GENDER WAGE DIFFERENTIALS IN PRIVATE AND PUBLIC SECTOR JOBS

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Abstract:

In this study gender wage differentials in private and public sector jobs in Austria are calculated. Occupational attainment is considered as endogeneous by the use of an ordered response model. Results show that wage discrimination is also present in the public sector, though on a lower level. Both in private firms and for public servants a substantial part of this unwarranted differential is due to unequal professional advancement.

JEL: J24, J31, J45, J71

1. Introduction

Wages in the public sector have been extensively studied in the past decade; this is motivated by the rising pay-roll for government servants and growing public deficits. Owing to comparability legislation for wage policies in the U.S., research has concentrated on differentials between private and public sector workers (Quinn, 1979, Smith, 1976, Ehrenberg, Schwarz, 1986). This research concluded that men in public sector jobs enjoy a premium in pay over the private sector, but that rents for women are even higher[1] (Gunderson, 1989).

In this study we approach the problem from another point of view: the occurrence of gender wage differentials. Different outcomes in different jobs may be the result of different discrimination practices. In theory, employer discrimination should be eliminated by competition. Accordingly, sex discrimination is found to be higher in concentrated markets (Ashenfelter, Hannan, 1986). In the public sector things are different for several reasons. There is no profit constraint, so that employers are not forced by the market. On the other hand, equal pay and affirmative action policies should apply to the public sector first of all. Moreover, strict pay schemes should make unequal wages for equal jobs nearly impossible.

In looking at gender discrimination in the public sector we also want to concentrate on job assignments. Usually human capital wage functions are augmented by specific job characteristics to account for influences like union status, sectors or occupations. These variables indeed model characteristics of different jobs, but are mostly the result of optimizing behavior of the agents themselves (self-selection, Polachek, 1981), so that no discriminatory practice is possible. The salient point in public sector wage determination is the professional advancement of females compared to males. Decisions about promotions and occupational status are in the sole discretion of the supervisor and might be an easily accessible substitute for (illegal) pay discrimination.

In the next section wage functions with and without the inclusion of occupational status are run and pay gaps are calculated. In section 3 wage differentials are reported for different steps in the job ladder to look at earnings within more narrow jobs. Section 4 analyzes occupational advancement of men and women. Gender wage differentials are split

[1] Shackett and Trapani (1987) find out that wages are highest in private regulated industries.

up into endowment effects and effects due to discrimination both in promotion and in wage determination. The concluding section summarizes the results.

2. The pay gap

For the empirical study we use data from the Austrian microcensus of 1983. Wage determination in Austria is known to be highly centralized. In the private sector only basic wages are fixed, allowing a positive wage drift in firms. In the public sector all workers are covered by collective bargaining;[2] this includes public servants. As only 8 % of governmental employees were blue-collar workers we confined the comparison to white-collar public and private workers[3] aged 20-60, a procedure which is also sensible as we consider career advancement on job ladders and are thus obliged to have a strict hierarchy of job positions.The 1983-survey asked individuals for their job-position, characterized by the skill intensities and/or training requirements. The reported categories are: (i) unskilled (no schooling and training requirements), (ii) low skilled (apprenticeship or equivalent education), (iii) medium skilled (middle school level or equivalent), (ii) high skilled (high school degree or equivalent), (v) leading (university degree or equivalent) and (vi) leading manager in large firms/institutions.

For our remaining sample we receive a gender wage gap of 38% in the private and only 12% in the public sector. In the first place, wage functions following eq. (1) and (2) are run for the 4 subpopulations (male private, female private, male public, female public). In eq. (2) in addition to the set of explanatory variables X_i used in (1) also a vector of dummies for the occupational job positions P_i of the workers is included.

$$(1)\ \ln W_i = aX_i + u_i$$

$$(2)\ \ln W_i = aX_i + bP_i + u_i$$

where $\ln W_i$ denote average hourly earnings measured in logarithmic terms, and i is an individual index. It should be noted, that the error terms of equations (1) and (2) suffer from potential biases. One source of bias comes from the endogeneity of the work

[2]This gives the opportunity to test the effects of collective bargaining on wage differentials in the public sector, which is seldomly possible (Ehrenberg, Schwarz, 1986, p. 1251), for rare exceptions see Asher, Popkin, (1984) or Figart (1987).

