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Public–private sector segmentation in the Pakistani labour market

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ABSTRACT

This study investigates public–private sector wage differentials for male and female waged employees in Pakistan. This is done using latest nationally representative data from the Pakistan Living Standards Measurement Survey (PSLM) 2005. We adopt three methodologies to obtain robust estimates of the wage differential and the results reveal that public sector workers enjoy large wage premia. The gross pro-public wage differential is much larger for women than for men. Our findings also show that while private and public sector workers' differing characteristics 'explain' a larger proportion of the private–public wage gap for men, this is not the case for women.

1. Introduction

The Pakistani labour market is characterised by very large wage gaps between public and private sector workers. In 2004–2005, men in government jobs earned on average Rs. 8211/month ($137)[1] and in private jobs only Rs. 5420 ($90), a private:public ratio of 1:1.5. This raw public–private wage difference is even higher for women, the private:public wage ratio being 1:3, i.e. Rs. 6614/month ($110) in public jobs and Rs. 2160/month ($36) in private sector jobs[2]. However, raw differences in earnings can be misleading if the public sector employs individuals with superior observed or unobserved characteristics (unobserved characteristics are those that are observed by employers but are unavailable to researchers). To obtain estimates of wage differentials between similar public and private sector employees, i.e. to get a sense of the true extent of private–public wage gap, it is important to control for worker characteristics.

This paper asks whether there is discrimination in the Pakistani labour market as between the public and private sectors and investigates whether the extent of private–public differentiation is equally large for men and women. To do this, it investigates the determinants of public and private sector wages and examines public–private wage differentials, drawing on the latest nationally representative household dataset: the Pakistan Living Standards Measurement Survey (henceforth PSLM, 2005). The paper therefore does the following: (1) estimates earnings functions by employment sector; (2) decomposes the public–private wage differentials into 'explained' and 'unexplained' components; and (3) does the above

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[1] The Exchange rate used everywhere in this study is the historical rate as at January 2005 of 1USD = Rs. 59.57.
[2] Wage figures from the Pakistan Living Standards Measurement Survey (2005).

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separately for male and female wage employees. Gender differentiated analysis is especially important as the wage determination process and labour market differentiation by sector may be different for men and women.

To our knowledge, this is the first study in Pakistan to look at these issues separately for male and female wage employees. Much of the interest in this debate in Pakistan is recent and is covered in three studies. One of these finds very small public sector wage premia over formal private sector jobs (Nasir, 2000) while the other two (Hyder, 2007; Hyder & Reilly, 2005) report finding quite substantial wage differentials across the two sectors. The finding of a large public–private wage differential is not uncommon in other developing countries. For instance, a study in India finds that the public-sector wage premium ranges from 62 to 102% over private-formal sectors and between 164 and 259% over private-informal sector jobs (Glinskaya & Lokshin, 2005). The narrower public–private pay gaps reported in the former study are linked to the somewhat arbitrary exclusion of informal sector workers[3].

The present study sheds light on an important dimension missing from both previous studies—that of gender. It is useful to study whether the extent of private–public wage gap differs by gender because existing studies suggest that the Pakistani labour market is characterised by exceedingly different outcomes for men and women in other respects such as labour force participation, occupational attainment and conditional earnings (Aslam, 2007a,b; Kingdon & Soderbom, 2007).

The empirical strategy used here relies on three main methodologies: (1) Ordinary Least Squares (OLS); (2) sample-selectivity corrected estimates, (using multinomial logits to control for labour force participation and for private–public sector choice, conditional on wage employment); and (3) household fixed-effects. The last (household fixed-effects) approach is a stringent control for worker unobservables that commonly generate biases in parameter estimates[4]. Finally, the preferred estimates are used to decompose the public–private wage gaps using the familiar Oaxaca's methodology, separately for males and females.

Several explanations have been proposed for the existence of the unexplained portions of public–private wage differentials. These include supply demand models, vote-maximisation models, human-capital models, segmented labour markets, rents, and bargaining models among others (Bender, 1998, pp. 178). Gunderson (1978, 1979) suggests that the main difference between the two sectors is that while the public sector budget faces a political constraint by voters, the private sector is characterised by profit maximisation. In this case, the public-sector wage floor is set by market forces and that public sector employers justify higher pays as their employees are engaged in vote-producing activities. Bender (1998), summarising literature on wage-differentials from developed and developing countries, argues that a large part of the differential between the two sectors can be attributed to the role of trade unions[5]. However, it must be noted that the evidence on wage-differentials is based on equations using different econometric specifications and the findings can be quite sensitive to the choice of methodology adopted.

