earnings effects of the individuals' "actual" and not of their potential work experience.

A summary of definitions of the variables employed in the subsequent empirical analysis is given in Table A of Appendix. Sample statistics for selected variables are shown in Table 1 below. A complete list of sample means for all males and females in private/public-sector employment and separately for the four labour markets considered is found in Tables B and D of Appendix. A detailed presentation of the underlying data and definitions of crucial variables are given in Asplund (1992b).

3. PRIVATE/PUBLIC-SECTOR EARNINGS STRUCTURES

3.1. Sectoral earnings effects of education and experience

The regression results obtained from estimating private/public-sector human capital earnings functions for each gender are displayed in Tables 2 and 3. The bivariate probit estimates are reported in Table E of Appendix.

Before turning to the estimated earnings effects of the various explanatory variables included in the estimations, three features of the results deserve attention. First, the estimated earnings functions succeed in explaining substantially more of the observed earnings dispersion in the public sector. Second, the standardized hourly earnings differential between central and local government employees (PUBLOCAL) is found to be negligible among both males and females. Finally, the exclusion of fringe benefits from the dependent variable leaves the parameter estimates of the various explanatory variables roughly unchanged.[7]

For both genders, the average level of schooling is higher among public-sector employees (Table 1). If education is equally rewarded in the two sectors, then the returns to schooling would be expected to be higher in the public sector. However, the schooling coefficients estimated for males suggest that there are no notable differences[8] in educational returns between private-sector and public-sector male employees. In both sectors, the average return to an additional year in above-primary schooling amounts to some 9 per cent (estimation results not shown). Also the returns to different educational degrees are very similar in the two sectors, ranging from some 11 per cent for graduation at the LOWER VOCATIONAL and professional level to over 90 per cent for a university GRADUATE degree when compared to the average return received by persons with basic education only (columns 1 and 3 in Table 2).[9]

Among female employees, on the other hand, the average return to an additional year in postcompulsory schooling seems to be substantially higher for females in public-sector employment (about 10 per cent compared with some 5½ per cent for private-

Table 2. Private/public-sector estimates for male employees obtained from eq. (6).[1] The dependent variable is log hourly earnings inclusive of fringe benefits. (Occupation controls are included in columns 2 and 4.)

Variable	Private sector males (1)	(2)	Public sector males (3)	(4)
CONSTANT	3.3555** (.0738)	3.5185** (.0997)	3.2312** (.1032)	3.3986** (.1152)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1057** (.0221)	0.0794** (.0213)	0.1055** (.0340)	0.0792** (.0326)
UPPER VOCATIONAL	0.2928** (.0277)	0.1487** (.0298)	0.2605** (.0390)	0.1733** (.0419)
SHORT NON-UNIV	0.4684** (.0521)	0.2032** (.0572)	0.4955** (.0516)	0.2514** (.0700)
UNDER-GRADUATE	0.5336** (.1160)	0.2433* (.1143)	0.4904** (.0650)	0.2280** (.0828)
GRADUATE	0.6622** (.0522)	0.3754** (.0570)	0.6460** (.0488)	0.3820** (.0691)
EXP	0.0178** (.0047)	0.0100* (.0045)	0.0280** (.0061)	0.0248** (.0063)
EXP^2 /1000	-0.2461* (.1198)	-0.1037 (.1146)	-0.4895** (.1563)	-0.4496** (.1604)
MARRIED	0.0557* (.0281)	0.0317 (.0266)	0.0884* (.0395)	0.0729* (.0386)
$CHILD^{0-6}$	-0.0034 (.0221)	-0.0042 (.0209)	-0.0532* (.0303)	-0.0382 (.0278)
$CHILD^{7-17}$	0.0544** (.0212)	0.0439* (.0201)	0.0536* (.0285)	0.0469* (.0271)
CAPITAL	0.1676** (.0233)	0.1484** (.0222)	0.0686* (.0314)	0.0755** (.0295)
TEMPEMPL	-0.0594 (.0387)	-0.0549 (.0367)	0.0016 (.0447)	-0.0185 (.0432)
PART-TIME	0.1219 (.0924)	0.0988 (.0879)	0.3146** (.0946)	0.3071** (.0994)
PUBLOCAL			-0.0008 (.0253)	-0.0019 (.0244)
PIECE-RATE	0.0481* (.0244)	0.0660** (.0234)	0.1695* (.0851)	0.1780** (.0757)
NODAYWORK	0.0493* (.0215)	0.0880** (.0207)	0.0541* (.0292)	0.0501* (.0282)
UNEMPL	-0.0754** (.0294)	-0.0602* (.0277)	-0.1682** (.0527)	-0.1410** (.0513)
UNION	-0.0023 (.0198)	0.0204 (.0188)	0.0045 (.0367)	0.0424 (.0351)

Table 2. (cont.)

Variable	Private sector males (1)	(2)	Public sector males (3)	(4)
Occupational status indicators:				
OCC31 (management)		0.4038** (.0769)		0.6128** (.1082)
OCC32 (research)		0.2113** (.0804)		0.1962** (.0704)
OCC33 (education)		0.1385 (.1650)		0.2076** (.0662)
OCC34 (oth.seniors		0.1917** (.0810)		0.1305* (.0621)
OCC41 (supervisors)		0.0931 (.0716)		0.0346 (.0493)
OCC42 (indep. clericals)		-0.0403 (.0802)		-0.1371 (.0986)
OCC43 (rout. clericals)		0.1098 (.1146)		-0.1983* (.1083)
OCC44 (oth. lower-level non-manuals)		0		0
OCC51 (workers, agriculture)		-0.0862 (.0983)		-0.1148 (.0911)
OCC52 (workers, manufacturing)		-0.0266 (.0757)		-0.0421 (.0747)
OCC53 (workers, oth. prod.)		-0.0864 (.0745)		-0.0585 (.0609)
OCC54 (workers, service)		-0.1413* (.0714)		-0.0998* (.0442)
LAMBDA1(ϵ,μ_1) (working sel.)	-0.0120 (.0585)	-0.0510 (.0563)	0.0835 (.0672)	0.0723 (.0729)
LAMBDA2(ϵ,μ_2) (sector sel.)	-0.0380 (.0413)	0.0108 (.0566)	0.0152 (.0332)	0.0705 (.0439)
R^2 adj.	0.3146	0.3885	0.5422	0.5934
SEE	0.3137	0.2952	0.2416	0.2249
F-value	26.11	25.42	22.23	19.59
Number of obs.	1423	1423	485	485

[1] Standard errors are given in parentheses below the estimates. Bivariate sample selection estimates where LAMBDA1(ϵ,μ_1) gives the selectivity bias associated with the individual's labour force status and LAMBDA2(ϵ,μ_2) measures the selectivity bias arising from choosing between the two sectors. The bivariate probit estimates are displayed in Table E of Appendix. The estimated earnings equations also include seven one-digit industry sector controls (INDU1,INDU2/3,INDU4,INDU5,INDU6, INDU8,INDU9), employment in transport and communication (INDU7) being the excluded variable. It may be noted that the addition of industry sector controls has no significant impact on the regression results.

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

Table 3. Private/public-sector estimates for female employees obtained from eq. (6).[1] The dependent variable is log hourly earnings inclusive of fringe benefits. (Occupation controls are included in columns 2 and 4.)

Variable	Private sector females (1)	(2)	Public sector females (3)	(4)
CONSTANT	3.4499** (.0848)	3.5568** (.0973)	3.3162** (.0831)	3.3330** (.0832)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	-0.0429 (.0265)	-0.0387 (.0255)	0.0445 (.0300)	0.0189 (.0290)
UPPER VOCATIONAL	0.1446** (.0313)	0.0974** (.0310)	0.2099** (.0326)	0.1331** (.0327)
SHORT NON-UNIV	0.3358** (.0708)	0.2506** (.0722)	0.3944** (.0404)	0.1994** (.0438)
UNDER-GRADUATE	0.4374** (.0735)	0.3557** (.0767)	0.5525** (.0513)	0.2318** (.0608)
GRADUATE	0.4622** (.0856)	0.3278** (.0879)	0.6897** (.0491)	0.3882** (.0582)
EXP	0.0098* (.0048)	0.0071 (.0047)	0.0123** (.0047)	0.0094* (.0046)
EXP^2 /1000	-0.1215 (.1195)	-0.0780 (.1171)	-0.0992 (.1217)	-0.0694 (.1185)
MARRIED	-0.0171 (.0241)	-0.0205 (.0232)	-0.0346 (.0246)	-0.0364 (.0237)
$CHILD^{0-6}$	0.0298 (.0260)	0.0163 (.0252)	0.0372 (.0254)	0.0412* (.0240)
$CHILD^{7-17}$	-0.0148 (.0222)	-0.0051 (.0216)	0.0200 (.0226)	0.0222 (.0214)
CAPITAL	0.1240** (.0250)	0.1169** (.0244)	0.0469* (.0277)	0.0672** (.0264)
TEMPEMPL	0.0638 (.0404)	0.0527 (.0396)	0.0939** (.0326)	0.0542* (.0316)
PART-TIME	0.2482** (.0402)	0.2667** (.0397)	0.3731** (.0469)	0.3557** (.0451)
PUBLOCAL			-0.0288 (.0295)	-0.0045 (.0285)
PIECE-RATE	0.0010 (.0331)	0.0449 (.0340)	-0.0937 (.1134)	-0.1404 (.1086)
NODAYWORK	0.1180** (.0236)	0.1403** (.0242)	0.1694** (.0234)	0.2194** (.0239)
UNEMPL	-0.0491 (.0338)	-0.0245 (.0330)	-0.0421 (.0382)	-0.0224 (.0370)
UNION	-0.0606** (.0240)	-0.0437* (.0239)	0.0482 (.0321)	0.0530* (.0308)

Table 3. (cont.)

Variable	Private sector females (1)	(2)	Public sector females (3)	(4)
Occupational status indicators:				
OCC31 (management)		0.2768** (.0862)		
OCC32 (research)		-0.0957 (.0980)		0.2326** (.0759)
OCC33 (education)		0.1437 (.2271)		0.4028** (.0476)
OCC34 (oth.seniors)		0.1168* (.0624)		0.1783** (.0460)
OCC41 (supervisors)		0.0005 (.0661)		0.1469** (.0527)
OCC42 (indep. clericals)		-0.0655 (.0550)		0.2038** (.0452)
OCC43 (rout. clericals)		-0.0385 (.0562)		0.0939* (.0428)
OCC44 (oth. lower-level non-manuals)		0		0
OCC51 (workers, agriculture)		-0.3134** (.1219)		0.0482 (.1543)
OCC52 (workers, manufacturing)		-0.1845** (.0642)		
OCC53 (workers, oth. prod.)		-0.1567** (.0632)		-0.1014* (.0453)
OCC54 (workers, service)		-0.1614** (.0503)		-0.0413 (.0342)
LAMBDA1(ϵ,μ_1) (working sel.)	-0.1225** (.0524)	-0.0991* (.0498)	0.0020 (.0455)	0.0369 (.0478)
LAMBDA2(ϵ,μ_2) (sector sel.)	0.0371 (.0278)	0.0098 (.0340)	-0.0163 (.0315)	0.0765* (.0371)
R^2 adj.	0.2340	0.2728	0.3620	0.4266
SEE	0.3135	0.3039	0.2907	0.2741
F-value	13.90	12.13	19.64	19.33
Number of obs.	1099	1099	888	888

[1] Standard errors are given in parentheses below the estimates. Bivariate sample selection estimates where LAMBDA1(ϵ,μ_1) gives the selectivity bias associated with the individual's labour force status and LAMBDA2(ϵ,μ_2) measures the selectivity bias arising from choosing between the two sectors. The bivariate probit estimates are displayed in Table E of Appendix. The estimated earnings equations also include seven one-digit industry sector controls (INDU1,INDU2/3,INDU4,INDU5,INDU6, INDU8,INDU9), employment in transport and communication (INDU7) being the excluded variable. It may be noted that the addition of industry sector controls has no significant impact on the regression results.

* Denotes significant estimate at a 5 % risk level.

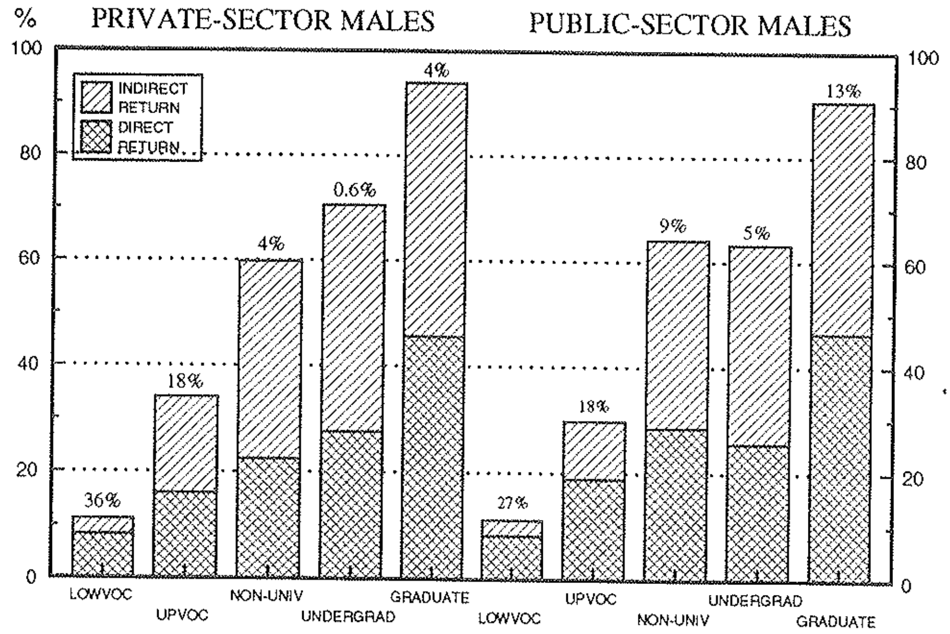
** Denotes significant estimate at a 1 % risk level.

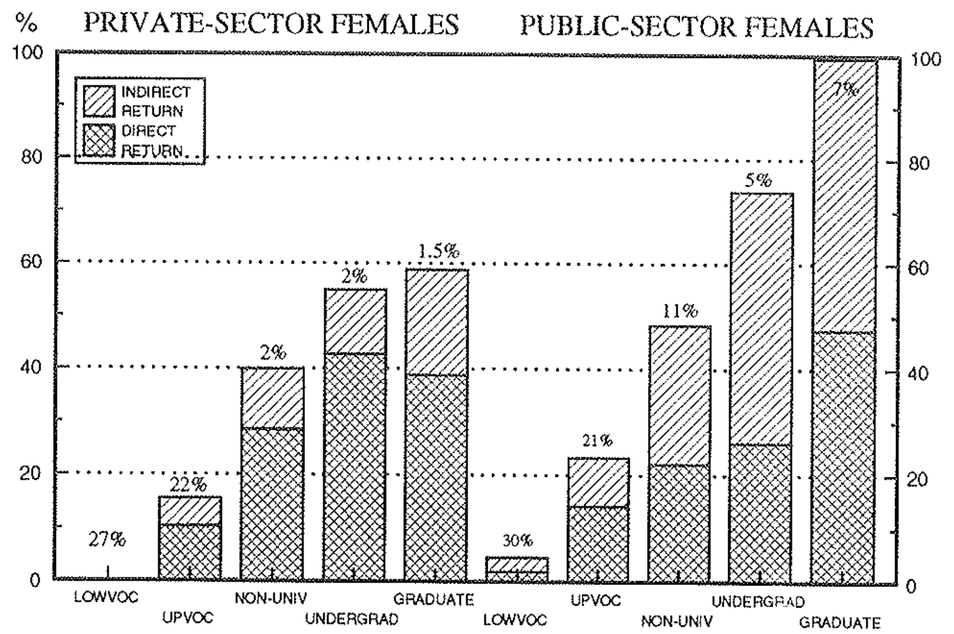
sector females). Yet, this result cannot be generalized for the estimated returns to educational degrees (columns 1 and 3 in Table 3). Irrespective of sector, women with completed LOWER VOCATIONAL and professional education tend to have no relative income advantage over women with basic education only. The estimation results further point to small, if any, differences between sectoral returns to non-university higher education degrees (SHORT NON-UNIV). Opposite to this, females with graduation from UPPER VOCATIONAL education and females with an UNDERGRADUATE or GRADUATE university degree seem to be more highly rewarded in the public sector. In fact, at these three educational levels females in public-sector employment tend to receive approximately the same return to their schooling investment as their male counterparts in the private and public sectors. But apart from this, the estimated returns to educational degrees are, on average, lower in both sectors for females.