[3]Most studies (Gunderson, 1979, Smith, 1976) use manufacturing workers as a control group, which is not suitable in our opinion, as the performed jobs are too diverse.

decision. This concerns only the female samples, since only a negligible proportion of men did not participate in the labor force. Since we restrict the sample to either private or public-sector white-collar workers, there is a second sample-selection problem: occupational choice. This type of selectivity of course concerns men and women alike. To account for these problems we calculated selectivity correction terms from univariate (men) and bivariate (women) probit regressions, respectively.[4] These appear as regressors on the r.h.s of equations (1) and (2).

Control variables, besides the usual human capital proxies, years of education and actual work experience, include agglomeration advantages (city size), marital status and children, weekly working time and regional and industry dummies. In the public sector regressions, special dummies for being a teacher are brought in. Calculated years of work interruptions appear in all equations.

The results in table 1[5] concentrate on eq. (2)[6]. The explanatory power is higher in the public sector because of stricter salary scales. All coefficients - except those for work interruptions and foreigners - show the expected sign. Work interruptions are calculated as "age - minimal years of schooling for a certain degree - 6 - actual experience". This indicator has to be interpreted with care because of imprecise measurement. It should be a sum of true interruptions with negative wage effects and spurious ones, i.e. years of job experience or additional education, with a positive wage impact. This should be especially important for men where true interruptions are supposed to occur more seldomly. In addition, due to collective bargaining agreements in the public sector, women enjoy pay advancements on a reduced scale even during periods of maternity leave. The share of foreigners in the sample is very small, ranging from 2% for men in the private sector to 0.3% for men in the public sector, which may mainly consist of academics. Teachers receive a bonus of 4-5% which may be due to their shorter hours of work per week. With one exception sample selectivity bias does not seem to be a serious problem. For women, neither the selectivity variable for participation, nor for sectoral choice indicates a significant sample selection problem. The same is true for male workers in the public sector. However, it seems to be of some importance in the case of male private sector employees.

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[4]Bivariate probit analyses for women takes into account that observation of occupational choice is contingent upon participation in the labor force.

[5]T-values are based on the correct asymptotic covariance matrix for the two step estimator (Greene, 1992, p. 605f.).

[6]Results of the regressions without occupational dummies are presented in the appendix.

Table 1: Wage functions (t-values in parentheses)

	Private sector		Public sector	
	Males	Females	Males	Females
Constant	4.341 (40.3)	3.902 (37.8)	4.309 (18.2)	4.180 (2.5)
Years of Schooling	0.010 (2.3)	0.025 (3.4)	0.021 (2.4)	0.030 (0.5)
Apprentice training	0.002 (0.1)	-0.010 (0.6)	0.014 (0.8)	-0.040 (0.4)
Experience	0.028 (12.0)	0.018 (7.2)	0.021 (7.5)	0.019 (1.7)
Experience^2	-0.0004 (8.4)	-0.0002 (3.2)	-0.0002 (3.8)	-0.0002 (1.5)
Work interruptions	0.008 (5.8)	0.003 (3.0)	0.008 (4.9)	0.008 (1.8)
City size (<2000)				
2-1000	-0.004 (0.2)	0.016 (0.9)	-0.001 (0.1)	0.035 (0.4)
10-100 000	0.015 (0.6)	0.083 (3.7)	0.023 (1.0)	0.064 (0.7)
> 100 000	0.034 (1.5)	0.098 (4.8)	0.073 (3.0)	0.004 (0.1)
Married	0.102 (5.3)	-0.119 (2.1)	0.026 (1.0)	0.130 (0.3)
Children <4 years	0.048 (7.0)	0.014 (1.6)	0.076 (10.6)	0.035 (0.5)
Weekly working time	-0.015 (13.2)	-0.010 (11.8)	-0.017 (11.8)	-0.026 (16.0)
Region (Swiss border)				
West	-0.080 (3.4)	-0.080 (2.9)	-0.076 (2.3)	-0.050 (0.6)
Central	-0.109 (5.0)	-0.103 (4.1)	-0.046 (1.5)	-0.013 (0.2)
East	-0.096 (4.2)	-0.039 (1.4)	-0.052 (1.6)	0.002 (0.3)
Foreigner	0.088 (1.8)	0.019 (0.4)	0.288 (1.9)	0.061 (0.3)
Occupation (Seasonal)				
Production	0.104 (2.9)	-0.060 (1.0)	-	-
Service	0.108 (3.5)	-0.035 (0.8)	-	-
Industry (Manufacturing)				
Mining, Construction	0.033 (1.3)	-0.034 (0.8)	-	-
Trade, Hotels	-0.018 (1.0)	-0.068 (3.6)	-	-
Transport, Energy	-0.030 (1.3)	-0.038 (1.1)	-	-
Finance, Insur.	0.011 (0.6)	-0.006 (0.3)	-	-
Other service	-0.089 (4.6)	0.008 (0.4)	-	-
Teacher	-	-	0.053 (2.2)	0.041 (1.3)
Professional position (unskilled)				
Low skilled	0.070 (3.3)	0.115 (5.5)	0.096 (2.3)	0.121 (2.0)
Medium skilled	0.188 (9.3)	0.254 (11.7)	0.239 (5.9)	0.358 (6.2)
High skilled	0.291 (12.9)	0.314 (9.7)	0.313 (7.4)	0.393 (5.0)
Leading	0.453 (15.2)	0.573 (9.3)	0.490 (10.4)	0.578 (6.2)
Leading manager	0.594 (18.2)	0.767 (6.6)	0.608 (8.3)	-
λ_1	-0.076 (2.4)	-0.008 (0.2)	-0.084 (1.1)	0.149 (0.5)
λ_2	-	0.015 (0.3)	-	-0.167 (0.3)
$\bar{R}^2$	0.542	0.361	0.516	0.671
SEE	0.263	0.271	0.267	0.262
N	1849	2053	1429	561