Past evidence investigating sectoral wage-differentials in developing countries suggests that wage gaps are negative and often large, sometimes in the region of 500% (Bender, 1998, pp. 213). In terms of differences by gender across the two sectors, much work stems from developed countries where the findings reveal large differences in the pay determination process by gender. In some studies public–private differentials are *larger* for women compared to men (see for instance Gunderson, 1979; Hou, 1993 cited in Bender, 1998; Venti, 1987). The latter finding is sometimes seen to suggest that there must be *less* gender wage discrimination in the public sector (Smith, 1976 cited in Bender, 1998).

The public-sector in Pakistan is marked by wage compression and, hence, is similar to that in many other developing countries. Wages in the public sector are largely determined through the political process rather than on the basis of productivity (Ali, 1998; Nasir, 2000). As in other countries, the government is also a popular employer because it offers permanency, job flexibility and fringe benefits often not available to poorly protected private sector employees. Given the poor incentive mechanisms prevailing in this sector and also because of a lack of supervision and ineffective monitoring, the public sector is marked by frequent job-shirking and absenteeism among employees (Ali, 1998). Somewhat perversely, however, these features also make government jobs especially attractive to women in view of their family commitments.

The government of Pakistan is a large employer, employing individuals across a range of occupations, skills-sets and personal characteristics. The public sector absorbs almost 30% of wage earners aged 15–65 in the sample used in this study while the private waged sector employs the remaining 70% Pakistan Living Standards Measurement Survey 2005 (PSLM 2005). The majority of the latter are in 'unprotected' private sector waged jobs[6]. The wages of public sector employees in Pakistan are formulated under the Regulations Wing of the Finance Division. The government's 'Basic Pay Scale' consists of a basic pay and various allowances with a small annual increment built into the pay scale. The basic pay and allowances all vary by grade and department (such as health, education, etc.). For instance, a primary school teacher is hired in grade 7 and her basic pay would be between Rs. 2555/month ($43) to Rs. 6755/month ($113) with an annual increment of Rs. 140 ($2). A

[3] We are grateful to a referee of this journal for pointing this out.

[4] Because of small sample sizes, we are constrained to estimating these models only on sub-samples of male wage earners.

[5] Whether trade unions generate the large public–private wage differentials in Pakistan has not been formally studied largely due to the paucity of data. However, one wonders whether trade unions can play any role, let alone an effective role, in collective bargaining to generate these differentials between the two sectors in Pakistan. This is because while the Industrial Relations Ordinance (IRO, 1969) allows trade union activity, in reality collective bargaining activities are largely suppressed through various other legislation and ordinances. For instance, workers in hospitals and in civil and defence services cannot form unions. Where workers are allowed to form unions, they cannot carry out strike actions and the new IRO passed in 2002 allows the government to effectively end any trade union strike that has lasted more than 15 days (Human Development in South Asia, The Employment Challenge, 2003, pp. 113).

[6] (US Department of Labour at http://www.dol.gov/ilab/media/reports/oiea/wagestudy/FS-Pakistan.htm and http://en.wikipedia.org/wiki).

medical doctor, on the other hand, would be hired in grade 17 with basic pay ranging from Rs. 7140/month ($120) to Rs. 17,840/month ($299) and an increment of Rs. 535 ($9) annually (Office Memorandum, GOP, 2005).

Wages of unskilled workers are set by the government through the National Minimum Wage Commission. These laws are, however, applicable only to industrial and commercial establishments employing 50 workers or more. These laws also cover workers in the formal sector, leaving the large informal sectors and the agricultural sector workers unprotected. While the minimum wage was Rs. 1500/month ($25 at $1 = Rs. 60) when first instituted in 1992, it has been raised to Rs. 4000/month (about $67) in 2006. The minimum wage for skilled workers is set by provincial labour regulatory bodies. Whether these laws are formally implemented in the work place cannot be determined due to paucity of data.