The estimation results reported in Asplund (1992a) indicate that the earnings effects of formal schooling mediated by the individual's position in the occupational hierarchy tend to be notably larger for male than for female employees. This points to a more rigid occupational structure of male earnings. It might, as a consequence, be of interest to examine whether the interactions between education and occupational status differ, not only by gender, but also across sectors. This is done by adding a set of occupation indicators to the sectoral earnings equations (columns 2 and 4 in Tables 2 and 3).

Following this approach, the schooling coefficients obtained from estimating earnings equations exclusive of occupation controls will measure the total effect of education on earnings. When occupation indicator variables are introduced, the parameter estimates of the schooling variables will capture only the direct earnings effect of schooling. Calculations reported in Asplund (1992a) indicate that the difference between the two groups of schooling coefficients provides a rough proxy of the degree of occupational rigidity, i.e. of the indirect effect education has on earnings by influencing the individual's occupational attainment.

Figure 1. Decomposition of the estimated average return to different educational levels, by sector and gender, into a "direct" and an "indirect" (through occupational attainment) effect, the reference being the average return received by persons with basic education only.





Note. The figure on the top of each pile gives the percentage sample share of each educational level in the employee category in question (e.g. 36%+18%+4%+0.6%+4%+residual%=100%). The residual percentage share in each category refers to employees with basic education only.

Source: Antilogs of the schooling coefficients in Tables 2-3.

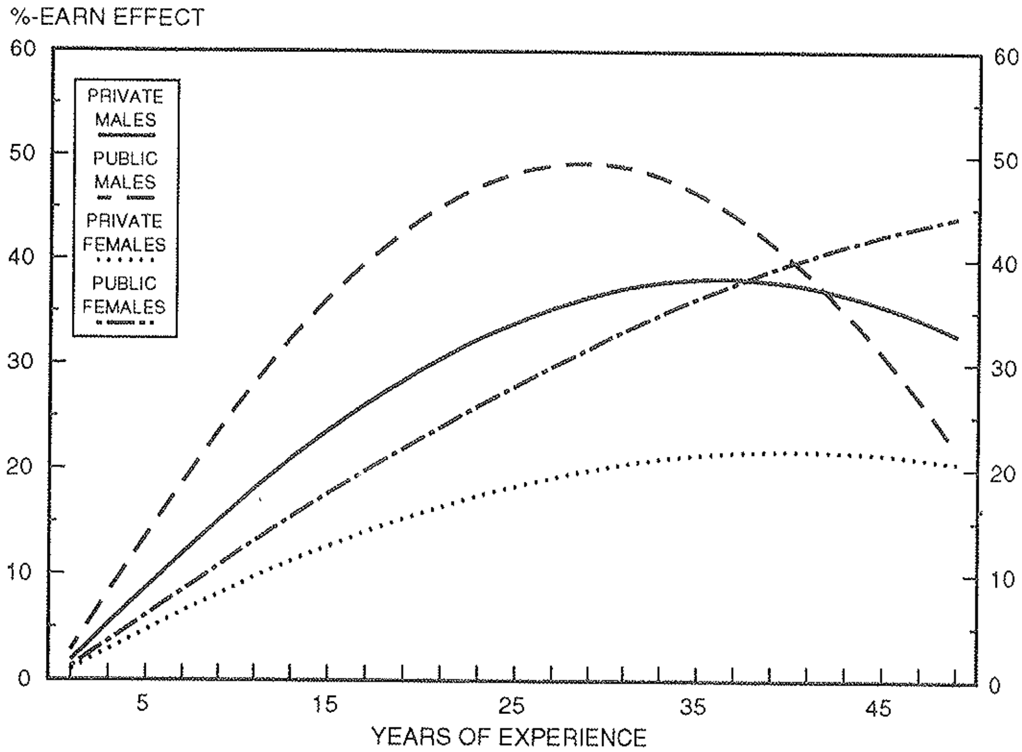
Comparison of the parameter estimates on the educational level variables in Tables 2 and 3 reveals that the addition of occupation controls to the sectoral earnings equations results in a substantial reduction in the estimated returns to formal education for male employees, the only exception being graduation from lower vocational education. Moreover, a similar pattern of reductions occurs in both sectors (Figure 1). Again the situation for females in public-sector employment resembles strongly that for male employees. Specifically, for all three employee categories a large portion of the earnings effects of education seems to arise through the influence that education has on the individual's occupational chances.

Figure 1 also shows that the inclusion of occupation controls causes an almost negligible drop in the schooling coefficients estimated for private-sector female employees. Hence, formal education tends to have a more direct effect on the earnings of females in private-sector employment, while the earnings of public-sector employees as well as of private-sector male employees seem to be more strongly influenced by occupational status. This suggests that private-sector females are likely to encounter a clearly different labour market situation when it comes to wage rigidity in general and to pecuniary returns on investment in formal education in particular.

The highly different pattern of educational returns estimated for females employed in the private sector may in part be explained by their typically lower level of formal education and the allocation on sectors and occupations that this may imply. However, controlling for the employees' occupational social status does not offer a full solution to the different labour market situation of private-sector women. Instead it seems as if the major explanations are to be looked for in differences not only across but also within occupational categories; females in private-sector employment tend to crowd into certain occupational categories and, moreover, to hold less-paid jobs within these categories (cf. Table C of Appendix). This aspect is analysed somewhat more in detail in Section 4.

The initial earnings effect of labour market experience (EXP) turns out to be highest, or almost 3 per cent, for public-sector

Figure 2. Experience-earnings profiles by sector and gender



Note. The percentage growth in earnings is calculated as the antilog of ($\alpha_1 EXP - \alpha_2 EXP^2$).

Source: Calculations based on the experience coefficients reported in Tables 2 and 3.

male employees. It is notably lower (close to 2 per cent) for males employed in the private sector. Earnings growth of females, when first entering the labour market, is of approximately the same magnitude in both sectors (about 1 per cent), and significantly lower compared with the earnings effects of experience estimated for their male counterparts.

The estimation results further point to an upward-sloping concave experience-earnings profile for male employees only (Figure 2). The magnitudes of the parameter estimates on the experience variables for men indicate that earnings growth decreases fairly slowly and reaches zero earlier for public-sector than for private-sector males. More exactly, a maximum of some 50 per cent cumulative growth[10] in public-sector male earnings is reached after 28 years of work experience, while the cumulative growth in private-sector male earnings peaks at roughly 38 per cent after no less than 36 years in the labour market. Thus the experience-earnings curves tend to be steeper and to also fall off more rapidly in the public sector.

Compared to female employees, the experience profiles of men are clearly steeper in both sectors. Indeed, the estimates point to a very moderate but fairly constant growth rate of female earnings with increased experience. Yet, also for women the experience curve turns out to be steeper in the public sector.

But in interpreting the estimated experience coefficients, it should be kept in mind that they are obtained from a single cross section of individuals, implying that no cohort effects are accounted for. Specifically, the experience-earnings profiles drawn in Figure 2 are based on the assumption that the estimated cross-sectional coefficients for experience do capture the dynamics of changes in earnings over the individual's life cycle. Moreover, even similarity in the percentage effects on earnings of more work experience and/or additional schooling may result in highly different absolute returns if the average earnings levels vary by sector and/or gender.

Although the earnings effects of the other personal and job characteristics accounted for in the estimations and displayed in Tables 2 and 3 are no central issue of the present paper, a few comments may, nonetheless, be justified. (See further the discussion in Asplund (1992c).) In both sectors, family responsibilities (MARRIED,CHILD) tend to have a positive effect on male earnings, but typically no significant influence on female earnings. Residence within the capital region (CAPITAL) is associated with a notable income advantage of both males and females in private-sector employment. The earnings differentials between public-sector employees living inside and outside the capital region are substantially lower, which is obviously attributable to the standard salary schemes and schedules applied in the public sector.

The estimated coefficients for the various job characteristics point to several interesting similarities and dissimilarities between private- and public-sector employees.[11] Presumably a major part of these findings can be given institutional explanations such as differences across the two sectors in the conditions of employment and the method of wage setting. Thus temporary employment (TEMPEMPL) seems to imply significantly higher hourly earnings as compared to a permanent job relationship for public-sector females only. A most conspicuous relative income advantage is obtained for part-time employees (PART-TIME); only private-sector males seem to encounter no significant differential between average pay in full-time and part-time employment. The results further point to clearly higher earnings of male employees in jobs covered by some other compensation system than wages/salaries paid on a monthly, weekly, or hourly basis (PIECE-RATE). For females, such extraordinary compensation systems imply a small, if any, income disadvantage. In both sectors, both genders are compensated if working in jobs entailing inconvenient working hours (NODAYWORK).

Periods of temporary unemployment or layoffs during the survey year (UNEMPL) typically implied a negative earnings effect, which turns out to have been much stronger for public- than for private-sector males. The negative coefficients on the unemployment variable obtained for females do not differ significantly from zero.[12] The influence of union membership (UNION) is found to be insignificant also when analysing earnings differentials separately for the private and the public sector (cf. Asplund (1992a,1992c)). There is one notable exception, though, namely the significant negative earnings effect of union membership obtained for females in private-sector employment. A possible explanation to this somewhat unexpected outcome is that the union variable captures some unobserved effects, such as worse working conditions, which are less strongly present in the other three employee categories. Further, the magnitudes and significance levels of the parameter estimates on the occupational indicator variables show much the same general pattern as the coefficients estimated for all male and female employees (cf. Asplund (1992a) and Figure A of Appendix).

Finally, there is some evidence of both working (LAMBDA1) and sector (LAMBDA2) selectivity bias present in the estimations for females. This suggests that estimation of sector-specific earnings equations for female employees using ordinary least squares techniques, thereby assuming a random distribution of women on labour force status groups and/or sectors, might result in inconsistent parameter estimates. For men, on the other hand, the insignificant coefficients for the selection variables indicate that consistent estimates would be obtained also when less sophisticated estimation methods are used. For comparison, regression results obtained from estimating sectoral earnings equations for each gender using (1) the standard Heckman (1979) two-stage procedure for correction of sector (but not for working) selectivity, and (2) ordinary least squares techniques are reported in Tables F and G of Appendix.

3.2. Sectoral earnings effects of seniority and OJT

The estimation results reported in Tables 2 and 3 above suggest that for both genders, starting wages/salaries are typically lower in the public than in the private sector. This is, however, compensated by a much faster growth rate of public-sector earnings with increased labour force experience. One possible interpretation of this result is that private-sector employees receive more specific training in the Becker (1962,1964) sense[13] compared with public-sector employees. An alternative explanation would be the conditions of employment and the method of earnings determination adopted in the public sector. In particular, the public sector represents a large, hierarchical internal labour market bound by fairly strict rules of employment and with the length of the employmentship playing a major part in promotion.

In order to assess this type of effects, the sector-specific earnings equations are augmented with survey information on the number of years employed at the present employer (SENiority, tenure) and on the occurrence of formal on-the-job training sponsored by the employer (OJT).[14] Tables 4 and 5 display the estimated coefficients for the included experience and training variables only. The parameter estimates on the other explanatory variables are very close to their counterparts in Tables 2-3 and are therefore not reported or commented on.

As can be seen from Table 4, the earnings effects of general experience (EXP) tend to be substantially higher for public-

Table 4. Sectoral earnings effects of experience for men obtained from estimating the human capital earnings model (eq. (6)) augmented with information on years with the current employer (SENiority) and participation in formal on-the-job training (OJT).[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR MALE EMPLOYEES (1)	(2)	(3)	(4)	PUBLIC-SECTOR MALE EMPLOYEES (1)	(2)	(3)	(4)
EXP	0.0147** (.0048)	0.0063 (.0047)	0.0135** (.0047)	0.0063 (.0046)	0.0237** (.0058)	0.0223** (.0056)	0.0231** (.0057)	0.0217** (.0055)
EXP^2 /1000	-0.2217* (.1214)	-0.0704 (.1173)	-0.1990* (.1184)	-0.0642 (.1157)	-0.4268* (.1431)	-0.4229** (.1389)	-0.4075** (.1412)	-0.4062** (.1371)
SEN	0.0050** (.0014)	0.0044** (.0013)	0.0043** (.0014)	0.0038** (.0013)	0.0036* (.0019)	0.0035* (.0018)	0.0035* (.0019)	0.0035* (.0018)
OJT			0.1296** (.0196)	0.1033** (.0190)			0.0824** (.0233)	0.0728** (.0218)
Occupation controls	no	yes	no	yes	no	yes	no	yes
R^2 adj.	0.3275	0.3988	0.3439	0.4109	0.5610	0.6202	0.5714	0.6280
SEE	0.3119	0.2938	0.3080	0.2907	0.2276	0.2091	0.2247	0.2067
F-value	26.14	25.33	27.10	25.93	22.73	20.93	22.88	21.09
Number of obs.	1395	1395	1395	1395	477	477	477	477

[1] Bivariate sample selection estimates with standard errors in parentheses below the estimates. The estimated earnings equations also include the personal and job characteristics displayed in Tables 2-3 above (LOWER VOCATIONAL, UPPER VOCATIONAL, SHORT NON-UNIV, UNDERGRADUATE, GRADUATE, MARRIED, $CHILD^{0-6}$, $CHILD^{7-17}$, CAPITAL, TEMPEMPL, PART-TIME, PUBLOCAL, PIECE-RATE, NODAYWORK, UNEMPL, UNION) as well as seven one-digit industry sector controls (INDU1, INDU2/3, INDU4, INDU5, INDU6, INDU8, INDU9). The omitted educational level variable is BASIC =primary education (about 9 years or less), and the reference industry sector is employment in transport and communication (INDU7). The coefficients estimated for these indicator variables are very close to their counterparts in Tables 2-3 and are therefore not reported here.

* Significant estimate at a 5 % risk level.

** Significant estimate at a 1 % risk level.

Table 5. Sectoral earnings effects of experience for women obtained from estimating the human capital earnings (eq. (6)) augmented with information on years with the current employer (SENiority) and participation in formal on-the-job training (OJT).[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR FEMALE EMPLOYEES (1)	(2)	(3)	(4)	PUBLIC-SECTOR FEMALE EMPLOYEES (1)	(2)	(3)	(4)
EXP	0.0057	0.0038	0.0054	0.0036	0.0049	0.0038	0.0050	0.0040
	(.0049)	(.0048)	(.0048)	(.0048)	(.0048)	(.0046)	(.0048)	(.0046)
EXP^2 /1000	-0.0841	-0.0499	-0.0736	-0.0411	-0.0375	-0.0195	-0.0381	-0.0217
	(.1206)	(.1181)	(.1187)	(.1175)	(.1205)	(.1158)	(.1204)	(.1159)
SEN	0.0056**	0.0048**	0.0053**	0.0046**	0.0100**	0.0078**	0.0100**	0.0077**
	(.0016)	(.0016)	(.0016)	(.0016)	(.0018)	(.0017)	(.0018)	(.0017)
OJT			0.0890**	0.0701**			-0.0100	-0.0279
			(.0219)	(.0216)			(.0213)	(.0208)
Occupation controls	no	yes	no	yes	no	yes	no	yes
R^2 adj.	0.2421	0.2766	0.2527	0.2828	0.3873	0.4440	0.3867	0.4445
SEE	0.3115	0.3027	0.3091	0.3013	0.2846	0.2697	0.2846	0.2694
F-value	13.90	11.98	14.18	12.03	20.89	20.02	20.16	19.55
Number of obs.	1092	1092	1092	1092	882	882	882	882

[1] Bivariate sample selection estimates with standard errors in parentheses below the estimates. The estimated earnings equations also include the personal and job characteristics displayed in Tables 2-3 above (LOWER VOCATIONAL, UPPER VOCATIONAL, SHORT NON-UNIV, UNDERGRADUATE, GRADUATE, MARRIED, $CHILD^{0-6}$, $CHILD^{7-17}$, CAPITAL, TEMPEMPL, PART-TIME, PUBLOCAL, PIECE-RATE, NODAYWORK, UNEMPL, UNION) as well as seven one-digit industry sector controls (INDU1, INDU2/3, INDU4, INDU5, INDU6, INDU8, INDU9). The omitted educational level variable is BASIC =primary education (about 9 years or less), and the reference industry sector is employment in transport and communication (INDU7).
The coefficients estimated for these indicator variables are very close to their counterparts in Tables 2-3 and are therefore not reported here.