The general picture yields an experience profile flatter than that of other countries (Wagner, 1990); rewards for experience differ by gender in the private but not in the public sector. The diverging influence of marriage on pay in the private sector is consistent with Becker's (1985) argument that women spend less effort on each hour of market work owing to the changing strain of housework following marriage and the setup of a common household. As we use after-tax hourly incomes, progressive taxation is reflected in the working time variable. As expected, occupational positions raise the explanatory power of the wage functions by a large amount, on average by 0.083 percentage points.

Following the well-known discrimination decomposition of Blinder (1973) and Oaxaca (1973) we can split up wage differentials in the following way ($\overline{X}$, $\overline{P}$ denoting mean values):

$$(3)\quad \ln W_m - \ln W_f = (a_m\overline{X}_m + b_m\overline{P}_m) - (a_f\overline{X}_f + b_f\overline{P}_f) =$$
$$= \left[(\overline{X}_m - \overline{X}_f)a_f + (\overline{P}_m - \overline{P}_f)b_f\right] \quad \text{.....endowments I}$$
$$+ \left[(a_m - a_f)\overline{X}_m + (b_m - b_f)\overline{P}_m\right] \quad \text{... discrimination I}$$

Table 2: Decomposition of wage differentials

	Private		Public	
	B-O-I*	B-O-II**	B-O-I	B-O-II
Excluding occ.positions				
- Endowments	-0.015	0.011	-0.071	-0.050
- Discrimination	0.383	0.357	0.189	0.168
Including occ. positions				
- Endowments	0.051	0.074	-0.084	0.019
- Discrimination	0.317	0.294	0.202	0.099
Actual Differentials	0.368	0.368	0.118	0.118

* Blinder-Oaxaca Measure I endowment differences evaluated at female coefficients.

** Blinder-Oaxaca Measure II: endowment differences evaluated at male coefficients.

If occupational positions are included in the wage functions, a discrimination component of 30% for white-collar workers is on the upper bound of international (Gunderson, 1989, Cain, 1986) and former Austrian (Christl, 1985) evidence. In the public sector where gross wage differentials are smaller, nonetheless high discrimination components are found. Females in the public sector are generally better educated than men; endowing them with male productivity characteristics will lead to even lower wages. Comparing the two Blinder-Oaxaca measures for the decomposition of differentials, we find surprisingly small differences for private white-collar workers, which makes a separate Neumark (1985) decomposition unnecessary. In the public sector results are a bit unclear. Excluding occupational status the results are fairly stable, which is not the case in the lower part of the panel: unexplained differentials range from 10-20%, if occupational positions are included in the wage regressions.

We proceed in two directions to shed further light on this issue. i) By looking at wage differentials within occupational positions we can analyse discrimination patterns on different levels of the job ladder; the groups under consideration become also more homogenous in this respect. ii) In a second step (chapter 4) the process of moving up the job ladder is modelled separately to show differences in occupational attainment and the consequences for wages.

3. Wage differentials by occupational rank

Following the specifications in Table 1, separate wage regressions by sex and occupational status are run. Owing to the low number of females in upper positions, several categories are lumped together.