Our findings show that the large raw gaps in public and private sector earnings persist for both males and females even after conditioning on observed characteristics (such as education levels, experience, and province and region fixed-effects). More interestingly, controlling for worker unobservables using a household fixed-effects methodology does not eliminate public–private differentials for men, pointing to the existence of large sectoral differences and also indicating that ‘rent-seeking’ rather than our inability to control for unobservables may be generating these effects. This study also finds the public–private wage gap to be much larger for women than for men signifying labour market differentiation to be far greater for women. Finally, the decomposition exercise reveals that, in the case of men, a higher proportion of the public–private wage-gap can be explained by observed characteristics than in the case of females.

This paper is organized as follows. Section 2 presents the econometric methodology underlying this study. Section 3 discusses the characteristics of the data while Section 4 presents the results. Section 5 summarises and concludes.

2. Econometric approach

Consistent estimation of earnings functions is important not only in its own right but also, in this study, for arriving at estimates that can be used for a robust decomposition of the private and public sector wage-differentials. There are two main approaches used in the literature to identify public–private wage differentials—the ‘single equation method’ (or the dummy variable approach) and the ‘separate equations method’. In the first, a single earnings function is estimated with the usual covariates incorporating one or more dummy variables to represent sectoral choice. This usually takes the following form:

$$\mathrm{Ln}Y_i = \beta + \beta_{1i}\mathbf{X}_i + \beta_2\mathbf{PUBLIC}_i + \varepsilon_i \tag{1}$$

where Ln Y_i is the log of wages of individual i, $\mathbf{X_i}$ is a vector of observed characteristics of individual i (including experience, gender, schooling, etc.), $\mathbf{PUBLIC}_i$ is a dummy variable equalling ‘1’ if individual is employed in the public sector and ‘0’ otherwise and ε_i is the individual-specific error. The coefficient β_2 measures the premium, if any, to belonging to the public sector.

The ‘dummy variable’ approach suffers two main drawbacks. First, it constrains the vector of all other coefficients to be identical across the sectors. For example, while there may be significant differences in how education is rewarded in the public and private sectors, the β coefficients (used to measure the return to education) are not allowed to vary by sector. The second problem has to do with the potential endogeneity of the dummy variable representing sectoral choice. This problem arises because entry into the public or private sector may be determined by variables often unobserved by researchers. For example, more able or motivated individuals may strive for employment, say, in the public sector. Moreover, if employers also reward these traits in the form of higher earnings, these unobserved characteristics will reside in the error term ε_i in (1). The potential (positive) correlation between $\mathbf{PUBLIC}_i$ and ε_i violates the basic conditions of the classical linear model and would generate an upward bias in all parameter estimates.

The alternative, ‘separate equations’ approach overcomes one of the drawbacks of the ‘single equations’ method—by allowing the vector of coefficients to vary by sector. Thus, the two earnings functions to be estimated are

$$\mathrm{Ln}\, Y_{1i} = \beta_{1i}\mathbf{X}_i + \varepsilon_{1i} \tag{2}$$

$$\mathrm{Ln}\, Y_{2i} = \beta_{2i}\mathbf{X}_i + \varepsilon_{2i} \tag{3}$$

where $\mathbf{X}$ is the vector of explanatory variables, β represents the corresponding vector of coefficients, ε are the i.i.d error terms and the subscripts 1 and 2 denote the public and private sectors, respectively.

However, OLS estimation of Eqs. (2) and (3) may not yield consistent results. This is mainly because endogeneity of $\mathbf{PUBLIC}_i$ in the ‘single equation’ method translates into ‘sample selection’ in a multi-equation framework. The sectoral sub-sample of workers is a potentially non-random draw from the population. The standard approach in the literature addresses this concern by including an additional regressor (lambda) in earnings functions which corrects for the bias generated through sector-selection (Heckman, 1979).

However, two further problems remain unaddressed. First, while recognising the bias arising from selection into a given sector (and correcting for it using a Heckman–Lee procedure in the ‘double equations’ method or instrumenting for the sectoral dummy in the ‘single equations’ approach), a second source of endogeneity/selection, though recognised in various studies, is often not corrected for in econometric estimates of public–private wage differentials. This bias occurs because most earnings functions are estimated on samples of wage-earners only and exclude non-labour force participants, the unemployed, the self-employed, etc. (termed ‘others’ in this paper). Correcting for this second source of sample selection would require a further correction term in the earnings functions for consistent estimation.