* Significant estimate at a 5 % risk level.

** Significant estimate at a 1 % risk level.

sector than for private-sector male employees. This weaker earnings effect of general experience obtained for private-sector males is, however, accompanied with a notable impact of seniority (SEN) and formal on-the-job training (OJT). More exactly, seniority accounts for about one fourth of the initial earnings effect of total work experience (i.e. EXP+SEN) for private-sector males but for only about one tenth for males employed in the public sector (Figure 3).

These results may be seen as supportive of the human capital interpretation of the seniority variable as reflecting earnings effects of primarily aquired specific skills and not of job duration. Put differently, the jobs private-sector men typically hold seem to involve substantial opportunities and/or requirements of investment in training and other productivity improving measures. Male employees in the public sector, on the other hand, tend to acquire mainly general skills, i.e. skills which are thought to be transferable between jobs and/or employers. Possibly this can be interpreted as suggesting that the large internal labour market of the public sector has also come to offer notable career opportunities not necessarily linked to the length of the current employmentship in the public sector.

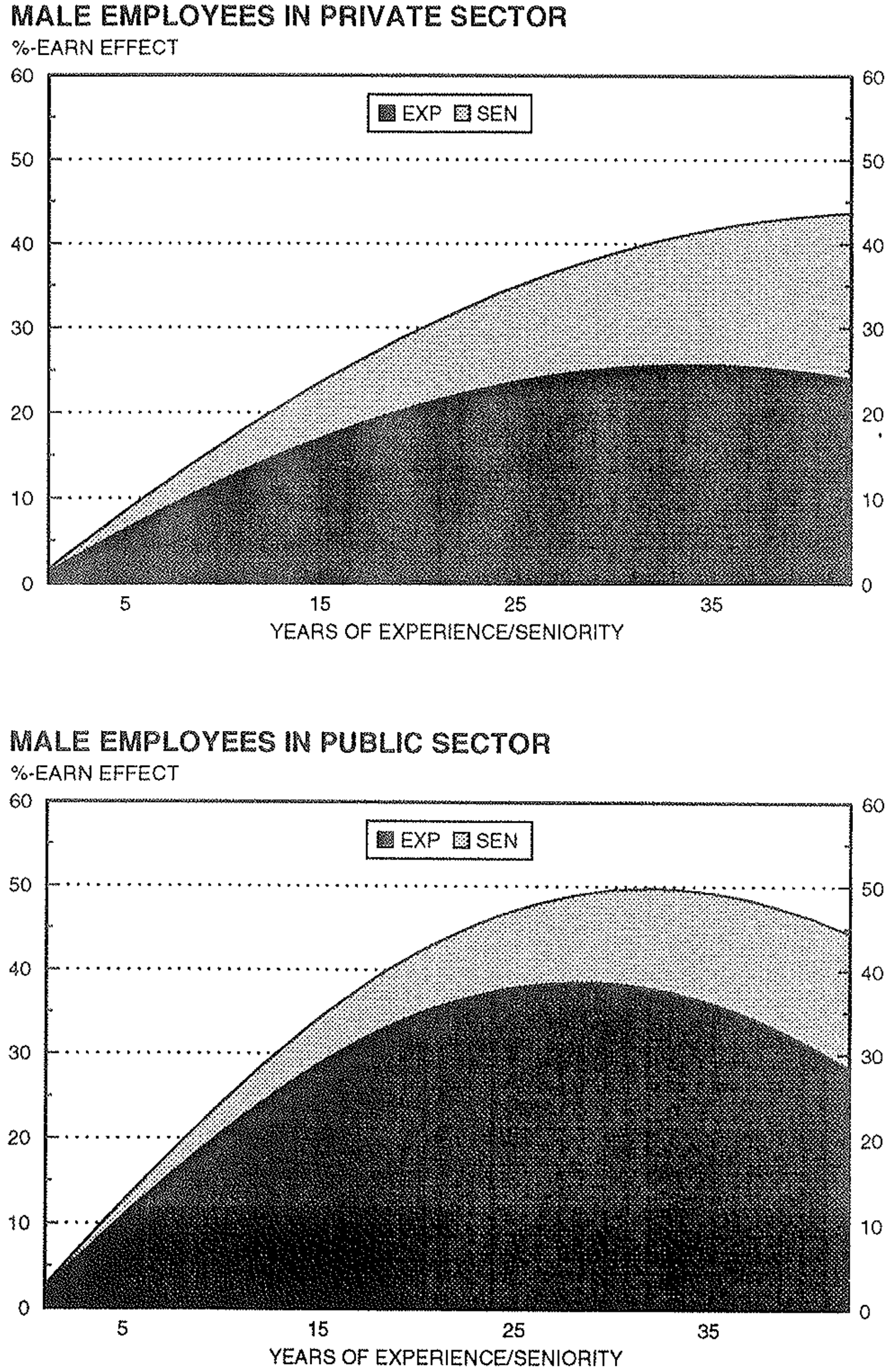
A potential explanation of the somewhat unexpected results concerning experience effects obtained for private- and public-sector men may thus be the differing promotional patterns in the two sectors and the fact that the public sector has weaker possibilities of using wages to create incentives for work effort and higher productivity.[15] But on the other hand, because of the different occupational structures of the private and public sectors and the obvious difficulty of properly accounting for these differences, there is also a possibility that the estimated experience effects reflect to some extent the strong dominance of manual workers in the private sector and of non-manual workers in the public sector. This question is addressed in the next section.

The estimation results point, once again, to a more important role of seniority than of general experience in the determination of female earnings (cf. the results reported in Asplund (1992c)). Indeed, for women seniority accounts for a half or more of the initial earnings effect of total work experience (Figure 4). But apart from this, the situation differs remarkably between females employed in the two sectors. The earnings effects of seniority seem to arise mainly from formal and informal on-the-job training for private-sector females, but almost exclusively from the duration of the current employmentship for public-sector females. Indeed, despite a very high participation rate in formal on-the-job training courses among public-sector women during the survey year, this type of training is found to have had no significant effect on their average earnings level.

All in all, then, the estimation results obtained when augmenting the sectoral earnings equations with seniority and on-the-job training variables point to highly differing earnings effects of labour force experience across both genders and sectors. Earnings growth in the private sector seems to be strongly influenced by the employees' investment in specific training. The most notable difference between private-sector males and females is the significantly stronger earnings effect of general skills obtained for male employees.

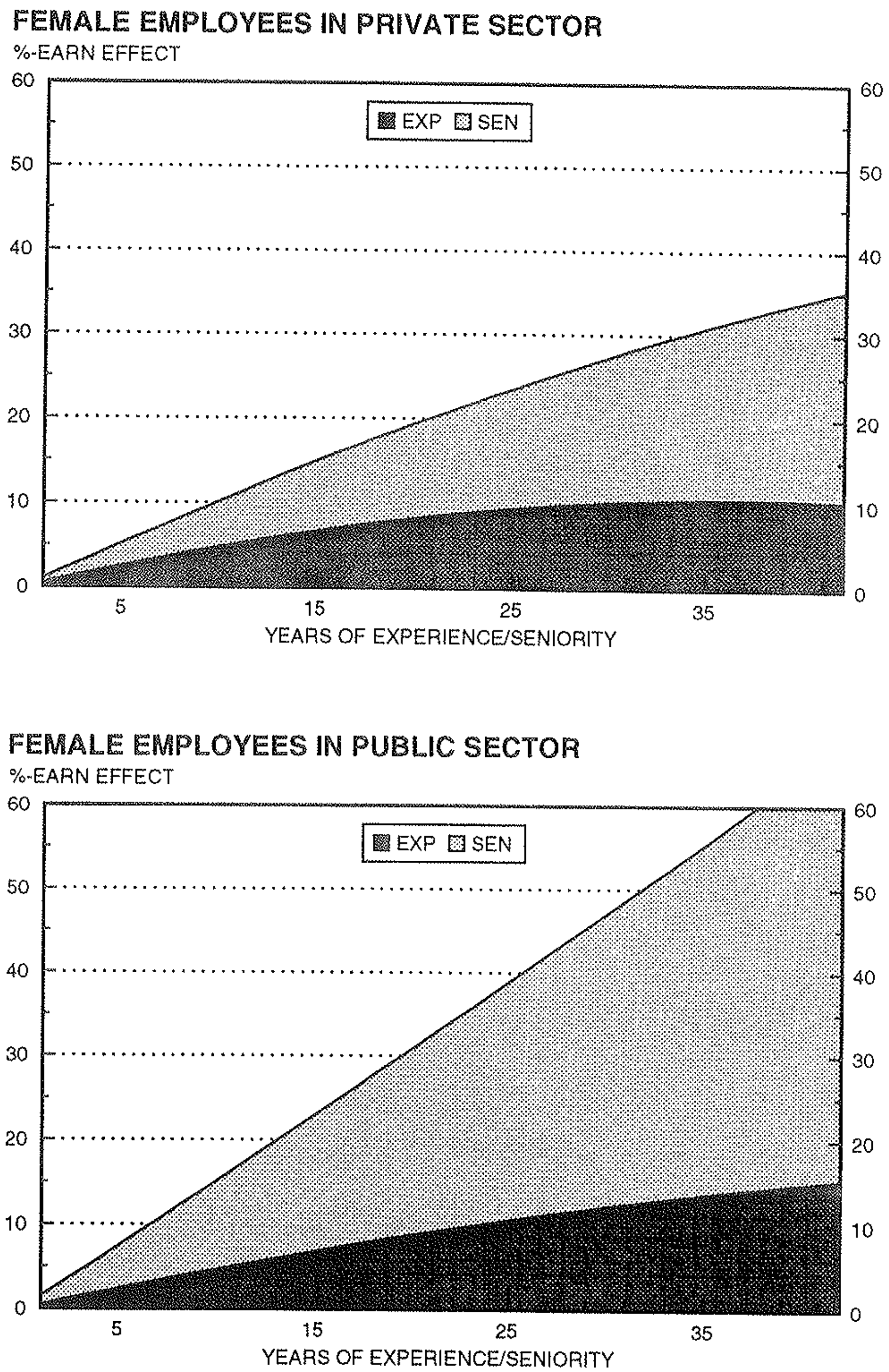
The gender differences in the public sector are much more outstanding. Males employed in the public sector receive a fairly high return on their investments in general skills and OJT, whereas the estimated effects of these two human capital variables on public-sector female earnings are small or negligible. The reverse holds for the estimated earnings effects of seniority. One possible explanation to these fundamental differences in the labour market situation for males and females in public-sector employment may be the fact that a major part of the females are employed at the local government level and that this affects both their promotional pattern and the type of on-the-job training they receive.

Figure 3. Earnings profiles for general experience and seniority, private/public-sector male employees



Source: Calculations for a hypothetical individual staying with the same employer up to 42 years (sample maximum of seniority) based on the estimated coefficients reported in column 3 of Table 4. The calculations are done as in Figure 2 above.

Figure 4. Earnings profiles for general experience and seniority, private/public-sector female employees



Source: Calculations for a hypothetical individual staying with the same employer up to 42 years (sample maximum of seniority) based on the estimated coefficients reported in column 3 of Table 5. The calculations are done as in Figure 2 above.

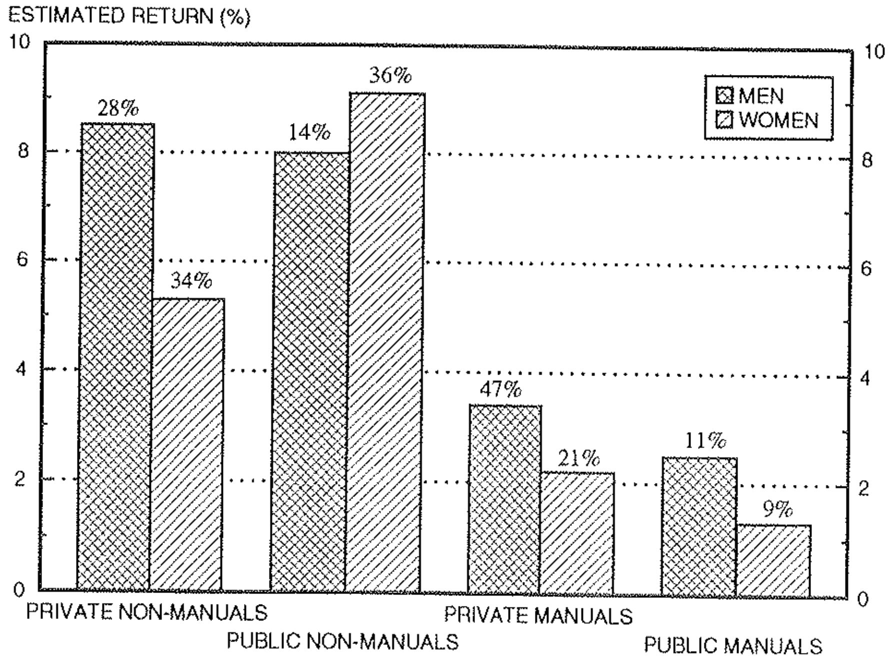
4. SECTORAL EARNINGS STRUCTURES OF MANUALS AND NON-MANUALS

The estimation results reported in the previous section displayed several fundamental similarities and dissimilarities between the earnings structure of the private and the public sector. In particular, the earnings position of private-sector women was found to be fairly weak compared not only to that of male employees but also to that of females employed in the public sector. However, the differing impact of the investigated earnings determinants and, especially, of the various human capital proxies may at least partly originate in the distinct occupational structures of the two sectors. Suppose that non-manual and manual workers are paid different returns on equal units of human capital. In that case, the larger share of manual workers in the private sector could offer an explanation to the observed earnings differentials as well as to the finding of an overall weaker earnings position of females in private-sector employment.

With this point in mind, each sector is divided into two labour markets, one for non-manuals and one for manuals. Both genders thus face four labour markets. (Sample means are given in Table D of Appendix.) In order to maintain comparability with the estimation results for occupational categories reported in Asplund (1992a), employees classified into the occupational social status category of workers in agriculture, forestry and commercial fishing (OCC51) are excluded from the subsequent analysis. As will become apparent from the regression results to be presented below, the general trends in the determination of non-manual and manual earnings found in Asplund (1992a) largely persist also when controlling for sectoral differences.

Various specifications of the human capital earnings model in eq. (17) are estimated separately for the four labour markets and the two genders, whereby correction for potential selection bias arising from labour market choice is done as outlined in Section 2. The gender-specific maximum likelihood estimates for the multinomial logit model in eq. (15) are displayed in Tables I and J of Appendix. In the following, emphasis is on differences between labour markets and genders in the estimated earnings effects of investments in formal education, training and work

Figure 5. Average returns to additional years in schooling after completed compulsory schooling, estimated by gender and labour market



Note. The figure on the top of each pile gives the percentage sample share of, respectively, male and female employees in the labour market in question.

Source: Estimation results not shown.

experience. The gender-specific coefficients for all explanatory variables included in the estimated earnings equations are given in Tables K and L of Appendix.