Table 3: Decomposition* of wage differentials by occupational rank

Private sector

Rank	Wage differential	Endowment I	Discrimination I	Endowment II	Discrimination II
0	0.330	0.053	0.277	-0.005	0.335
1	0.285	0.057	0.228	-0.004	0.289
2	0.246	0.052	0.194	-0.004	0.250
3	0.243	0.048	0.195	-0.051	0.297
4,5**	0.174	0.010	0.164	-0.097	0.271

Public Sector

Rank	Wage differential	Endowment I	Discrimination I	Endowment II	Discrimination II
0	0.220	.008	0.212	-0.098	0.318
1	0.210	.054	0.156	-0.142	0.352
2	0.133	.084	0.049	-0.110	0.243
3, 4,5**	0.141	.018	0.123	-0.051	0.192

* Control variables as in Table 1 without selection and rank terms. Occasionally dummy variables had to be eliminated because of zero cell density.
**Lumped together owing to small number of females for the separate wage regressions.

Raw wage differentials - shown in Table 3 - are decreasing in both sectors. The comprehended upper rank groups must be taken with care, because they comprise different male/female ratios in the respective positions. In the private sector the differential falls to 0.09 in position 4, rising again in the top management group 5 to 0.264. In the public sector upper pay bounds reduce differentials to 0.088 in rank 3 and 0.069 in rank 4 (not reported in Table 3).

Different endowments cannot explain different performance in any sub-group, the unexplained residuals are large. Following the raw differential pattern, discrimination coefficients decline slightly with rising skill level. This confirms results for Germany in certain respects: Brandes et al (1989) also describe discrimination highest in low skill groups, but they did not find such a regularly decreasing pattern as we did for Austria.

4. Advancement in occupational positions

Promotion into better positions often incurs training costs. Besides productivity, the profitability of this investment depends upon the expected future job tenure of the candidate. If women have a shorter time horizon in market work because of the comparative advantage in outside opportunities, a profit maximizing firm will require higher qualification standards for promotion in the case of women compared to men. (Lazear and Rosen, 1990). Consider a situation where men and women are equally productive in labor market activities, but differ in their quit propensities. It is natural to assume that the probability of separation is higher for women than for men, if women have a comparative advantage in outside (non-market) opportunities. Suppose the firm cannot observe the non-market productivity of the individuals, but has some idea about their

distribution. If the distribution of men's non-market skills stochastically dominates the distribution of those of women, firms will demand strictly higher promotion requirements for women.

Similar arguments are used in the context of dual labor markets (Bulow and Summers, 1986). Also in this setting, productivity differentials between men and women are absent. However, the shorter time horizon of women in primary sector jobs would require the firms to pay higher wages for women to induce them not to shirk. As a consequence men will enter the good jobs, where women will be crowded in the secondary sector.

It is obvious that the crucial variable in assigning equally productive workers to different jobs is the workers' separation probability. We will return to this issue below. In 1983 Austrian white-collar and public-sector workers were distributed in the following way:

Table 4: Actual rank distribution (%)[7]

	0	1	2	3	4	5
Private sector						
Males	14.0	21.4	27.5	20.0	11.1	6.0
Females	14.1	37.0	37.5	8.7	2.4	0.3
Public sector						
Males	3.2	20.6	35.6	27.8	11.4	1.4
Females	9.6	12.5	24.5	49.0	4.4	0.0

Whereas men are clearly overrepresented in the upper ranks of the private sector, no such simple diagnosis is possible for public servants, due to the crowding of female teachers in position 3. Since six ranked career positions are observed, the advancement model leads empirically to an ordered response specification, which can be written as follows. Assume that actual abilities of the workers can be represented by

$$(4)\ \delta_i = \beta Z_i + \varepsilon_i.$$

Let μ_j be the productivity threshold to be overcome for promotion from job level (j-l) to level j. The probability of being in state j of the job ladder is given by

[7]Rank assignment in public and private sector is not directly comparable.

$$(5)\ P_j = \Pr(\beta Z_i + \varepsilon_i > \mu_j) - \Pr(\beta Z_i + \varepsilon_i > \mu_{j+1})$$
$$= G(\mu_{j+1} - \beta Z_i) - G(\mu_j - \beta Z_i).$$

G is the distribution function of ε_i. Let us take a closer look at the ability thresholds, which we assume to be determined differently for men and women. To account for females' lower work attachment we included years of past work interruptions h and expected future fertility risk α as variables affecting women's thresholds, work interruptions h are included for men as well. In particular, we proxy women's threshold $\mu_{j,f}$ by the following relation

$$(6)\ \mu_{j,f} = \mu_j + \phi h_f + \alpha_{lk}$$
$$\text{with } \alpha_{lk} = 1 - \prod_{n=1}^{5} (1 - \Pr(\text{birth} \mid \text{age } l+n,\ \text{number of children } k).$$

α_{lk} gives the probability that a women of age 1 having k children will bear a further child within the next five years.[8]

Assuming standard normality for ε_i gives rise to an ordered probit model[9] which estimates the interesting parameters ß und μ_j by maximum likelihood techniques. However, using score tests described by Machin, Stewart (1990), it turned out that the normality assumption is rejected by the data. As the errors are found to deviate in curtosis from a normal distribution, we choose a logistic density, which has a higher peak but declines faster than a normal density (Cramer, 1991, p.15). We thus receive an ordered logit model.