In order to allow for both types of selection (i.e. the wage participation versus non-participation decision and the choice of public or private sector conditional on participation), one could use a multinomial logit model (MNL) in the first step to represent three choices: (1) waged work in the public sector; (2) waged work in the private sector; and (3) 'other', i.e. non-labour force participation, self-employment, etc. The correction-terms generated from the first step can then be incorporated in the earnings functions as additional regressors as suggested by Lee (1983). We adopt this approach in this paper as it attempts to address both selection issues without imposing strong prior the selection process [7].

Finally, the 'schooling' variable in earnings functions is also potentially endogenous. While 'separate equation' models with employment choice estimated through a multinomial model will control for the two potential selection issues, the endogeneity of schooling remains unresolved. The household fixed-effects technique provides a useful though not entirely convincing solution for dealing simultaneously with the issues surrounding endogeneity *and* sample selectivity. This approach rests on the grounds that arguably a good part of the unobserved heterogeneity (generating endogeneity in schooling and selection biases) is common to family members and any differences in unobserved ability and their impact in determining education should be lower *within* rather than *between* families. By introducing sub-samples of households with at least two wage-earning individuals of a given gender in a household (and also in a given sector of employment in this case), this 'first-differencing' approach effectively controls for all household-level variables common across these individuals within a household. Furthermore, as most studies (including the current one) control for sample-selection using observed household-level variables such as demographic composition and asset-ownership as exclusion restrictions determining participation (and sectoral choice), controlling for household fixed-effects also simultaneously controls for sample selection issues (Behrman & Deolalikar, 1995; Pitt & Rosenzweig, 1990). The parameter estimates generated through this technique are also used to decompose the wage gap between public and private sector employees[8].

Summarising, our empirical strategy will be as follows. We start by estimating earnings functions using Ordinary Least Squares (OLS) on a pooled sample of wage-employed males and females working in public and private sectors. Then we will split the sample into workers (male and female) employed either in the public or in the private sectors. Recognising that these estimates constrain the vector of coefficients for both genders to be identical, albeit across sectors, our final OLS estimates will separately estimate earnings functions on males in public and private sector jobs and on females in public and private sector employment. As sub-sample analysis imposes sample-selection concerns, correction-terms generated from a first-step multinomial logit will be included in earnings functions to determine whether selectivity poses any potential biases in our sample. Finally, individual earnings in public and private sectors (for both males and females) will be decomposed (using Oaxaca's method) into the portion 'explained' by characteristics and the 'unexplained' portion to determine whether public–private wage differentials prevail and if they can be explained by differences in the endowments of workers.

Broadly speaking, two different empirical approaches can be used for studying public–private sector wage differentials. The first of these includes the PUBLIC sector dummy variable as a predictor of 'discrimination' in a pooled regression of earnings functions. However, as already mentioned above, this approach yields biased results because it assumes that the wage structure is the same for both sectors. A second approach employs 'decomposition' to separate the observed wage gap between sectors into the components that are 'explained' by differences in characteristics across sectors and the 'unexplained' portions. This method was first developed by Blinder (1973) and Oaxaca (1973), and later extended to overcome the index number problem (Cotton, 1988; Neumark, 1988).

We decompose the public–private wage gap using the technique proposed by Oaxaca (1973). OLS, selectivity-corrected and household fixed-effects estimates of earnings functions for men and women are used to predict earnings. The wage-gap is decomposed into two components: (1) the portion 'explained' by differences in characteristics of workers in either of the two sectors and (2) the residual, 'unexplained' portion reflecting differences in wage structures or rewards across public and private sectors. The unexplained component could represent differential rewards to possessing characteristics or 'rents' in the labour market. However, if there are important differences in the unobserved or unmeasured characteristics of public and private sector employees, then the residual component cannot be purely attributed to 'rents' and one must wonder whether differences in unobserved or unmeasured individual ability or even motivation generate this gap in earnings across individuals in the two sectors. However, Oaxaca and Ransom (1999) note that results of the decomposition exercise are sensitive to the underlying method used to estimate wage gaps.