The regression results reported in Section 3 pointed to small or negligible differences in educational returns between private- and public-sector male employees. The rates of return to education estimated for females in public-sector employment were, except for lower-level educations, found to exceed those received by private-sector women and, moreover, to be very close to the returns paid to male employees. Broadly speaking, these overall trends persist when a further distinction is made between employees in non-manual and manual jobs.

The average return to an additional year in schooling beyond compulsory schooling is estimated at some 8 per cent for males in non-manual jobs and at roughly 3 per cent for males in manual jobs (Figure 5). In other words, there seem to be small, if any, differences across sectors in the average returns to schooling estimated for the broad categories of non-manual and manual workers. The returns to different educational degrees estimated for non-manuals and manuals are also very similar in magnitude across sectors (Table 6 below). Further, irrespective of sector, male employees in manual jobs tend to receive significantly lower returns on their investments in above-primary education as compared to male employees in non-manual jobs.

The average return to an additional year in postcompulsory education is estimated at close to 9 per cent for female non-manuals in public-sector employment (Figure 5). The average rate of return obtained for female non-manuals employed in the private sector is notably lower, or some 5 per cent. Obviously this difference in sectoral returns on education is largely attributable to the substantially higher return to a university graduate degree estimated for public-sector females (Table 7 below). For females in manual jobs, the estimated average return to an additional year's schooling beyond primary education turns out to be insignificant in both sectors and thus significantly lower compared to the average returns received by non-manual females. The insignificant educational return obtained for public-sector manual workers is presumably to most part due to the small variance in educational endowments within the category, while the return estimated for private-sector manuals becomes insignificant only after controlling also for the individuals' industry affiliation.

Comparison of gender-specific rates of return to additional years in postcompulsory schooling suggests that female public-sector non-manuals receive approximately the same average return as male non-manuals, whereas female private-sector non-manuals are paid a markedly lower average return on their investments in formal education. Among manual workers there seem to be small, if any, differences in educational returns across genders and sectors. As to the estimated returns to different educational levels, the most conspicuous gender gap is, once again, obtained for degrees at the lower vocational and professional level.

The sectoral estimates on earnings effects of labour force experience reported in Section 3 indicated that the experience-earnings profiles of male employees tend to be steeper and to also fall off more rapidly in the public sector. When a distinction is made between non-manuals and manuals in the two sectors, these differences across sectors tend to disappear. It may also be noted that the significance levels of the quadratic experience term point to a concave earnings profile for manual workers only.

Further, the sectoral estimates of experience-imputed earnings growth displayed a very moderate but fairly constant growth rate of female earnings with increased experience. Moreover, also among women the experience profile was found to be steeper in the public sector. However, as for male employees, estimation of separate earnings equations for private- and public-sector females employed in, respectively, non-manual and manual jobs reveals no outstanding differences in the estimated earnings effects of total work experience across sectors.

The sectoral estimates reported in Section 3 also suggested that in both sectors, the experience profiles are steeper for men. Comparison of the gender-specific experience coefficients across the four labour markets indicates, in turn, that the gender gaps in the estimated earnings effects of labour market experience are small or negligible, except in public-sector non-manual jobs. However, it is worthwhile emphasizing once again that this impressive similarity in the percentage effects on earnings of more work experience will result in different absolute returns because of largely differing average earnings levels across genders and labour markets.

A division of the earnings effects of total work experience into earnings effects of, respectively, general experience and seniority produced somewhat surprising sectoral estimates for male employees (Table 4 above). The earnings effects of general experience, i.e. of skills which by definition are transferable between jobs and/or employers, were found to be notably stronger for males in public-sector employment, while the earnings effects of seniority turned out to be much stronger for males employed in the private sector. Clearly the opposite finding would have been expected. Dividing the sample employees in each sector into non-manuals and manuals adds to the understanding of this rather puzzling outcome in the sense that now the sectoral differences in earnings growth attributable to general experience and seniority have turned into remarkable conformity within the two broad occupational categories under study.

Specifically, as can be seen from Table 6, the sectoral earnings effects of work experience estimated for male employees in non-manual jobs are of approximately the same size and almost exclusively attributable to general experience (EXP); the estimated coefficients for the quadratic experience variable and the seniority variable (SEN) do not differ significantly from zero. Also among manual workers, there is a strong similarity in the estimated returns to general experience and seniority across sectors. However, compared to their non-manual counterparts, the earnings effects of general experience are significantly lower and those of seniority notably stronger. This outcome may be taken to reflect the different types of working tasks performed by the two employee categories. The most conspicuous remaining earnings difference among male employees is the insignificant return on formal on-the-job training (OJT) obtained for manual workers in the public sector.

The sectoral estimates for females discussed in the previous section pointed to a relatively important role of seniority in the determination of both private- and public-sector female earnings. The overall conclusions drawn for female employees from these sectoral estimates on labour force experience seem to largely hold also when examining separately non-manual and manual females employed in the two sectors (Table 7). There are small, if any, differences among females in the estimated earnings effects of general experience (EXP) across the four labour markets considered. In both sectors, the returns to seniority (SEN) and formal on-the-job training (OJT) are insignificant for female manual workers. Among non-manual females, on the other hand, the seniority effect is found to be stronger in the public sector, whereas the sector's return to OJT is estimated to be negative but insignificant. This is to be compared with female non-manuals employed in the private sector, for whom the weaker seniority effect is accompanied with a relatively strong return on investments in formal OJT. Giving these findings a human capital interpretation, the seniority effects estimated for public-sector non-manuals seem to reflect primarily the length

Table 6. Private/public-sector estimates obtained from eq. (17) for male employees in, respectively, non-manual and manual jobs.[1] The dependent variable is log hourly earnings inclusive of fringe benefits. (The estimation results are fully reported in Table K of Appendix.)

Variable	PRIVATE-SECTOR MEN Non-manuals	Manuals	PUBLIC-SECTOR MEN Non-manuals	Manuals
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1276* (.0550)	0.0614** (.0216)	0.1152 (.0717)	0.0531* (.0272)
UPPER VOCATIONAL	0.2396** (.0486)	0.0932 (.0610)	0.1987** (.0717)	0.0402 (.0692)
SHORT NON-UNIV	0.3622** (.0667)		0.3268** (.0899)	
UNDER-GRADUATE	0.3489** (.1382)		0.3411** (.1151)	
GRADUATE	0.5375** (.0704)		0.5158** (.1002)	
EXP	0.0187** (.0068)	0.0093* (.0043)	0.0236** (.0092)	0.0108* (.0054)
EXP^2 /1000	-0.1812 (.1549)	-0.1766* (.0922)	-0.1601 (.2880)	-0.2376* (.1077)
SEN	0.0001 (.0026)	0.0052** (.0016)	-0.0039 (.0048)	0.0064** (.0020)
OJT	0.1108** (.0320)	0.0940** (.0221)	0.0916** (.0327)	0.0322 (.0244)
UNION	-0.0540* (.0323)	0.0376 (.0285)	-0.0228 (.0577)	-0.0490 (.0459)
LAMBDA (labour market) selectivity)	0.0432 (.0512)	-0.0578 (.0406)	-0.0334 (.0459)	-0.0194 (.0382)
R^2 adj.	0.3447	0.1912	0.5436	0.3620
SEE	0.3373	0.2752	0.2501	0.1744
F-value	11.09	9.65	12.68	5.80
No. of obs.	519	843	256	204

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980). The earnings equations are corrected for potential selectivity bias arising from labour market choice. The corresponding multinomial logit estimates are reported in Table I of Appendix.

* Denotes significant estimate at a 5 % risk level.
** Denotes significant estimate at a 1 % risk level.

Table 7. Private/public-sector estimates obtained from eq. (17) for female employees in, respectively, non-manual and manual jobs.[1] The dependent variable is log hourly earnings inclusive of fringe benefits. (The estimation results are fully reported in Table L of Appendix.)

Variable	PRIVATE SECTOR WOMEN Non-manuals	PRIVATE SECTOR WOMEN Manuals	PUBLIC SECTOR WOMEN Non-manuals	PUBLIC SECTOR WOMEN Manuals
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	-0.0229 (.0366)	-0.0402 (.0356)	0.0225 (.0342)	-0.0055 (.0448)
UPPER VOCATIONAL	0.0809* (.0416)	0.2620** (.1022)	0.1311** (.0330)	0.2137 (.1514)
SHORT NON-UNIV	0.2830** (.0982)		0.2915** (.0531)	
UNDER-GRADUATE	0.4190** (.0722)		0.4570** (.0552)	
GRADUATE	0.3899** (.1153)		0.6047** (.0597)	
EXP	0.0121* (.0061)	0.0132* (.0061)	0.0044 (.0049)	0.0082 (.0114)
EXP^2 /1000	-0.2545* (.1335)	-0.2051 (.1316)	-0.0102 (.1177)	-0.0989 (.1905)
SEN	0.0062** (.0016)	-0.0011 (.0024)	0.0110** (.0016)	0.0031 (.0025)
OJT	0.0752** (.0224)	0.0490 (.0355)	-0.0286 (.0215)	-0.0432 (.0290)
UNION	-0.0724* (.0326)	-0.0426 (.0420)	0.0671* (.0405)	-0.0747 (.0934)
LAMBDA (labour market selectivity)	0.0443 (.0304)	0.0916** (.0376)	-0.0542 (.0451)	0.0110 (.0791)
R^2 adj.	0.2503	0.2208	0.4054	0.0808
SEE	0.3128	0.2962	0.2852	0.2723
F-value	9.68	6.00	18.60	1.91
No. of obs.	677	407	698	177

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980). The earnings equations are corrected for potential selectivity bias arising from labour market choice. The corresponding multinomial logit estimates are reported in Table J of Appendix.

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

of the current employmentship, while the seniority effects obtained for private-sector non-manuals rather point to an important role of acquired specific skills.

Comparison of the earnings effects of general experience and seniority across genders and labour markets reveals no significant gender gap in the estimated earnings effects of general experience within the broad categories of private-sector non-manuals and manuals. Contrary to this, notable gender gaps show up in the estimated earnings effects of seniority and formal OJT. The role of seniority in the determination of private-sector earnings is much stronger for females than for males in non-manual jobs, but clearly weaker for females than for males in manual jobs. Among both employee categories, men receive a higher return on investments in formal OJT in the private sector.

A picture similar to that observed for private-sector non-manuals emerges when comparing public-sector men and women in non-manual jobs; the return to seniority is substantially higher among female non-manuals, whereas the return to formal OJT is significantly higher among male non-manuals. But contrary to the situation in the private sector, the earnings of male non-manuals tend to be more strongly affected also by general skills. The gender differences in experience and OJT effects among manual workers employed in the public sector are almost negligible.

Finally a few comments on the estimates of the union membership (UNION) and selectivity (LAMBDA) variables also displayed in Tables 6 and 7. Not surprisingly, a significant negative earnings effect on the part of unionized employees is obtained for private-sector non-manuals. Less expected is perhaps the finding that the same effect shows up for both male and female non-manuals. Union membership, when measured by means of a simple indicator variable, seems to have no significant earnings effect in the other labour markets. The only exception is female non-manuals in public-sector employment for whom union membership is found to imply a negative earnings effect. Their relative income disadvantage is caused by a fairly small number of non-unionized females working mainly in temporary and/or part-time jobs in the local government service sector (SIC9).

The estimated coefficients for the selectivity term indicate that there is generally no serious selectivity bias arising from labour market choice influencing the estimation results. Accordingly ordinary least squares techniques would produce consistent parameter estimates. These are reported in Tables M and N of Appendix. There is one exception, though. The strong significant selectivity effect obtained for females in private-sector manual jobs suggests, when evaluated at the sample mean level of LAMBDA, that females entering this particular labour market earn, on average, some 5 per cent less than a female with identical observable characteristics drawn at random from the labour force would be expected to earn in that labour market. It may be noted that a much stronger crowding-in effect was estimated for females in non-manufacturing jobs in Asplund (1992a, Table 2).

5. CONCLUDING REMARKS

The present paper has focused on displaying similarities and dissimilarities in the earnings structure of the private and the public sector in Finland. Special attention has thereby been paid to the sectoral returns on investments in human capital received by male and female employees, with a further distinction made between males and females employed in, respectively, non-manual and manual jobs in the two sectors. In summarizing the general findings of the paper, emphasis will be on comparing the estimated human capital effects on sectoral earnings with empirical evidence for the 1980s available for the other Nordic countries. The comparison concerns private/public-sector estimates only, since no previous evidence has been found based on a division of employees into non-manual and manual workers by sector *and* gender. It should, however, be kept in mind that the differences in estimates across the Nordic countries that will become evident in the following are not necessarily "true" differences but may at least partly be due to differences in, inter alia, sample data sets used, variables included in the estimations, and estimation methods employed.

The regression results point to small, if any, differences in educational returns between private- and public-sector male employees. The rates of return on education estimated for females in public-sector employment are generally found to exceed those received by private-sector women and, moreover, to be very close to the returns paid to male employees. However, in both sectors females tend to get a very moderate return on low educations, which shows up in notable gender gaps at the lower end of the educational scale. Further, non-manuals in the public sector and male non-manuals in the private sector are estimated to receive approximately the same average return to additional years in postcompulsory schooling, whereas female non-manuals in private-sector employment are paid a significantly lower average return on their investments in formal education. Among manual workers, there seem to be no notable differences in educational returns across genders and sectors.

Estimation results obtained for the other Nordic countries display a slightly different pattern of educational returns. Estimates for Sweden reveal no significant differences in educational returns across sectors when education is measured in years (Zetterberg, 1988), but point to a lower return in the public sector when education is accounted for by means of educational level indicators (Kazamaki & D'Agostino, 1992). The two studies provide an ambiguous picture of the gender gaps in sectoral returns on education in Sweden. The Norwegian estimates reported by Barth & Mastekaasa (1990) point to a higher return on education in the private sector for both genders. However, there seem to be no significant gender gaps in educational returns within the two sectors in Norway. Estimates reported for Denmark (Pedersen et al., 1990) display a higher return on education in the public sector for both genders, and a higher return for men than for women within each sector.

The estimation results reported in the present paper also suggest that for both genders, the experience-earnings profiles are on average steeper in the public sector. Further, in both sectors the experience curves are found to be flatter for women. However, when a distinction is made between non-manuals and manuals in the two sectors, these differences across sectors and genders become small or vanish; the gender gap in the earnings effects of work experience is significant among public-sector non-manuals only.

Again the empirical evidence obtained for Finland differs from that reported for the other Nordic countries. In particular, the experience effects have been estimated to be stronger in the private sector in Norway for both genders (Barth & Mastekaasa, 1990) and in Sweden for women (Zetterberg (1988), Kazamaki & D'Agostino (1992)). The earnings effects of labour force experience are estimated to be of approximately the same magnitude in the two sectors for Swedish men. This holds for both genders in Denmark (Pedersen et al., 1990). The Swedish estimates point to small, if any, gender gaps in the estimated experience effects, while the Danish and Norwegian results suggest that the return to experience is lower for women in both sectors.

A division of the earnings effects of total work experience into earnings effects of, respectively, general experience and seniority produce somewhat surprising sectoral estimates for Finnish male employees; the earnings effects of general experience are found to be notably stronger for males in public-sector employment, while the earnings effects of seniority turn out to be clearly stronger for males employed in the private sector. Intuitively, a stronger seniority effect would be expected in the public sector; this is the general finding in the other Nordic countries (cf. Arai (1991), Kazamaki & D'Agostino (1992), Asplund et al. (1993)).

However, dividing the sample employees in each sector into non-manuals and manuals changes markedly this rather puzzling picture of experience effects on earnings obtained for Finnish men; irrespective of sector, the earnings effects of general experience are found to be significantly higher for non-manual males, while those of seniority are estimated to be significantly higher for manual males. This outcome may well be taken to reflect the different types of working tasks performed by non-manual and manual workers. The most conspicuous remaining earnings differential among male employees is the insignificant return on formal on-the-job training obtained for manual workers in the public sector.