The model yields a fairly good fit as can be seen by the likelihood-ratio test and the percentage of right predictions. The Akaike Information Criterion (AIC) is positive in all cases, i.e. the logit specification yields a better approximation of the data than the probit model. The thresholds μ_j are well determined and significantly different from each other, signifying that the classification in six distinct and ranked categories makes sense. For white-collar workers inferior promotion possibilities for women can be deducted from the

[8]For calculation we employed data collected by the Austrian Statistical Office (Demographisches Jahrbuch).

[9]Brown, et al (1980) and Miller (1987) study occupational attainment in a similar context, without having a clear ranking of positions at their disposal.

Table 5: Estimates of ordered logit model for professional status (t-values in parentheses)

	Private sector		Public sector	
	Males	Females	Males	Females
Constant	-6.285 (11.8)	-6.247 (11.2)	-4.656 (6.5)	-4.138 (3.4)
Years of Schooling	0.609 (23.9)	0.748 (27.8)	0.627 (20.4)	0.658 (12.0)
Apprentice training	0.450 (4.6)	0.623 (6.7)	0.261 (2.2)	0.402 (1.8)
Experience	0.052 (3.1)	0.044 (2.5)	0.027 (1.4)	-0.012 (0.3)
Experience^2	-0.0008 (2.3)	-0.0009 (2.2)	0.0001 (0.3)	0.0006 (0.6)
Work interruptions	0.029 (3.1)	-0.018 (2.4)	0.015 (1.3)	-0.030 (1.7)
City size (<2000)				
2-1000	0.256 (2.1)	0.189 (1.5)	0.063 (0.4)	-0.352 (1.1)
10-100 000	0.309 (2.1)	0.147 (1.0)	0.200 (1.2)	-0.705 (2.0)
> 100 000	0.369 (2.7)	0.392 (2.0)	-0.018 (0.1)	-0.813 (2.7)
Married	0.235 (1.8)	-0.111 (1.1)	0.208 (1.3)	-0.134 (0.6)
Children <4 years	0.140 (2.9)	0.027 (0.5)	0.060 (1.1)	-0.061 (0.5)
Weekly working time	0.017 (2.4)	0.011 (1.9)	0.026 (2.0)	0.011 (0.6)
Region (Swiss border)				
West	0.277 (1.7)	0.101 (0.5)	0.504 (2.1)	0.182 (0.3)
Central	0.343 (2.3)	0.059 (0.3)	0.254 (1.1)	0.491 (1.0)
East	0.358 (2.3)	-0.039 (0.2)	0.386 (1.6)	0.114 (0.2)
Foreigner	-0.185 (0.5)	-1.018 (2.8)	1.030 (1.7)	-0.238 (0.4)
Occupation (Seasonal)				
Production	0.283 (1.0)	0.211 (0.6)	-	-
Service	0.653 (2.8)	0.833 (3.3)	-	-
Industry (Manufacturing)				
Mining, Construction	0.185 (1.0)	-0.045 (0.2)	-	-
Trade, Hotels	-0.779 (6.1)	-0.841 (6.4)	-	-
Transport, Energy	-0.681 (4.3)	-0.222 (1.0)	-	-
Finance, Insurance	-0.009 (0.7)	0.207 (1.7)	-	-
Other service	-0.709 (9.1)	0.162 (1.3)	-	-
Teacher	-	-	0.380 (2.0)	1.870 (7.2)
Fertility index α_{1k}	-	-0.716 (2.3)	-	-1.511 (1.8)
μ_1	1.454 (21.6)	2.367 (30.8)	2.441 (16.0)	1.395 (8.5)
μ_2	3.069 (34.6)	5.228 (44.3)	4.652 (27.3)	3.606 (15.0)
μ_3	4.841 (40.0)	7.602 (34.3)	7.351 (35.4)	9.199 (22.4)
μ_4	6.630(45.8)	10.570 (25.7)	10.374 (31.9)	-
Log L	-2689.3	-2393.0	-1682.7	-516.30
LR-Test	1074.6	1063.30	879.6	469.92
AIC*	15.7	21.0	6.2	5.92
% right predictions	37.3	54.2	50.9	62.5
N	1903	2199	1448	571