3. Data and descriptive statistics

The data used in this study are drawn from the latest, nationally representative household survey from Pakistan: the PSLM, 2005. This dataset is based on a sample of more than 70,000 households from rural and urban regions across the four provinces (Punjab, Sindh, NWFP and Balochistan) and from *each* district in the country (from more than a 100 districts).

[7] If one strongly believed that labour market choices are made in a sequential way in Pakistan, i.e. that individuals first choose whether to be wage employed and, conditional on being that decide to work either in the public or private sector, the bivariate probit may have been the model of choice.

[8] Pure family-effect models are most plausible for identical twins but not so convincing for father-son, mother-daughter or sibling pairs. This is because while identical twins have the same genes and may have faced similar backgrounds, this may not be the case of parent–child pairs or even for siblings within the same household.

Table 1
Description of variables used in OLS, FE and the MNL functions, ages 15–65

Variable	Description
WAGEWORK	Participation in salaried wage work during the past month
MALE	Dummy variable, equals 1 if individual is male, 0 otherwise
AGEYRS	Age in completed years
AGE2	Square of age
NOEDU	Dummy equals 1 if individual has completed 0 years of education, 0 otherwise
LESSPRIM	Dummy equals 1 if individual has completed 1–4 years of education, 0 otherwise
PRIMARY	Dummy equals 1 if individual has completed 5–7 years of education, 0 otherwise
MIDDLE	Dummy equals 1 if individual has completed 8 or 9 years of education, 0 otherwise
MATRIC	Dummy equals 1 if individual has completed 10 years of education, 0 otherwise
INTER	Dummy equals 1 if individual has completed FA/FSc, 0 otherwise
BACHELORS	Dummy equals 1 if individual has completed BA/BSc, 0 otherwise
MA	Dummy equals 1 if individual has completed MA, MSc, 0 otherwise
ODEGREES	Dummy equals 1 if individual has obtained a degree in engineering, MBBS, computers, agriculture, MPhil/PhD. Or other, 0 otherwise
URBAN	Dummy equals 1 if resides in urban area, 0 otherwise
PUNJAB	Province is Punjab, yes = 1, no = 0
SINDH	Province is Sindh, yes = 1, no = 0
NWFP	Province is NWFP, yes = 1, no = 0
BALOCHISTAN	Province is Balochistan, yes = 1, no = 0
PUBLIC	Sector of work is public = 1, private = 0
CHILD5	Number of children aged 5 or less in the household
ADULT70	Number of adults aged 70 or more in the household
MARRIED	Married, yes = 1, no = 0
OALAND	Own agricultural land, yes = 1, no = 0
ONALAND	Own non-agricultural land, yes = 1, no = 0
LAMBDA	Selectivity term

However, we rely on the Household Income and Expenditure Survey (HIES) portion of the survey on a sub-sample of some 14,000 households on which detailed information needed for the calculation of earnings functions was collected.

The HIES-section of the PSLM asked detailed employment questions from all individuals aged 10 and above. However, in line with past work, our analysis is restricted to individuals aged 15–65 (55,723 observations) and those in wage employment (relegating individuals reporting self-employment in agriculture or non-agricultural activities, unpaid family work, unemployment or non-labour force participation to the ‘Other’ category). This leaves us with a sample of 10,884 individuals (9640 males and 1244 females); i.e. roughly 20% individuals aged 15–65 are wage employees (a higher proportion, 34% males, and only about 4.5% females in this age group are wage employees)[9]. Finally, for the purpose of this study, the unique feature of this survey (compared to past household datasets such as the Pakistan Integrated Household Survey) is that is asks all individuals aged 10 and above to report their ‘sector of employment’. The question allows for five categories: (1) public; (2) private business; (3) private; (4) NGO; and (5) other. As we have restricted the sample to wage workers only, the answer to this question will determine whether a *wage employee* works in either of the above sectors. We collapse the private-business, private, NGO and ‘other’ category into a single category called ‘private’. Among the 10,884 wage workers, roughly 27% report employment in the public sector and the remaining are deemed to be working in private-sector jobs[10]. There is no difference in sectoral choice by gender (2623 males and 352 females in waged work are public sector employees in this sample).