The sectoral estimates for females suggest that seniority has a relatively important role in the determination of both private- and public-sector female earnings. However, when examining separately non-manuals and manuals employed in the two sectors, these seniority effects show up for non-manual employees only, and more strongly in the public sector. It is also noteworthy that despite a relatively high participation rate in formal on-the-job training programmes among female employees, only private-sector women in non-manual jobs tend to receive a significant return on such investments.

In view of this, it is hardly surprising that the estimates point to notable gender gaps in the estimated earnings effects of both seniority and formal on-the-job training. Among non-manuals, the seniority effects are much stronger for females than for males, while the reverse holds among manuals in private sector jobs. Also, in all three employee categories men generally receive a higher return on investments in formal on-the-job training. Among public-sector manual workers, the gender differences in experience and on-the-job training effects are almost negligible.

All in all, the empirical evidence on the earnings structure of the private and the public sector in Finland presented in this paper does seem to indicate that the earnings effects of especially general and specific work experience estimated from sectoral earnings equations are to some extent "distorted" by the strong dominance of manual workers in the private sector and of non-manual workers in the public sector. A further division of the employees in each sector into non-manuals and manuals provides support of this hypothesis; the estimation results now point to a high degree of similarity in sectoral returns on investments in human capital within the broad categories of non-manual and manual workers, implying that the differences in human capital returns tend to be larger between occupational categories than between sectors.

Finally, the estimation results do not seem to point to the presence of a "double-imbalance" problem in the Finnish labour market. Instead the comparatively high returns on investments in human capital received in the public sector can be expected to attract also high-educated individuals to the sector's large and, until recent years, rapidly growing number of both upper- and lower-level non-manual jobs.[16]

Footnotes:

1. The human capital earnings equation is derived and commented on in Asplund (1992d).

2. This compromise is based on two observations. First, empirical evidence reported in Asplund (1992a,1992c) and in Section 3 of the present paper suggests that this source of selection bias is generally no serious problem in the employed data set. Second, a tractable model specification maintaining the labour force participation decision for the sample individuals would require that decision to be independent of the decision on which of the four labour markets to enter (cf. Dolton et al., 1989). As indicated by the estimates on the selectivity (RHO) variable in Table E of Appendix, this can be argued to be a fairly strong a priori restriction on the regressions for females.

3. The Type 1 distribution is occasionally referred to as the exponential or Gumbel distribution (Johnson & Kotz, 1970).

4. The multinomial logit model is preferred to the unordered multinomial probit model because it is less difficult to estimate, and to the ordered probit model because it does not require a sequential ranking of labour markets which would involve arbitrary judgements. Moreover, the ordered probit model has been found to predict less well than the multinomial logit model (cf. de Beyer & Knight (1989) and Reilly (1991)).

5. Specifically, the probability of participation in labour market m, $Prob(LM_i=m)$, is estimated in relation to the labour market, say, k chosen for the purpose of normalization. This implies estimation of three functions of the form

$$(i)\quad \ln\left[\frac{Prob(LM_i=m)}{Prob(LM_i=k)}\right] = \delta_m + \theta_m V_i + \eta_{im}, \qquad m = 1,2,3,4 \quad m \neq k$$

where $Prob(LM_i=m)/Prob(LM_i=k)$ is the ratio of the probability of participating in labour market m to that of participating in labour market k, and δ is a constant term. A comparison of any labour markets m and r can then be derived as

$$(ii)\quad \ln\left[\frac{Prob(LM_i=m}{Prob(LM_i=r)}\right] = \ln\left[\frac{Prob(LM_i=m)}{Prob(LM_i=k)}\right] - \ln\left[\frac{Prob(LM_i=r)}{Prob(LM_i=k)}\right]$$
$$= (\delta_m - \delta_r) + (\theta_m - \theta_r)V_i + (\eta_{im} - \eta_{ir})$$

6. Since the income data come from the tax rolls, the information concerns the tax value of taxable fringe benefits. In 1987, the tax value of fringe benefits was, on average, 80 per cent of their market value. Further, virtually all fringe benefits had by 1987 become subject to taxation, the most important exceptions being reasonable health and recreational benefits financed by the employer.

7. Cf. the regression results reported in Table H of Appendix. This outcome may be partly explained by the fact that the tax rolls provide information merely on the tax value of fringe benefits subject to taxation. Moreover, of the male employees retained in the actual estimating data only some 16 per cent are recorded to have received taxable fringe benefits in 1987. The corresponding share for female employees is close to 20 per cent.

8. A t-test for testing the statistical significance of the difference between single coefficients estimated for males/females employed in, respectively, the private and the public sector cannot be done as the sectoral parameter estimates are correlated due to the correction for potential sector selection bias undertaken in the estimations. I am indebted to Markus Jäntti for making me aware of this problem.

9. Following Halvorsen & Palmquist (1980), the percentage differential for indicator variables obtained from estimating various specifications of the semilogarithmic human capital earnings function is calculated as $(e^{\alpha} - 1)*100$.

10. The cumulative earnings effect of labour market experience (EXP) measures total percentage additions to earnings due to experience and is calculated as the antilog of $(\alpha_1 EXP - \alpha_2 EXP^2)$.

11. The quite large number of public-sector employees (both manual and non-manual workers) in industry sectors other than the service sector (SIC9) explain the introduction of the full set of job-related variables also into the public-sector earnings equations.

12. A more frequent occurrence of temporary unemployment and layoffs in 1987 among especially less-paid private-sector males may offer a potential explanation to the negative earnings effect of UNEMPL obtained for that category. The strong negative earnings effect estimated for public-sector males may, in turn, be the outcome of the special arrangements in force at that time. In particular, a fairly large number of the public-sector staff was still in 1987 paid by means of budgetary employment grants (Finnish Labour Review 1/1990, Tables 22-23). Apart from this, also central and local government jobs were to a certain extent filled with the aid of pay subsidies. The lack of similar effects for female employees may simply be due to their lower average

hourly earnings level and smaller wage dispersion (cf. Table 1).

13. If acquiring general skills, i.e. skills which are by definition transferable across jobs/employers, the employee is thought to have to pay all the costs of his or her training and to also receive the full return from his or her accumulated investment in training. If acquiring specific skills, on the other hand, the employee and the employer are likely to share both the costs and the returns associated with the training. In the first (second) case, starting wages will be lower (higher) and earnings growth with increased experience faster (slower). For critics of the Becker theorem, see e.g. Ballot (1992).

14. The estimated coefficients of the quadratic seniority term were throughout insignificant, and the variable was therefore abandoned in the estimations. (Cf. the results reported in Asplund (1992a,1992c).) For a more detailed discussion of the theoretical considerations and the empirical implications of this approach, see Asplund (1992c,1992d).

15. Arai (1991) presents evidence on this for Sweden.

16. Possibly this can be taken to actualize the wage-twist policy option discussed in Pedersen et al. (1990): "When (these) high tax rates reflect that a major part of the labour force is employed in the public sector, more than they reflect a high share of cash transfers, a new policy option is open for government. If the public sector can use its role as a dominant employer to reduce relative public sector wages, it can lessen the effects both on and from the tax pressure with a given relative size of public sector employment. Such a wage policy would further tend to reduce the number of public employees by voluntary mobility from the public sector. In a political sense this could be an attractive option as an alternative to direct cuts of in-kind transfers which will usually be strongly resisted by both the public employees in question and by the users of public institutions" (p. 126). The authors conclude that the hypothesis about a wage-twist policy being pursued in Denmark in the period 1976-85 cannot be rejected.

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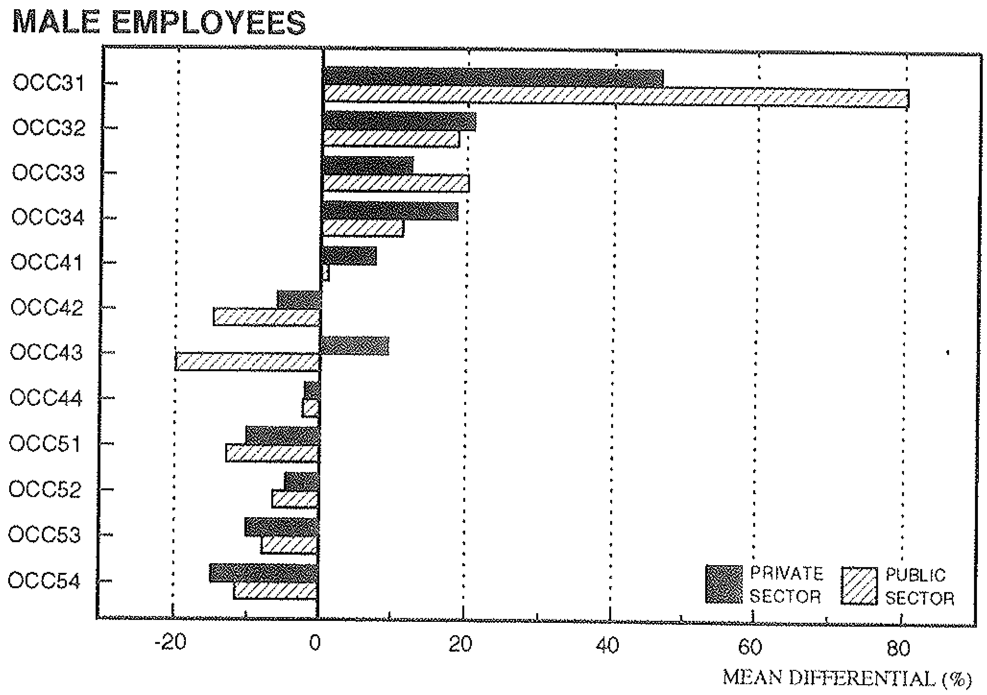
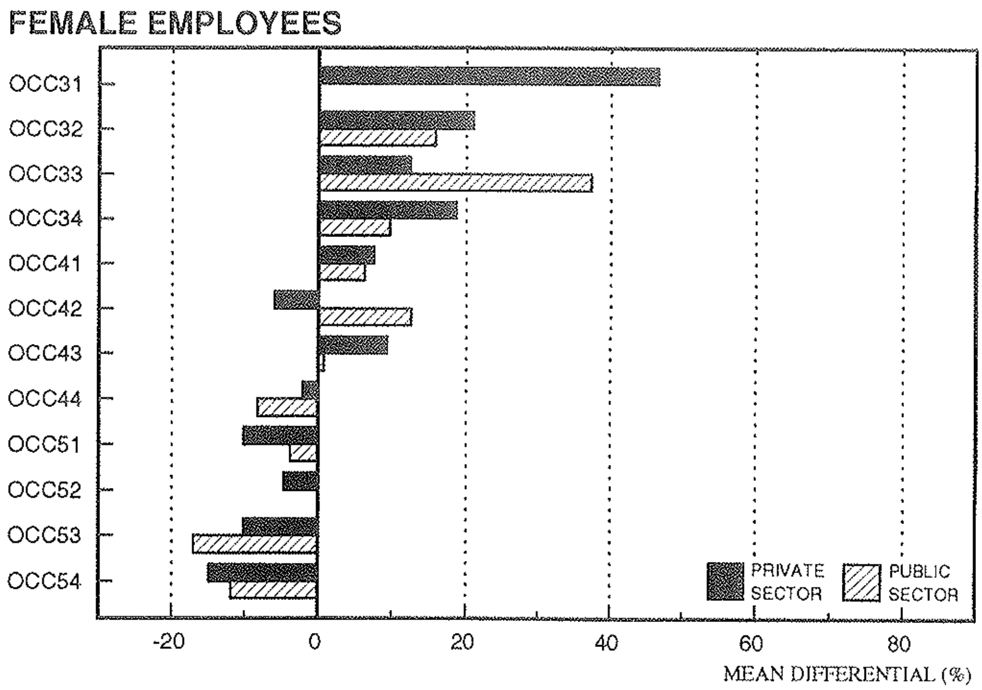
APPENDIX

Figure A. Employment weighted mean differentials in hourly earnings levels inclusive of fringe benefits between 12 occupational social status categories after having controlled for various background factors, by gender and sector





Source: Calculated from the occupation coefficients reported in Tables 2-3 in the text and the sample means given in Table B of Appendix. For the formulae used, see footnote 11 in Asplund (1993).

Table A. Summary of definitions of included variables

Variable	Definition
EARN	Average hourly earnings (in FIM) calculated from the before-tax annual wage/salary income recorded in the tax rolls and an estimated amount of annual normal working hours. The earnings data include most types of compensation, including fringe benefits.
ln EARN	Natural logarithm of EARN.
SCHOOL	Years of formal schooling evaluated from register information on the highest single education completed using the Finnish standard classification of education.
S	Years of formal schooling with basic education (9 years of schooling) set equal to zero.
BASIC	Indicator for persons with basic education only (about 9 years or less).
LOWER VOCATIONAL	Indicator for persons with completed lower-level of upper secondary education (about 10-11 years).
UPPER VOCATIONAL	Indicator for persons with completed upper-level of upper secondary education (about 12 years).
SHORT NON-UNIV	Indicator for persons with completed lowest level of higher education (about 13-14 years).
UNDERGRADUATE	Indicator for persons with completed graduate university education (about 15 years).
GRADUATE	Indicator for persons with completed graduate university education (more than 16 years).
EXP	Self-reported total years of labour market experience.
SEN	Seniority, i.e. self-reported years with the present employer.
AGE	Physical age of the individual.
MARRIED	Indicator for married persons and singles living together.
$CHILD^{0-6}$	Indicator for children aged 0 to 6 living at home.
$CHILD^{7-17}$	Indicator for children aged 7 to 17 living at home.
$CHILD^{0-17}$	Indicator for children aged 0 to 17 living at home.
CAPITAL	Indicator for residence within the capital region (the region of Helsinki).
UUSIMAA	Indicator for residence in the province of Uusimaa but outside the capital region.
OTHER SOUTH	Indicator for residence in the southern parts of Finland other than Uusimaa.
MIDDLE	Indicator for residence in the middle parts of Finland.
NORTH	Indicator for residence in the northern parts of Finland.
PUBLOCAL	Indicator for employment in the local government (municipality) sector.
TEMPEMPL	Indicator for persons who self-reportedly are in temporary employment.
PART-TIME	Indicator for persons who self-reportedly are in part-time employment.

PIECE-RATE	Indicator for persons who are not being paid on an hourly, weekly or monthly basis.
NODAYWORK	Indicator for persons who are not in regular day-time work.
UNEMPL	Indicator for persons who have been temporarily unemployed or laid off during the survey year.
UNION	Indicator for unionized employees.
OJT	Indicator for persons who self-reportedly have received employer-sponsored formal on-the-job-training during the survey year.
OCC31	Indicator for senior officials and upper management.
OCC32	Indicator for senior officials and employees in research and planning.
OCC33	Indicator for senior officials and employees in education and training.
OCC34	Indicator for other senior officials and employees.
OCC41	Indicator for supervisors.
OCC42	Indicator for clerical and sales workers, independent work.
OCC43	Indicator for clerical and sales workers, routine work.
OCC44	Indicator for other lower-level employees with administrative and clerical occupations.
OCC51	Indicator for workers in agriculture, forestry and commercial fishing.
OCC52	Indicator for manufacturing workers.
OCC53	Indicator for other production workers.
OCC54	Indicator for distribution and service workers.
INDU1	Indicator for employment in agriculture, forestry and fishing.
INDU2/3	Indicator for employment in mining, quarrying and manufacturing.
INDU4	Indicator for employment in electricity, gas and water.
INDU5	Indicator for employment in construction.
INDU6	Indicator for employment in trade, restaurants and hotels.
INDU7	Indicator for employment in transport and communication.
INDU8	Indicator for employment in financing, insurance, reale estate and business services.
INDU9	Indicator for employment in public, social and personal services.
WORK1	Indicator for persons in technical, physical science, social science, humanistic and artistic work.
WORK2	Indicator for persons in managerial, administrative, and clerical work.
WORK3	Indicator for persons in commercial work.
WORK4	Indicator for persons in agriculture, forestry, and fishing.
WORK5	Indicator for persons in manufacturing work, mining and quarrying.
WORK6	Indicator for persons in transport and communication work.
WORK7	Indicator for persons in health care and social work.
WORK8	Indicator for persons in other service work.