* Akaike Information Criterion, Logit vs. Probit:

$\frac{1}{2}(\text{Log } L_L - M_L)$ - $\frac{1}{2}$ $(\text{Log } L_p - M_P)$, $M_{(.)}$ ... number of independent variables in Logit or Probit model

much higher ability standards (μ_j) they must meet. In the case of public servants the picture is more irregular: women climb up the first steps of the job ladder more easily than men, but suffer from a "career stop" in middle-management positions.[10] Schooling as well as training have a high influence on occupational attainment. In the public sector, experience is not rewarded in the same manner as in private sector jobs, especially for women. As predicted from theory, variables catching the shorter horizon in market work of women (work interruptions and fertility risk) reduce women's advancement.[11]

Furthermore, the detrimental impact of expected fertility risk on occupational attainment in the private sector is lower with higher education and job experience.[12] Two explanations for this phenomenon are possible: i) the desire for having children might be lower for those groups, or ii) advancement consequences are not so severe. Some experimentation with aggregate quit rates - on an occupational or industry level - turned out unsuccessful both for men and women. Coefficients were never significant and often changed signs.

Comparing the results of table 5 with the wage functions helps to clarify some peculiarities. Periods of home time lead to a career setback for women, but besides that no separate negative impact on wages can be detected. Higher rewards for females from superior hierarchical positions in the wage functions are consistent with different advancement standards by gender: women have to do better than men to be promoted, following human capital theory they get higher money payoffs for achieved advancements.

Extending the decomposition of wage differentials in section 2, we proceed to incorporate effects of unequal promotion schemes. In wage functions where occupational positions are included, these positions cannot be taken as exogenously given. Instead, occupational attainment can be decomposed in effects resulting from endowments and discrimination in the usual way, using results from Table 5. Finally, we arrive at a decomposition of raw wage differentials arising from four sources: endowments vs. discrimination in wage determination as such and wage effects of endowments vs. discrimination originating from

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[10]Teachers are better graded from the very beginning.

[11]See Winter-Ebmer, Zweimüller (1991) for a more thorough discussion of the issue of efficient vs. discriminatory promotion in the context of expected length of stay on the job.

[12]The equation with interaction terms read as follows (t-values in par.):

$$-3.21\,\alpha_{lk} + 0.17\ \text{school} * \alpha_{lk} + 0.16\ \exp * \alpha_{lk}.$$
$$(2.7) \qquad (1.6) \qquad (2.9)$$

the occupational promotion process (following eqs. 3 and 5 with $\overline{G}$ as average over individuals: $(\overline{G} = \frac{1}{N}\sum_{i=1}^{N} G(...))$.[13]

$$
\begin{aligned}
7)\ \ln W_m - \ln W_f &= (a_m \overline{X}_m + b_m \overline{P}_m) - (a_f \overline{X}_f + b_f \overline{P}_f) \\
&= \left[(a_m - a_f)\overline{X}_m + (b_m - b_f)\overline{P}_m\right] \quad \text{..... discrimination in wages} \\
&+ \left[(\overline{X}_m - \overline{X}_f)a_f\right] \quad \text{..... endowments in wages} \\
&+ \left[\overline{G}(\mu_{j+1,m} - \beta_m Z_m) - \overline{G}(\mu_{j,m} - \beta_m Z_m) - \overline{G}(\mu_{j+1,m} - \beta_f Z_m) + \overline{G}(\mu_{j,f} - \beta_f Z_m)\right] b_f \quad \text{... discrimination in promotion} \\
&+ \left[\overline{G}(\mu_{j+1,f} - \beta_f Z_m) - \overline{G}(\mu_{j,f} - \beta_f Z_m) - \overline{G}(\mu_{j+1,f} - \beta_f Z_f) + \overline{G}(\mu_{j,f} - \beta_f Z_f)\right] b_f \quad \text{... endowments in promotion}
\end{aligned}
$$

Tab.6: Decomposition of wage differentials

	Private		Public	
	B-O-I	B-O-II	B-O-I	B-O-II
Wage Determination				
Endowments	-.035	-.002	-.099	.008
Discrimination	.317	.294	.202	.099
Promotion Effects				
Endowments	.024	.037	-.041	-.009
Discrimination	.062	.039	.056	.020
Total Differential	.368	.368	.118	.118

In the private sector, 4-6 percentage points of the gender wage differential comes from discrimination in professional advancement.[14] A similar result shows up in the public sector. Here the range in the estimates by the two methods of decomposition is slightly higher: from 2.0 percentage points (B-O-II) to 5.6 points (B-O-I).