Table 1 describes the variables used in estimation and Table 2 shows summary statistics for the full sample and for the three employment sectors: private, public and ‘Other’. The dependent variable in the earnings functions is LNMEARN (log of monthly earnings in rupees). The standard Mincerian function stipulates that individual earnings are a function of experience and education. Experience is often computed as (age – years of completed schooling – 5) with the view that individuals start school aged 5 and enter the labour market upon completing schooling. This can be misleading for Pakistan not only because children may not enter school aged 5 (and this may differ by gender), but also because a large proportion of the labour force is illiterate, having never attended school at all. Another constraint in the PSLM is that it asks individuals about ‘completed levels of schooling’ rather than ‘completed years of schooling’. For these reasons, AGEYRS (and the quadratic AGE2) are used to proxy for experience and experience squared. Education is denoted in the form of dummy variables indicating levels of completed schooling with no education (NOEDU) as the base category. Unless otherwise stated, province and regional fixed effects are included in all earnings functions to control for any provincial or regional differences in earnings.

[9] These proportions of wage employees are relatively smaller compared with the PIHS (2002) figures reported in Aslam (2007a). While roughly similar proportions of both men and women (aged 15–65) are wage employees in the PIHS (2002) and PSLM (2005) data sets—23 and 20%, and 42% men and almost 7% women reported wage employment according to the PIHS (2002), these figures fall to only 34 and 4.5% according to the PSLM (2005).

[10] The proportion of wage-employed individuals in the public sector from the PSLM dataset (26%) is much smaller than the proportions reported by Nasir (2000) using the Labour Force Survey, 1997 (56%) and by Hyder and Reilly (2005) using the 2002 Labour Force Survey (45%).

Table 2
Summary statistics of variables used in the MNL selection equations and in earnings functions, males and females 15–65 in waged work

Variable	Mean characteristics of males				Mean characteristics of females			
	Waged work in		'Other'	All	Waged work in		'Other'	All
	Public	Private			Public	Private		
WAGEWORK	1.00 (0.00)	1.00 (0.00)	0.00 (0.00)	0.34 (0.47)	1.00 (0.00)	1.00 (0.00)	0.00 (0.00)	0.05 (0.21)
PUBLIC	–	–	–	0.094 (0.29)	–	–	–	0.013 (0.11)
LNMEARN	8.73 (0.63)	8.01 (0.69)	–	–	8.53 (0.72)	7.16 (0.87)	–	–
AGEYRS	38.68 (9.81)	31.53 (12.25)	31.79 (14.96)	32.37 (14.06)	33.84 (9.50)	31.05 (12.31)	32.06 (13.51)	32.05 (13.43)
AGE2	1592.49 (775.22)	1143.99 (899.95)	1234.59 (1123.58)	1245.34 (1045.75)	1235.03 (722.60)	1115.72 (881.22)	1210.16 (997.98)	1207.42 (991.49)
LESSPRIM	0.02 (0.15)	0.07 (0.26)	0.06 (0.23)	0.06 (0.23)	0.00 (0.05)	0.03 (0.16)	0.03 (0.17)	0.03 (0.17)
PRIMARY	0.07 (0.26)	0.13 (0.33)	0.11 (0.31)	0.11 (0.31)	0.00 (0.00)	0.04 (0.20)	0.08 (0.27)	0.08 (0.27)
MIDDLE	0.07 (0.26)	0.15 (0.36)	0.16 (0.37)	0.15 (0.36)	0.02 (0.13)	0.03 (0.18)	0.08 (0.27)	0.07 (0.26)
MATRIC	0.25 (0.43)	0.15 (0.36)	0.22 (0.42)	0.21 (0.41)	0.26 (0.26)	0.08 (0.27)	0.10 (0.30)	0.10 (0.31)
INTER	0.12 (0.32)	0.04 (0.19)	0.06 (0.23)	0.06 (0.23)	0.16 (0.36)	0.04 (0.21)	0.03 (0.18)	0.04 (0.19)
BACHELORS	0.20 (0.40)	0.04 (0.20)	0.04 (0.20)	0.06 (0.23)	0.25 (0.43)	0.06 (0.24)	0.03 (0.16)	0.03 (0.17)
MA	0.10 (0.30)	0.01 (0.11)	0.01 (0.09)	0.02 (0.13)	0.17 (0.38)	0.03 (0.18)	0.00 (0.06)	0.01 (0.08)
ODEGREES	0.05 (0.22)	0.01 (0.10)	0.01 (0.10)	0.01 (0.12)	0.06 (0.23)	0.02 (0.14)	0.00 (0.05)	0.00 (0.06)
URBAN	0.55 (0.50)	0.47 (0.50)	0.38 (0.49)	0.42 (0.49)	0.62 (0.49)	0.47 (0.50)	0.41 (0.49)	0.41 (0.49)
SINDH	0.26 (0.44)	0.25 (0.43)	0.24 (0.43)	0.25 (0.43)	0.18 (0.39)	0.18 (0.38)	0.23 (0.42)	0.23 (0.42)
NWFP	0.21 (0.41)	0.20 (0.40)	0.21 (0.40)	0.21 (0.40)	0.31 (0.46)	0.08 (0.27)	0.23 (0.42)	0.23 (0.42)
BALOCHISTAN	0.24 (0.43)	0.12 (0.33)	0.16 (0.36)	0.15 (0.36)	0.11 (0.32)	0.03 (0.17)	0.14 (0.34)	0.13 (0.34)
CHILD5	1.25 (1.34)	1.14 (1.37)	1.19 (1.49)	1.18 (1.45)	1.15 (1.48)	0.95 (1.25)	1.28 (1.50)	1.27 (1.49)
ADULT70	0.13 (0.40)	0.12 (0.37)	0.14 (0.40)	0.13 (0.39)	0.21 (0.47)	0.14 (0.39)	0.15 (0.41)	0.15 (0.41)
MARRIED	0.87 (0.33)	0.60 (0.49)	0.52 (0.50)	0.57 (0.49)	0.70 (0.46)	0.52 (0.50)	0.67 (0.47)	0.67 (0.47)
ONALAND	0.06 (0.24)	0.03 (0.18)	0.05 (0.23)	0.05 (0.22)	0.11 (0.31)	0.05 (0.22)	0.05 (0.22)	0.05 (0.22)
N	2623	7017	18,539	28,179	352	892	26,300	27,544