Table B. Sample statistics by sector for all male and female employees retained in the actual estimating data[1]

Variable	MALE EMPLOYEES Private sector	Public sector	FEMALE EMPLOYEES Private sector	Public sector
EARN	49.12	50.84	39.65	42.86
	(23.57)	(25.59)	(24.31)	(22.71)
ln EARN	3.81	3.85	3.60	3.68
	(0.38)	(0.37)	(0.36)	(0.37)
SCHOOL (S + 9)	10.77	11.70	10.54	11.55
	(1.76)	(2.39)	(1.65)	(2.11)
BASIC (1,0)	0.3746	0.2907	0.4522	0.2624
LOWER VOCA-TIONAL (1,0)	0.3577	0.2701	0.2684	0.2996
UPPER VOCA-TIONAL (1,0)	0.1841	0.1753	0.2220	0.2117
SHORT NON-UNIV (1,0)	0.0401	0.0887	0.0218	0.1081
UNDER-GRADUATE (1,0)	0.0056	0.0454	0.0209	0.0529
GRADUATE (1,0)	0.0379	0.1299	0.0146	0.0653
EXP	17.15	18.34	16.27	15.98
	(10.98)	(10.39)	(9.95)	(9.65)
SEN	8.58	11.23	8.42	8.82
	(8.36)	(8.93)	(8.04)	(8.14)
MARRIED (1,0)	0.7344	0.7650	0.7234	0.7410
$CHILD^{0-6}$ (1,0)	0.2621	0.2474	0.1920	0.2309
$CHILD^{7-17}$ (1,0)	0.3183	0.3629	0.3576	0.3840
CAPITAL (1,0)	0.1855	0.1691	0.2320	0.1768
TEMPEMPL (1,0)	0.0611	0.1155	0.0682	0.1813
PART-TIME (1,0)	0.0091	0.0206	0.0682	0.0518
PUBLOCAL (1,0)	-	0.4969	-	0.7522
PIECE-RATE (1,0)	0.1483	0.0227	0.1110	0.0079
NODAYWORK (1,0)	0.2228	0.2557	0.2466	0.2511
UNEMPL (1,0)	0.1174	0.0784	0.0955	0.1014
UNION (1,0)	0.6859	0.8577	0.7307	0.8761
OJT (1,0)	0.3290	0.4644	0.3159	0.4580
OCC31 (1,0)	0.0808	0.0144	0.0191	-
OCC32 (1,0)	0.0450	0.0763	0.0136	0.0180
OCC33 (1,0)	0.0028	0.1443	0.0018	0.1047
OCC34 (1,0)	0.0372	0.0639	0.0528	0.0754
OCC41 (1,0)	0.1047	0.1113	0.0419	0.0428
OCC42 (1,0)	0.0731	0.0165	0.2884	0.1036
OCC43 (1,0)	0.0077	0.0103	0.1474	0.1104
OCC44 (1,0)	0.0162	0.1010	0.0546	0.3356
OCC51 (1,0)	0.0232	0.0350	0.0073	0.0079
OCC52 (1,0)	0.4020	0.0907	0.1902	-
OCC53 (1,0)	0.0991	0.0907	0.0491	0.0642
OCC54 (1,0)	0.1082	0.2454	0.1338	0.1374
INDU1 (1,0)	0.0358	0.0330	0.0082	0.0113
INDU2/3 (1,0)	0.4336	0.0103	0.3412	0.0034
INDU4 (1,0)	0.0176	0.0309	0.0064	0.0045
INDU5 (1,0)	0.1511	0.1237	0.0227	0.0045
INDU6 (1,0)	0.1462	0.0124	0.2821	0.0146
INDU7 (1,0)	0.0737	0.2289	0.0400	0.0619
INDU8 (1,0)	0.0773	0.0227	0.1665	0.0338
INDU9 (1,0)	0.0647	0.5381	0.1328	0.8660
Number of obs.	1423	485	1099	888

[1] The figures in parentheses below the continuous variables give the standard deviation of the variable in question.

Table C. Sample means for selected variables by gender, sector, and occupational status

	Private-sector male employees					Public-sector male employees				
	EARN	$EARN^{NO\ FRINGES}$	SCHOOL	EXP	No. of obs.	EARN	$EARN^{NO\ FRINGES}$	SCHOOL	EXP	No. of obs.
OCC31	80.77	76.58	12.31	20.35	115	114.79	113.40	14.00	27.00	7
OCC32	67.88	66.47	13.64	12.47	64	71.22	71.22	14.89	17.38	37
OCC33	56.92	56.92	14.50	13.75	4	72.02	71.37	14.76	18.70	70
OCC34	63.71	62.25	13.08	16.09	53	63.15	63.04	13.64	13.64	31
OCC41	53.59	52.70	11.22	20.13	149	47.41	47.37	11.13	18.65	54
OCC42	44.43	43.46	10.91	14.88	104	40.91	40.76	13.00	16.00	8
OCC43	60.50	59.92	12.09	14.18	11	33.50	33.48	11.80	9.60	5
OCC44	47.37	47.22	11.39	12.91	23	45.25	45.21	11.24	14.12	49
OCC51	36.96	36.68	9.97	17.54	33	35.06	35.02	9.35	21.71	17
OCC52	44.48	44.39	10.15	17.17	572	40.32	40.26	10.02	21.11	44
OCC53	40.05	39.92	10.07	16.79	141	39.04	39.01	10.34	18.20	44
OCC54	38.86	38.84	9.95	16.85	154	40.85	40.74	10.07	19.82	119

	Private-sector female employees					Public-sector female employees				
	EARN	$EARN^{NO\ FRINGES}$	SCHOOL	EXP	No. of obs.	EARN	$EARN^{NO\ FRINGES}$	SCHOOL	EXP	No. of obs.
OCC31	62.47	60.16	12.48	16.48	21	-	-	-	-	-
OCC32	49.82	48.98	14.07	11.20	15	45.96	45.76	13.12	12.88	16
OCC33	55.32	55.32	14.50	10.00	2	65.82	65.29	14.70	15.95	93
OCC34	53.51	53.21	12.41	16.02	58	50.95	50.77	14.10	12.85	67
OCC41	39.59	39.22	10.37	19.67	46	45.10	45.10	11.60	19.66	38
OCC42	39.43	39.14	10.68	16.04	317	43.11	42.90	11.16	17.51	92
OCC43	39.86	39.02	10.40	16.41	162	37.78	37.57	10.74	15.08	98
OCC44	48.93	47.93	10.73	14.87	60	40.57	40.50	11.28	14.52	298
OCC51	26.47	26.47	10.00	10.75	8	36.51	36.51	10.14	8.71	7
OCC52	33.03	33.00	9.88	16.85	209	26.29	26.29	11.00	18.00	1
OCC53	34.45	34.22	10.17	17.46	54	31.22	30.91	10.32	16.82	56
OCC54	38.19	38.06	10.04	15.76	147	35.20	35.12	9.80	20.15	122

Table D. Sample means for male and female employees by sector and occupational status
(Standard deviations of continuous variables are given in parentheses.)

Variable	MALE EMPLOYEES Private sector Non-manuals	Manuals	Public sector Non-manuals	Manuals	FEMALE EMPLOYEES Private sector Non-manuals	Manuals	Public sector Non-manuals	Manuals
EARN	60.47	42.67	59.27	40.45	42.50	35.12	45.21	33.82
	(28.43)	(17.36)	(23.54)	(8.91)	(26.02)	(20.69)	(23.52)	(17.00)
ln EARN	4.01	3.70	4.01	3.68	3.66	3.48	3.73	3.47
	(0.42)	(0.31)	(0.37)	(0.22)	(0.36)	(0.34)	(0.37)	(0.28)
SCHOOL (S + 9)	11.94	10.11	13.14	10.12	10.88	9.98	11.97	9.97
	(2.08)	(1.08)	(2.28)	(1.15)	(1.81)	(1.16)	(2.12)	(1.10)
BASIC (1,0)	0.1927	0.4745	0.0938	0.4853	0.3900	0.5577	0.1877	0.5480
LOWER VOCATIONAL (1,0)	0.1773	0.4698	0.1406	0.4412	0.2098	0.3612	0.2765	0.3842
UPPER VOCATIONAL (1,0)	0.4027	0.0546	0.2773	0.0637	0.3117	0.0762	0.2478	0.0678
SHORT NON-UNIV (1,0)	0.1079	0.0012	0.1562	0.0098	0.0325	0.0025	0.1375	-
UNDERGRADUATE (1,0)	0.0154	-	0.0859	-	0.0340	-	0.0673	-
GRADUATE (1,0)	0.1040	-	0.2461	-	0.0222	0.0025	0.0831	-
EXP	17.27	17.02	17.01	19.98	16.18	16.49	15.22	19.04
	(10.52)	(11.26)	(9.54)	(11.17)	(9.75)	(10.17)	(9.19)	(10.79)
SEN	9.04	8.35	10.71	12.01	8.71	8.05	8.72	9.41
	(8.55)	(8.27)	(8.69)	(9.30)	(8.24)	(7.70)	(8.10)	(8.41)
MARRIED (1,0)	0.8208	0.6845	0.8242	0.7206	0.7223	0.7297	0.7393	0.7458
$CHILD^{0-6}$ (1,0)	0.2929	0.2432	0.2695	0.2157	0.1994	0.1867	0.2522	0.1356
$CHILD^{7-17}$ (1,0)	0.3738	0.2871	0.4141	0.3039	0.3560	0.3636	0.4012	0.3277
CAPITAL (1,0)	0.2929	0.1281	0.1797	0.1765	0.2998	0.1253	0.1905	0.1299
TEMPEMPL (1,0)	0.0405	0.0700	0.1328	0.1029	0.0709	0.0614	0.1877	0.1638
PART-TIME (1,0)	0.0173	0.0047	0.0195	0.0245	0.0591	0.0786	0.0473	0.0678
PUBLOCAL (1,0)	-	-	0.5312	0.4510	-	-	0.7507	0.7571
PIECE-RATE (1,0)	0.0597	0.1886	-	0.0245	0.0162	0.2727	0.0860	-
NODAYWORK (1,0)	0.0963	0.2859	0.2070	0.3137	0.1551	0.3907	0.2478	0.2373
UNEMPL (1,0)	0.0328	0.1601	0.0547	0.1029	0.0665	0.1400	0.0788	0.1921
UNION (1,0)	0.5838	0.7758	0.8555	0.8872	0.6987	0.8010	0.8825	0.8870
OJT (1,0)	0.5607	0.1850	0.6016	0.2941	0.4165	0.1548	0.5229	0.2034
INDU1 (1,0)	0.0347	0.0024	0.0117	0.0049	0.0074	0.0025	0.0057	-
INDU2/3 (1,0)	0.3083	0.5243	-	0.0245	0.1846	0.6093	0.0029	0.0056
INDU4 (1,0)	0.0193	0.0166	0.0078	0.0637	0.0059	0.0074	0.0057	-
INDU5 (1,0)	0.0674	0.2076	0.0781	0.1912	0.0266	0.0172	0.0014	0.0169
INDU6 (1,0)	0.2563	0.0854	0.0117	0.0147	0.3471	0.1794	0.0100	0.0339
INDU7 (1,0)	0.0385	0.0985	0.1094	0.3971	0.0502	0.0221	0.0544	0.0904
INDU8 (1,0)	0.1792	0.0178	0.0156	0.0343	0.2393	0.0442	0.0387	0.0169
INDU9 (1,0)	0.0963	0.0474	0.7656	0.2696	0.1388	0.1179	0.8811	0.8362
Number of obs.	519	843	256	204	677	407	698	177

Table E. Full information maximum likelihood estimates of the bivariate probit model (eq. (15) in the text) explaining the probability of labour force participation and private-sector employment, by gender[1]

Variable	Males	Sample mean	Females	Sample mean
Working selection:				
CONSTANT	-6.7401** (.8271)		-5.5965** (.7310)	
BASIC EDUCATION	-0.2901** (.0648)	0.4506	-0.3250** (.0571)	0.4641
AGE	0.4752** (.0718)	37.25	0.3484** (.0625)	38.42
AGE^2	-0.0087** (.0019)	1572.7	-0.0047** (.0016)	1662.4
AGE^3 /1000	0.0381** (.0161)	73038	0.0034 (.0135)	78637
MARRIED	0.5614** (.0793)	0.6368	0.0786 (.0648)	0.6630
$CHILD^{0-17}$	0.3412** (.0929)	0.3487	-0.1818** (.0682)	0.3946
SOUTH	0.3514** (.0615)	0.6046	0.3940** (.0524)	0.6257
Private-sector selection:				
CONSTANT	0.8822** (.1787)		0.0322 (.1755)	
BASIC EDUCATION	0	0.3533	0	0.3673
VOCATIONAL (levels 3-4)	-0.1028 (.0900)	0.5173	-0.0880 (.0863)	0.4998
HIGHER EDUC. (levels 5-8)	-0.8178** (.1374)	0.1294	-0.5202** (.1301)	0.1329
EXP	-0.0150** (.0039)	17.46	-0.0058 (.0037)	16.14
CAPITAL	0	0.1814	0	0.2073
UUSIMAA	-0.0426 (.1441)	0.0896	-0.1606 (.1368)	0.0856
OTHER SOUTH	-0.0780 (.0992)	0.3789	-0.1949* (.0916)	0.3855
MIDDLE	-0.3950** (.1168)	0.1562	-0.4855** (.1115)	0.1530

Table E. (cont.)

Variable	Males	Sample mean	Females	Sample mean
NORTH	-0.3848** (.1116)	0.1939	-0.6956** (.1125)	0.1686
MARRIED	0.1261 (.1034)	0.7421	-0.0461 (.0812)	0.7321
PART-TIME	-0.3774 (.2762)	0.0121	0.2819* (.1347)	0.0609
NODAYWORK	0.0644 (.0934)	0.2311	0.4368** (.0860)	0.2486
WORK1	0	0.1845	0	0.1208
WORK2	0.5114** (.1216)	0.1132	0.7902** (.1129)	0.2672
WORK3	1.8839** (.3479)	0.0592	2.4844** (.2975)	0.0810
WORK4	-0.0138 (.1805)	0.0351	0.0967 (.3183)	0.0091
WORK5	0.6849** (.1056)	0.4308	2.2213** (.2228)	0.1243
WORK6	-0.4182** (.1288)	0.1059	-0.4200* (.1993)	0.0337
WORK7	-1.2612** (.2448)	0.0210	-0.8636** (.1242)	0.1691
WORK8	-0.7551** (.1653)	0.0503	-0.0028 (.1220)	0.1948
RHO(μ_1, μ_2)	0.1842 (.1311)		0.2725** (.1160)	
Log-Likelihood	-1996.3		-2508.9	
No of obs. in				
- working selection	2825		3193	
- sector selection	1908		1987	

[1] Standard errors are given in parentheses below the estimates.
* Denotes significant estimate at a 5 % risk level.
** Denotes significant estimate at a 1 % risk level.