The overall picture now shows an even higher part of discrimination. Both in the private

[13]Due to the non-linear character of the ordered logit model the use of variable means as in eq. 3 is not possible.

[14]This proportion corresponds to results for the private sector in the UK by Miller (1978) and Dolton, Kidd (1990).

and in the public sector endowment differences can account at most for 3.5 points (= -.002 + .037) in the case of white-collar private-sector workers (B-O-II). In the public sector the endowment differential either is zero (when females' endowments are rewarded like males', B-O-II) or indicates that women would on average deserve higher pay than their male collegues (B-O-I). The discrimination component ranges from .333-.379 in the private, and from .119-.258 in the public sector.

5. Conclusions

The overall picture of our results show that wage discrimination is very high in the private sector, but also astonishingly high in the public sector. In both the private and the public sector an important part of this differential is due to unequal promotion procedures. Whereas in the private sector women are crowded in lower ranks in the job hierarchy, female employees in the public sector suffer from a career stop in middle management positions.

The results of this study have to be treated with caution, especially those for the public sector because tenure in the respective institution could not be observed. Policies of equal treatment seem to be effective in the public administration although wage discrimination still exists and leading positions remain reserved for men.

Appendix: Wage functions without professional position (t-values in parentheses)

	Private sector		Public sector	
	Males	Females	Males	Females
Constant	4.027 (34.6)	3.629 (33.8)	4.147 (16.8)	4.227 (2.9)
Years of Schooling	0.048 (11.0)	0.061 (9.6)	0.051 (5.6)	0.053 (0.9)
Apprentice training	0.021 (1.3)	0.016 (1.0)	0.029 (1.6)	-0.016 (0.2)
Experience	0.032 (12.2)	0.021 (7.6)	0.021 (7.3)	0.019 (2.0)
Experience^2	-0.0005 (9.3)	-0.0002 (3.4)	-0.0002 (3.3)	-0.0002 (1.4)
Work interruptions	0.011 (7.0)	0.0026 (2.6)	0.009 (5.0)	0.005 (1.2)
City size (<2000)				
2-1000	0.004 (0.2)	0.028 (1.4)	-0.002 (0.1)	0.031 (0.4)
10-100 000	0.021 (0.8)	0.091 (3.9)	0.028 (1.0)	0.045 (0.5)
> 100 000	0.042 (1.6)	0.119 (5.2)	0.069 (2.7)	-0.011 (0.1)
Married	0.107 (5.0)	-0.131 (2.2)	0.037 (1.3)	0.076 (0.3)
Children <4 years	0.062 (8.1)	0.017 (1.8)	0.078 (10.3)	0.036 (0.6)
Weekly working time	-0.014 (10.8)	-0.009 (10.3)	-0.015 (10.1)	-0.024 (14.6)
Region (Swiss border)				
West	-0.060 (2.3)	-0.073 (2.5)	-0.055 (1.5)	-0.065 (0.8)
Central	-0.082 (3.5)	-0.098 (3.6)	-0.035 (1.1)	-0.0003 (0.1)
East	-0.069 (2.8)	-0.042 (1.5)	-0.036 (1.1)	-0.012 (0.2)
Foreigner	0.083 (1.5)	-0.012 (0.2)	0.332 (2.1)	0.045 (0.3)
Occupation (Seasonal)				
Production	0.121 (3.0)	-0.043 (0.7)	-	-
Service	0.160 (4.7)	0.013 (0.3)	-	-
Industry (Manufacturing)				
Mining, Construction	0.044 (1.6)	-0.030 (0.8)	-	-
Trade, Hotels	-0.069 (3.6)	-0.104 (5.3)	-	-
Transport, Energy	-0.083 (3.3)	-0.042 (1.2)	-	-
Finance, Insur.	-0.004 (0.2)	-0.0008 (0.1)	-	-
Other service	-0.139 (6.6)	0.013 (0.7)	-	-
Teacher	-	-	0.056 (2.2)	0.104 (3.4)
λ_1	-0.114 (3.1)	-0.013 (0.3)	-0.089 (1.1)	0.129 (0.5)
λ_2	-	0.021 (0.4)	-	-0.155 (0.4)
$\bar{R}^2$	0.437	0.282	0.448	0.592
SEE	0.298	0.287	0.286	0.278
N	1849	2053	1429	561

Literature

Ashenfelter, Orley, Hannan, Timothy: Sex Discrimination and Product Market Competition: The Case of the Banking Industry, The Quarterly Journal of Economics, 1986, 149-174.