Table F. Private/public-sector estimates of eqs. (1) and (2) for males using the standard Heckman two-stage procedure and ordinary least squares (OLS) techniques.[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	Private sector males Heckman	Private sector males OLS	Public sector males Heckman	Public sector males OLS
CONSTANT	3.3419^{**} (.0481)	3.3271^{**} (.0472)	3.3188^{**} (.0801)	3.2957^{**} (.0803)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1068^{**} (.0213)	0.1056^{**} (.0214)	0.0969^{**} (.0332)	0.0988^{**} (.0261)
UPPER VOCATIONAL	0.2940^{**} (.0264)	0.2916^{**} (.0286)	0.2498^{**} (.0380)	0.2522^{**} (.0342)
SHORT NON-UNIV	0.4675^{**} (.0508)	0.4518^{**} (.0444)	0.4893^{**} (.0519)	0.4994^{**} (.0590)
UNDER-GRADUATE	0.5331^{**} (.1157)	0.5210^{**} (.1153)	0.4796^{**} (.0648)	0.4878^{**} (.0566)
GRADUATE	0.6617^{**} (.0512)	0.6485^{**} (.0534)	0.6396^{**} (.0489)	0.6491^{**} (.0462)
EXP	0.0184^{**} (.0031)	0.0183^{**} (.0036)	0.0233^{**} (.0048)	0.0233^{**} (.0054)
EXP^2 /1000	-0.2615^{**} (.0738)	-0.2638^{**} (.0817)	-0.3491^{**} (.1099)	-0.3458^{**} (.1267)
MARRIED	0.0593^{**} (.0237)	0.0595^{**} (.0239)	0.0641^{*} (.0343)	0.0626^{*} (.0317)
$CHILD^{0-6}$	-0.0030 (.0219)	-0.0034 (.0197)	-0.0578^{*} (.0300)	-0.0583^{*} (.0271)
$CHILD^{7-17}$	0.0547^{**} (.0211)	0.0548^{**} (.0205)	0.0468^{*} (.0279)	0.0470^{*} (.0273)
CAPITAL	0.1684^{**} (.0232)	0.1696^{**} (.0277)	0.0650^{*} (.0313)	0.0618^{*} (.0319)
TEMPEMPL	-0.0598 (.0386)	-0.0601 (.0499)	-0.0025 (.0446)	-0.0022 (.0495)
PART-TIME	0.1180 (.0919)	0.1121 (.1990)	0.3500^{**} (.0886)	0.3542^{*} (.1621)
PUBLOCAL			-0.0027 (.0253)	-0.0014 (.0250)
PIECE-RATE	0.0488^{*} (.0243)	0.0508^{*} (.0220)	0.1665^{*} (.0852)	0.1657^{*} (.0796)
NODAYWORK	0.0493^{*} (.0215)	0.0486^{**} (.0207)	0.0506^{*} (.0292)	0.0552^{*} (.0291)

Table F. (cont.)

Variable	Private sector males Heckman	OLS	Public sector males Heckman	OLS
UNEMPL	-0.0753**	-0.0747**	-0.1579**	-0.1580**
	(.0294)	(.0309)	(.0517)	(.0595)
UNION	-0.0023	-0.0029	0.0045	0.0052
	(.0197)	(.0218)	(.0367)	(.0481)
LAMBDA2(ϵ,μ_2)	-0.0316		0.0173	
(sector sel.)	(.0414)		(.0336)	
R^2 adj.	0.3150	0.3152	0.5417	0.5424
SEE	0.3138	0.3166	0.2420	0.2488
F-value	27.15	28.27	23.00	23.95
Number of obs.	1423	1423	485	485

[1] Standard errors are given in parentheses below the estimates. Heckman (1979) estimates where LAMBDA2 (ϵ,μ_2) measures the selectivity bias arising from choosing between the two sectors. The probit estimates for the sector selection function are very close to those given in Table D above and are therefore not reported here. Since maximum likelihood estimates (see Asplund (1992a,1992c) for an explanation of this estimation method) were obtained for private-sector employees only, standard Heckman estimates are reported for both employee categories.
The OLS-estimates are corrected for heteroscedasticity according to White (1980).
The estimated earnings equations also include seven one-digit industry sector controls (INDU1,INDU2/3,INDU4,INDU5,INDU6, INDU8,INDU9), employment in transport and communication (INDU7) being the reference sector.

* Significant estimate at a 5 % risk level.

** Significant estimate at a 1 % risk level.

Table G. Private/public-sector estimates of eqs. (1) and (2) for females using the standard Heckman two-stage procedure and ordinary least squares (OLS) techniques.[1]
The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	Private sector females Heckman	Private sector females OLS	Public sector females Heckman	Public sector females OLS
CONSTANT	3.3155** (.0640)	3.3398** (.0775)	3.3219** (.0671)	3.3321** (.0764)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	-0.0228 (.0246)	-0.0163 (.0250)	0.0435 (.0284)	0.0413 (.0279)
UPPER VOCATIONAL	0.1733** (.0284)	0.1772** (.0352)	0.2090** (.0310)	0.2072** (.0334)
SHORT NON-UNIV	0.3585** (.0694)	0.3916** (.0788)	0.3924** (.0397)	0.3864** (.0347)
UNDER-GRADUATE	0.4772** (.0706)	0.5006** (.0709)	0.5506** (.0501)	0.5453** (.0426)
GRADUATE	0.4997** (.0831)	0.5303** (.1041)	0.6875** (.0473)	0.6820** (.0439)
EXP	0.0177** (.0037)	0.0174** (.0042)	0.0121** (.0040)	0.0122** (.0043)
EXP^2 /1000	-0.3227** (.0931)	-0.3114** (.0959)	-0.0936 (.1010)	-0.0964 (.0988)
MARRIED	-0.0092 (.0230)	-0.0081 (.0220)	-0.0348 (.0244)	-0.0356 (.0226)
$CHILD^{0-6}$	0.0290 (.0258)	0.0313 (.0274)	0.0373 (.0254)	0.0374 (.0297)
$CHILD^{7-17}$	-0.0079 (.0218)	-0.0061 (.0189)	0.0198 (.0225)	0.0197 (.0204)
CAPITAL	0.1335** (.0246)	0.1293** (.0272)	0.0471* (.0271)	0.0489* (.0250)
TEMPEMPL	0.0576 (.0406)	0.0588 (.0570)	0.0936** (.0326)	0.0924** (.0364)
PART-TIME	0.2313** (.0398)	0.2254** (.0602)	0.3736** (.0470)	0.3761** (.0851)
PUBLOCAL			-0.0301 (.0295)	-0.0343 (.0232)
PIECE-RATE	0.0018 (.0332)	-0.0081 (.0318)	-0.0942 (.1133)	-0.0941 (.1334)
NODAYWORK	0.1149** (.0236)	0.1149** (.0253)	0.1694** (.0234)	0.1693** (.0280)

Table G. (cont.)

Variable	Private sector females Heckman	OLS	Public sector females Heckman	OLS
UNEMPL	-0.0451 (.0339)	-0.04478 (.0388)	-0.0421 (.0381)	-0.0422 (.0492)
UNION	-0.0568** (.0240)	-0.0548* (.0272)	0.0478 (.0320)	0.0472 (.0376)
LAMBDA2(ϵ,μ_2) (sector sel.)	0.0459* (.0275)		-0.0130 (.0320)	
R^2 adj.	0.2298	0.2286	0.3627	0.3633
SEE	0.3145	0.3185	0.2907	0.2951
F-value	14.10	14.56	20.41	21.24
Number of obs.	1099	1099	888	888

[1] Standard errors are given in parentheses below the estimates. Heckman (1979) estimates where LAMBDA2 (ϵ,μ_2) measures the selectivity bias arising from choosing between the two sectors. The probit estimates for the sector selection function are fairly close to those given in Table D above and are therefore not reported here. Since maximum likelihood estimates (see Asplund (1992a,1992c) for an explanation of this estimation method) were obtained for private-sector employees only, standard Heckman estimates are reported for both employee categories.
The OLS-estimates are corrected for heteroscedasticity according to White (1980).
The estimated earnings equations also include seven one-digit industry sector controls (INDU1,INDU2/3,INDU4,INDU5,INDU6, INDU8,INDU9), employment in transport and communication (INDU7) being the reference sector.

* Significant estimate at a 5 % risk level.

** Significant estimate at a 1 % risk level.

Table H. Private/public-sector estimates of eqs. (1) and (2) for male and female employees using OLS.[1] The dependent variable is log hourly earnings *exclusive* of fringe benefits.

Variable	MALE EMPLOYEES Private sector	Public sector	FEMALE EMPLOYEES Private sector	Public sector
CONSTANT	3.3269** (.0469)	3.2924** (.0805)	3.3239** (.0781)	3.3279** (.0764)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1051** (.0211)	0.0990** (.0260)	-0.0152 (.0246)	0.0409 (.0280)
UPPER VOCATIONAL	0.2813** (.0281)	0.2532** (.0342)	0.1784** (.0348)	0.2062** (.0334)
SHORT NON-UNIV	0.4430** (.0419)	0.4958** (.0587)	0.3885** (.0788)	0.3834** (.0347)
UNDER-GRADUATE	0.4926** (.1122)	0.4801** (.0565)	0.5061** (.0707)	0.5426** (.0425)
GRADUATE	0.6158** (.0541)	0.6488** (.0459)	0.5176** (.0993)	0.6801** (.0440)
EXP	0.0181** (.0035)	0.0237** (.0055)	0.0176** (.0041)	0.0123** (.0043)
EXP^2 /1000	-0.2609* (.0806)	-0.3547** (.1277)	-0.3128** (.0947)	-0.0986 (.0988)
MARRIED	0.0557** (.0235)	0.0626* (.0316)	-0.0098 (.0218)	-0.0376* (.0226)
$CHILD^{0-6}$	-0.0026 (.0193)	-0.0577* (.0270)	0.0282 (.0273)	0.0392 (.0297)
$CHILD^{7-17}$	0.0527** (.0202)	0.0452* (.0272)	-0.0081 (.0187)	0.0210 (.0204)
CAPITAL	0.1607** (.0273)	0.0624* (.0317)	0.1291** (.0268)	0.0495* (.0250)
TEMPEMPL	-0.0578 (.0498)	-0.0014 (.0494)	0.0581 (.0568)	0.0934** (.0364)
PART-TIME	0.1285 (.1984)	0.3564* (.1624)	0.2137** (.0592)	0.3787** (.0851)
PUBLOCAL		-0.0011 (.0249)		-0.0336 (.0232)
PIECE-RATE	0.0479* (.0216)	0.1636* (.0791)	-0.0061 (.0316)	-0.0956 (.1340)
NODAYWORK	0.0494** (.0207)	0.0566* (.0291)	0.1174** (.0252)	0.1726** (.0280)
UNEMPL	-0.0727** (.0308)	-0.1581** (.0596)	-0.0446 (.0386)	-0.0450 (.0493)
UNION	0.0061 (.0215)	0.0052 (.0480)	-0.0499* (.0269)	0.0508 (.0377)
R^2 adj.	0.3031	0.5417	0.2236	0.3636
SEE	0.3126	0.2480	0.3156	0.2950
F-value	26.77	23.88	14.17	21.27
Number of obs.	1423	485	1099	888

[1] For notes, see Table G above.

Table I. Multinomial logit estimates of eq. (15) for labour market choice equations, male employees[1]

Variable	Public-sector manuals	Private-sector non-manuals	Private-sector manuals
CONSTANT	2.8908** (.6773)	0.2571 (.5516)	3.2677** (.6648)
BASIC EDUCATION	0	0	0
LOWER VOCATIONAL	-1.2079** (.4140)	-0.1057 (.3499)	-1.5837** (.4051)
UPPER VOCATIONAL	-3.1004** (.5137)	-0.3786 (.3161)	-3.5624** (.4865)
SHORT NON-UNIV	-3.4756** (.9120)	-1.0478** (.3616)	-5.2409** (1.3202)
UNDER-GRADUATE		-2.9725** (.5419)	
GRADUATE		-1.9363** (.3681)	
EXP	-0.0207 (.0157)	-0.0213* (.0102)	-0.0538** (.0154)
MARRIED	-0.1939 (.3759)	0.2787 (.2496)	-0.2027 (.3608)
CAPITAL	0	0	0
UUSIMAA	-1.1623* (.6175)	-0.3908 (.3615)	-0.9271 (.5807)
OTHER SOUTH	-0.9682* (.4327)	-0.6050** (.2451)	-0.6276 (.4172)
MIDDLE	-0.1940 (.5221)	-1.2156** (.3190)	-0.4829 (.5187)
NORTH	-0.4068 (.4816)	-1.2298** (.2865)	-0.6686 (.4775)
PART-TIME	-0.0501 (.9202)	-0.0735 (.6748)	-1.0711 (.9946)
NODAYWORK	-0.3178 (.3267)	-0.4912* (.2749)	0.1213 (.3200)
WORK1		1.8263** (.3766)	
WORK2		3.0743** (.4237)	
WORK3		2.9760** (.4425)	
WORK4-5	2.3875** (.5506)	1.7664** (.6330)	4.7957** (.5494)
WORK6	0.8369* (.3878)	0.3276 (.5307)	1.4164** (.3942)
WORK7-8	0	0	0

Log-Likelihood = -1037.6 Corr. pred.[2] = 76.7 %
Chi-square (54) = 2505.6 No. of obs. = 1858

[1] Standard errors are in parentheses below the estimates. The reference category is public-sector non-manuals.
[2] Percentage of correctly predicted labour market status.
* Denotes significant estimate at a 5 % risk level.
** Denotes significant estimate at a 1 % risk level.

Table J. Multinomial logit estimates of eq. (15) for market choice equations, female employees[1]

Variable	Public-sector manuals	Private-sector non-manuals	Private-sector manuals
CONSTANT	0.1433 (.3978)	-0.6037* (.3000)	0.1363 (.3822)
BASIC EDUCATION	0	0	0
LOWER VOCATIONAL	-0.9130** (.2272)	-0.4805** (.1999)	-0.9356** (.2305)
UPPER VOCATIONAL	-2.1618** (.3559)	-0.3682* (.1949)	-1.9190** (.3230)
SHORT NON-UNIV		-1.1649** (.3000)	-4.1677** (1.0291)
UNDER-GRADUATE		-1.0791** (.3282)	
GRADUATE		-1.5130** (.3694)	-4.3054** (1.3480)
EXP	0.0180* (.0102)	-0.0093 (.0080)	-0.0007 (.0103)
MARRIED	-0.2800 (.2304)	-0.1394 (.1551)	-0.3200 (.2232)
CAPITAL	0	0	0
UUSIMAA	-0.2405 (.4588)	-0.4843* (.2626)	0.1220 (.4139)
OTHER SOUTH	0.3372 (.3042)	-0.3739* (.1766)	0.2331 (.2853)
MIDDLE	0.1753 (.3491)	-0.7750** (.2250)	-0.4849 (.3505)
NORTH	0.4960 (.3366)	-1.0415** (.2270)	-0.4830 (.3518)
PART-TIME	0.4236 (.4262)	0.0918 (.3138)	1.4464** (.3563)
NODAYWORK	-0.7112** (.2196)	0.2515 (.1949)	0.6063** (.2038)
WORK1		1.2710** (.2231)	
WORK2		2.4320** (.1913)	
WORK3		5.9841** (.7307)	
WORK4-5	1.0079 (.7353)	0.6127 (1.1680)	5.2067** (.6159)
WORK6	-0.5673 (.3538)	0.2863 (.3994)	-0.8296* (.4122)
WORK7-8	0	0	0

Log-Likelihood = -1351.6 Corr. pred.[2] = 69.6 %
Chi-square (54) = 2342.0 No. of obs. = 1972

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1 Standard errors are in parentheses below the estimates. The reference category is public-sector non-manuals.
2 Percentage of correctly predicted labour market status.
* Denotes significant estimate at a 5 % risk level.
** Denotes significant estimate at a 1 % risk level.