Asher, Martin, Popkin, Joel: The Effect of Gender and Race Differentials on Public-Private Wage Comparisons: A Study of Postal Workers, Industrial and Labor Relations Review, 1984, 16-25.

Becker, Gary S.: Human Capital, Effort and the Sexual Division of Labor, Journal of Labor Economics, 1985, 533-58.

Blinder, Alan S.: Wage Discrimination: Reduced Form and Structural Estimates, Journal of Human Resources, 1973, 436-455.

Brandes, Wolfgang et al: Eingruppierungsunterschiede von Frauen und Männern beim Staat als Arbeitgeber, Mitteilungen aus der Arbeitsmarkt- und Berufsforschung, 1989, 245-276.

Brown, Randall S., Moon, Marilyn, Zoloth, Barbara S.: Incorporating Occupational Attainment in Studies of Male-Female Wage Differentials, Journal of Human Resources, 1980, 3-28.

Bulow, Jeremy I., Summers, Lawrence, H.: A Theory of Dual Labor Markets with Application to Industrial Policy, Discrimination, and Keynesian Unemployment, Journal of Labor Economics, 1986, 376-414.

Cain, Glen G.: The Economic Analysis of Labor Market Discrimination: A Survey, in: Ashenfelter, Orley, Layard, Richard (Eds.), Handbook of Labor Economics, (North-Holland), 1986, 693-786.

Christl, Josef: Verdienstdifferentiale zwischen Männern und Frauen, IHS, Vienna, mimeo, 1985.

Cramer, J. S.: The Logit Model, E. Arnold (London), 1991.

Demographisches Jahrbuch der Republik Österreich, Statistisches Zentralamt, Vienna, 1984.

Dolton, Peter J., Kidd, Michael P.: Occupational Access and Wage Discrimination, mimeo, University of Newcastle, 1990 .

Ehrenberg, Ronald G., Schwarz, Joshna L.: Public Sector Labor Markets, in: Ashenfelter, Orley, Layard, Richard (Eds.), Handbook of Labor Economics, (North-Holland), 1986, 1219-1268.

Figart, Deborah M.: Gender, Unions, and Internal Labor Markets: Evidence from the Public Sector in Two States, American Economic Review, 1987, 252-256 .

Greene, William H.: Limdep User's Manual, 1992, NY.

Gunderson, Morley: Earnings Differentials between the Public and Private Sectors, Canadian Journal of Economics, 1979, 228-242 .

Gunderson, Morley: Male-Female Wage Differentials and Policy Responses, Journal of Economic Literature, 1989, 46-72.

Lazear, Edward P., Rosen, Sherwin: Male-Female Wage Differentials in Job Ladders, Journal of Labour Economics, 1990, 106-123.

Machin, Stephen J., Stewart, Mark B.: Unions and the Financial Performance of British Private Sector Establishments, Journal of Applied Econometrics, 1990, 327-350.

Miller, Paul W.: The Wage Effect of the Occupational Segregation of Women in Britain, The Economic Journal, 1987, 885-896 .

Neumark, David: Employers' Discriminatory Behavior and the Estimation of Wage Discrimination, Journal of Human Resources, 1985, 279-295.

Oaxaca, Ronald: Male-Female Wage Differentials in Urban Labor Markets, International Economic Review, 1973, 693-709.

Polachek, Solomon W.: Occupational Self-Selection: A Human Capital Approach to Sex Differences in Occupational Structure, The Review of Economics and Statistics, 1981, 60-69 .

Quinn, Joseph F.: Wage Differentials Among Older Workers in the Public and Private Sectors, Journal of Human Resources, 1979, 41-62.

Shackett, Joyce R., Trapani, John M.: Earnings Differentials and Market Structure, Journal of Human Resources, 1987, 518-531.

Smith, Sharon: Government Wage Differentials by Sex, Journal of Human Resources, 1976, 185-199.

Wagner, Joachim: An International Comparison of Sector Wage Differentials, Economics Letters, 1990, 93-97.

Winter-Ebmer, Rudolf, Zweimüller, Josef: Unequal Promotion in Job Ladders, University of Linz, WP 9112, 1991.

Winter-Ebmer, Rudolf, Zweimüller, Josef: Occupational Segregation and Career Advancement, Economics Letters, 1992, 229-234.

Zweimüller, Josef: Survey Non-Response and Biases in Wage Regressions, Economics Letters, 1992, 105-109.