Table K. Private/public-sector estimates of eq. (17) for male employees in non-manual and manual jobs, respectively.[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR MEN Non-manuals	PRIVATE-SECTOR MEN Manuals	PUBLIC-SECTOR MEN Non-manuals	PUBLIC-SECTOR MEN Manuals
CONSTANT	3.3859** (.1143)	3.3458** (.0550)	3.3451** (.1591)	3.5193** (.1009)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1276* (.0550)	0.0614** (.0216)	0.1152 (.0717)	0.0531* (.0272)
UPPER VOCATIONAL	0.2396** (.0486)	0.0932 (.0610)	0.1987** (.0717)	0.0402 (.0692)
SHORT NON-UNIV	0.3622** (.0667)		0.3268** (.0899)	
UNDER-GRADUATE	0.3489** (.1382)		0.3411** (.1151)	
GRADUATE	0.5375** (.0704)		0.5158** (.1002)	
EXP	0.0187** (.0068)	0.0093* (.0043)	0.0236** (.0092)	0.0108* (.0054)
EXP^2 /1000	-0.1812 (.1549)	-0.1766* (.0922)	-0.1601 (.2880)	-0.2376* (.1077)
SEN	0.0001 (.0026)	0.0052** (.0016)	-0.0039 (.0048)	0.0064** (.0020)
OJT	0.1108** (.0320)	0.0940** (.0221)	0.0916** (.0327)	0.0322 (.0244)
MARRIED	0.0581 (.0442)	0.0433 (.0264)	0.0454 (.0518)	0.0158 (.0317)
$CHILD^{0-6}$	-0.0119 (.0355)	0.0080 (.0218)	-0.0042 (.0381)	-0.0470 (.0359)
$CHILD^{7-17}$	0.0907** (.0341)	0.0227 (.0227)	0.0566* (.0328)	0.0600* (.0295)
CAPITAL	0.1804** (.0367)	0.1795** (.0394)	0.0702 (.0494)	0.0740* (.0346)
TEMPEMPL	-0.2105* (.1023)	0.0071 (.0574)	-0.0457 (.0662)	0.1782** (.0733)
PART-TIME	0.2415 (.2305)	0.0078 (.3237)	0.6032** (.1741)	-0.0154 (.1751)
PUBLOCAL			0.0272 (.0338)	-0.0083 (.0294)
PIECE-RATE	0.1502* (.0664)	0.0534** (.0222)		0.1027 (.1043)

Table K. (cont.)

Variable	PRIVATE-SECTOR MEN Non-manuals	PRIVATE-SECTOR MEN Manuals	PUBLIC-SECTOR MEN Non-manuals	PUBLIC-SECTOR MEN Manuals
NODAYWORK	-0.1080 (.0656)	0.1217** (.0206)	-0.0005 (.0522)	0.0838** (.0296)
UNEMPL	0.0584 (.0816)	-0.0665* (.0343)	-0.2055* (.1051)	-0.2721** (.0808)
UNION	-0.0540* (.0323)	0.0376 (.0285)	-0.0228 (.0577)	-0.0490 (.0459)
INDU1	-0.2152** (.1168)			
INDU2/3	0.0421 (.0893)			
INDU1-3		0.0716* (.0330)	-0.2133** (.0734)	-0.0688 (.0565)
INDU4	0.0448 (.1111)	0.1293* (.0717)	0.2670** (.0966)	0.1192* (.0592)
INDU5	0.0231 (.1024)	0.1562** (.0394)	-0.0781 (.0596)	-0.0544 (.0337)
INDU6	-0.1090 (.0902)	0.0133 (.0472)	-0.3722* (.1842)	-0.1286 (.1149)
INDU7	0	0	0	0
INDU8	0.0579 (.0908)	-0.0584 (.0443)	0.0485 (.1981)	0.0999 (.0696)
INDU9	-0.1274 (.1094)	-0.0722 (.0466)	0.0221 (.0464)	-0.0867** (.0336)
LAMBDA (labour market selectivity)	0.0432 (.0512)	-0.0578 (.0406)	-0.0334 (.0459)	-0.0194 (.0382)
R^2 adj.	0.3447	0.1912	0.5436	0.3620
SEE	0.3373	0.2752	0.2501	0.1744
F-value	11.09	9.65	12.68	5.80
No. of obs.	519	843	256	204

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980). The earnings equations are corrected for potential selectivity bias arising from labour market choice. The corresponding multinomial logit estimates are reported in Table I above.

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

Table L. Private/public-sector estimates of eq. (17) for female employees in non-manual and manual jobs, respectively.[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR WOMEN Non-manuals	Manuals	PUBLIC-SECTOR WOMEN Non-manuals	Manuals
CONSTANT	3.3552** (.1013)	3.1952** (.1615)	3.4074** (.0806)	3.3157** (.2354)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	-0.0229 (.0366)	-0.0402 (.0356)	0.0225 (.0342)	-0.0055 (.0448)
UPPER VOCATIONAL	0.0809* (.0416)	0.2620** (.1022)	0.1311** (.0330)	0.2137 (.1514)
SHORT NON-UNIV	0.2830** (.0982)		0.2915** (.0531)	
UNDER-GRADUATE	0.4190** (.0722)		0.4570** (.0552)	
GRADUATE	0.3899** (.1153)		0.6047** (.0597)	
EXP	0.0121* (.0061)	0.0132* (.0061)	0.0044 (.0049)	0.0082 (.0114)
EXP^2 /1000	-0.2545* (.1335)	-0.2051 (.1316)	-0.0102 (.1177)	-0.0989 (.1905)
SEN	0.0062** (.0016)	-0.0011 (.0024)	0.0110** (.0016)	0.0031 (.0025)
OJT	0.0752** (.0224)	0.0490 (.0355)	-0.0286 (.0215)	-0.0432 (.0290)
MARRIED	-0.0294 (.0281)	0.0243 (.0319)	-0.0497* (.0251)	-0.0074 (.0454)
$CHILD^{0-6}$	0.0487 (.0359)	-0.0077 (.0389)	0.0554* (.0313)	-0.0868 (.0694)
$CHILD^{7-17}$	-0.0022 (.0227)	-0.0106 (.0294)	0.0197 (.0224)	0.0031 (.0442)
CAPITAL	0.1116** (.0337)	0.1028* (.0449)	0.0389 (.0249)	0.1350* (.0717)
TEMPEMPL	0.0050 (.0731)	0.2043* (.0997)	0.0847* (.0401)	0.0809 (.0901)
PART-TIME	0.2814** (.0746)	0.3376** (.0868)	0.4955** (.1100)	0.1224 (.0900)
PUBLOCAL			-0.0381 (.0264)	0.0051 (.0397)
PIECE-RATE	-0.0472 (.1213)	0.0758* (.0340)	-0.1750 (.1651)	

Table L. (cont.)

Variable	PRIVATE-SECTOR WOMEN Non-manuals	PRIVATE-SECTOR WOMEN Manuals	PUBLIC-SECTOR WOMEN Non-manuals	PUBLIC-SECTOR WOMEN Manuals
NODAYWORK	0.1105** (.0378)	0.1408** (.0298)	0.1550** (.0323)	0.1420** (.0530)
UNEMPL	-0.0433 (.0462)	-0.0071 (.0553)	-0.0053 (.0646)	0.0082** (.0622)
UNION	-0.0724* (.0326)	-0.0426 (.0420)	0.0671* (.0405)	-0.0747 (.0934)
INDU1-3	0.0553 (.0487)	0.0394 (.1377)	0.1161 (.1333)	
INDU4	0.0216 (.0682)	-0.1250 (.1684)	-0.0204 (.0695)	
INDU5	0.0461 (.0742)	0.0808 (.1379)	-0.1404** (.0583)	
INDU6	-0.0462 (.0501)	-0.0676 (.1355)	0.2641 (.2017)	
INDU7	0	0	0	
INDU8	0.1359** (.0491)	-0.2541 (.1566)	0.0329 (.0655)	
INDU9	0.0811 (.0600)	-0.0990 (.1440)	-0.0483 (.0588)	
LAMBDA (labour market selectivity)	0.0443 (.0304)	0.0916** (.0376)	-0.0542 (.0451)	0.0110 (.0791)
R^2 adj.	0.2503	0.2208	0.4054	0.0808
SEE	0.3128	0.2962	0.2852	0.2723
F-value	9.68	6.00	18.60	1.91
No. of obs.	677	407	698	177

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980). The earnings equations are corrected for potential selectivity bias arising from labour market choice. The corresponding multinomial logit estimates are reported in Table J above.

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

Table M. Private/public-sector estimates of eqs. (1) and (2) for male employees in, respectively, non-manual and manual jobs using ordinary least squares (OLS) techniques.[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR MEN Non-manuals	PRIVATE-SECTOR MEN Manuals	PUBLIC-SECTOR MEN Non-manuals	PUBLIC-SECTOR MEN Manuals
CONSTANT	3.4292^{**} (.1186)	3.3199^{**} (.0542)	3.2760^{**} (.1309)	3.4955^{**} (.0847)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	0.1299^{**} (.0554)	0.0557^{**} (.0215)	0.1250^{*} (.0690)	0.0528^{*} (.0272)
UPPER VOCATIONAL	0.2411^{**} (.0485)	0.0767 (.0596)	0.2204^{**} (.0633)	0.0366 (.0691)
SHORT NON-UNIV	0.3714^{**} (.0642)		0.3613^{**} (.0751)	
UNDER-GRADUATE	0.3837^{**} (.1294)		0.3923^{**} (.0844)	
GRADUATE	0.5538^{**} (.0653)		0.5584^{**} (.0751)	
EXP	0.0180^{**} (.0067)	0.0089^{*} (.0042)	0.0235^{**} (.0092)	0.0108^{*} (.0054)
EXP^2 /1000	-0.1658 (.1512)	-0.1769^{*} (.0918)	-0.1498 (.2854)	-0.2333^{*} (.1071)
SEN	0.0002 (.0026)	0.0054^{**} (.0016)	-0.0039 (.0048)	0.0065^{**} (.0020)
OJT	0.1096^{**} (.0318)	0.0964^{**} (.0222)	0.0930^{**} (.0323)	0.0312 (.0247)
MARRIED	0.0569 (.0446)	0.0421 (.0264)	0.0397 (.0515)	0.0150 (.0314)
$CHILD^{0-6}$	-0.0140 (.0353)	0.0075 (.0219)	-0.0015 (.0385)	-0.0484 (.0358)
$CHILD^{7-17}$	0.0923^{**} (.0339)	0.0246 (.0226)	0.0594^{*} (.0330)	0.0596^{*} (.0294)
CAPITAL	0.1724^{**} (.0368)	0.1792^{**} (.0393)	0.0587 (.0479)	0.0735^{*} (.0346)
TEMPEMPL	-0.2063^{*} (.1052)	0.0079 (.0574)	-0.0419 (.0670)	0.1768^{**} (.0740)
PART-TIME	0.2339 (.2330)	-0.0325 (.3148)	0.6039^{**} (.1726)	-0.0062 (.1748)
PUBLOCAL			0.0302 (.0338)	-0.0092 (.0299)
PIECE-RATE	0.1441^{*} (.0666)	0.0554^{**} (.0221)		0.0997 (.1025)

Table M. (cont.)

Variable	PRIVATE-SECTOR MEN Non-manuals	PRIVATE-SECTOR MEN Manuals	PUBLIC-SECTOR MEN Non-manuals	PUBLIC-SECTOR MEN Manuals
NODAYWORK	-0.0995 (.0660)	0.1223** (.0206)	0.0129 (.0468)	0.0844** (.0293)
UNEMPL	0.0501 (.0830)	-0.0657* (.0343)	-0.2059* (.1055)	-0.2724** (.0816)
UNION	-0.0538* (.0323)	0.0389 (.0285)	-0.0223 (.0574)	-0.0500 (.0459)
INDU1	-0.1931* (.1090)			
INDU2/3	0.0217 (.0935)			
INDU1-3		0.0911** (.0319)	-0.2210** (.0786)	-0.0767 (.0555)
INDU4	0.0284 (.1142)	0.1459* (.0723)	0.2677** (.0974)	0.1114* (.0549)
INDU5	0.0057 (.1057)	0.1757** (.0381)	-0.0706 (.0578)	-0.0598* (.0341)
INDU6	-0.1328 (.0946)	0.0240 (.0482)	-0.3706* (.1879)	-0.1279 (.1156)
INDU7	0	0	0	0
INDU8	0.0366 (.0946)	-0.0650 (.0410)	0.0627 (.1884)	0.1018 (.0704)
INDU9	-0.1347 (.1119)	-0.0623 (.0456)	0.0260 (.0461)	-0.0850** (.0335)
R^2 adj.	0.3447	0.1906	0.5447	0.3648
SEE	0.3373	0.2752	0.2498	0.1740
F-value	11.48	10.01	13.20	6.07
No. of obs.	519	843	256	204

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980).

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

Table N. Private/public-sector estimates of eqs. (1) and (2) for female employees in, respectively, non-manual and manual jobs using ordinary least squares (OLS) techniques.[1] The dependent variable is log hourly earnings inclusive of fringe benefits.

Variable	PRIVATE-SECTOR WOMEN Non-manuals	Manuals	PUBLIC-SECTOR WOMEN Non-manuals	Manuals
CONSTANT	3.3778** (.1030)	3.2861** (.1604)	3.3496** (.0752)	3.3312** (.1793)
BASIC EDUCATION	0	0	0	0
LOWER VOCATIONAL	-0.0140 (.0351)	-0.0231 (.0344)	0.0368 (.0324)	-0.0031 (.0412)
UPPER VOCATIONAL	0.0841* (.0412)	0.2824** (.1048)	0.1547** (.0332)	0.2217 (.1605)
SHORT NON-UNIV	0.3125** (.0933)		0.3398** (.0362)	
UNDER-GRADUATE	0.4377** (.0715)		0.5005** (.0451)	
GRADUATE	0.4196** (.1077)		0.6548** (.0438)	
EXP	0.0119* (.0061)	0.0131* (.0061)	0.0042 (.0049)	0.0080 (.0109)
EXP^2 /1000	-0.2505* (.1337)	-0.2037 (.1344)	0.0014 (.1190)	-0.0974 (.1868)
SEN	0.0065** (.0016)	-0.0007 (.0024)	0.0108** (.0016)	0.0031 (.0025)
OJT	0.0734** (.0224)	0.0530 (.0353)	-0.0286 (.0215)	-0.0440 (.0294)
MARRIED	-0.0288 (.0280)	0.0260 (.0326)	-0.0457* (.0250)	-0.0066 (.0450)
$CHILD^{0-6}$	0.0499 (.0360)	-0.0041 (.0394)	0.0541* (.0314)	-0.0860 (.0686)
$CHILD^{7-17}$	0.0001 (.0226)	-0.0062 (.0296)	0.0203 (.0223)	0.0024 (.0426)
CAPITAL	0.1084** (.0336)	0.1073** (.0457)	0.0329 (.0250)	0.1361* (.0698)
TEMPEMPL	0.0060 (.0733)	0.1981* (.1006)	0.0870* (.0400)	0.0795 (.0872)
PART-TIME	0.2759** (.0754)	0.3100** (.0891)	0.4881** (.1095)	0.1226 (.0898)
PUBLOCAL			-0.0308 (.0253)	0.0037 (.0391)
PIECE-RATE	-0.0546 (.1263)	0.0587* (.0326)	-0.1874 (.1619)	

Table N. (cont.)

Variable	PRIVATE-SECTOR WOMEN Non-manuals	PRIVATE-SECTOR WOMEN Manuals	PUBLIC-SECTOR WOMEN Non-manuals	PUBLIC-SECTOR WOMEN Manuals
NODAYWORK	0.1187** (.0397)	0.1371** (.0303)	0.1582** (.0318)	0.1457** (.0452)
UNEMPL	-0.0387 (.0461)	-0.0036 (.0562)	-0.0061 (.0645)	0.0076 (.0619)
UNION	-0.0698* (.0328)	-0.0445 (.0429)	0.0683* (.0405)	-0.0760 (.0919)
INDU1-3	0.0518 (.0493)	-0.0386 (.1369)	0.0838 (.1281)	
INDU4	0.0229 (.0704)	-0.0888 (.1703)	-0.0224 (.0675)	
INDU5	0.0405 (.0745)	0.0137 (.1372)	-0.1373** (.0586)	
INDU6	-0.0608 (.0520)	-0.0734 (.1358)	0.2509 (.1985)	
INDU7	0	0	0	
INDU8	0.1315** (.0497)	-0.2508 (.1581)	0.0178 (.0609)	
INDU9	0.0999* (.0593)	-0.0912 (.1441)	-0.0636 (.0515)	
R^2 adj.	0.2488	0.2123	0.4052	0.0865
SEE	0.3132	0.2979	0.2852	0.2715
F-value	9.95	5.97	19.26	2.04
No. of obs.	677	407	698	177

[1] Standard errors are in parentheses below the estimates and are adjusted for heteroscedasticity according to White (1980).

* Denotes significant estimate at a 5 % risk level.

** Denotes significant estimate at a 1 % risk level.

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