Note: 'Public' includes individuals reporting sector of employment to be 'government' and 'Private' constitutes those reporting employment in private or NGO sector. 'Other' constitutes individuals who are non-labour force participants, unemployed, livestock-sellers and self-employed (agriculture/non-agriculture). S.E.s are reported in parentheses.

Table 3
Occupational status and monthly earnings for wage earners by sector, gender (15–65)

Occupation	Male				Female			
	Public		Private		Public		Private	
	%	Earning (Rs. per month)	%	Earning (Rs. per month)	%	Earning (Rs. per month)	%	Earning (Rs. per month)
Senior officials	8.1	16,021	1.6	14,616	5.1	11,262	1.0	7633
Professionals	15.9	9,014	3.0	9,241	39.5	7,277	10.0	5595
Ass. professionals	12.0	10,308	2.8	5,805	12.2	8,718	3.6	2425
Clerks	15.4	6,653	2.1	4,960	5.7	4,634	0.8	4417
Service, shop, sale	26.3	6,719	34.2	3,892	26.7	5,483	24.2	1847
Skilled Agri	0.7	5,650	6.7	2,600	0.3	–	19.2	1202
Craft and Trade	0.8	6,135	5.8	4,169	0.6	1,500	6.4	1112
Plant, machine Op.	3.4	5,765	9.1	4,130	0.0	–	1.2	2690
Elementary	17.5	6,849	34.7	3,579	9.7	5,885	33.6	1634
All	100	8,215	100	4,150	100	6,632	100	2101

Finally, selectivity-correction in stage 1 (in the multinomial logit) requires finding exclusion restrictions—variables that determine sectoral choice but not earnings conditional on being in that sector. As always, this is a challenge. Conforming to past work, we use household demographic variables (CHILD5 and ADULT70) and land ownership (ONALAND). These are chosen on the belief that the presence of very small children or elderly individuals in the household and ownership of an asset like land may determine an individual's choice between the three options: public sector-waged employment, private sector-waged employment or 'Other' without having a direct effect on individual earnings. For example, public sector-waged employment often offers flexibility in working conditions which may suit a woman with younger children. Alternatively, land ownership may provide the 'safety net' that may encourage individuals to forgo the security of wage employment and become either self-employed, give up a current job and seek another (unemployed), or even completely exit the labour force.

Table 2 shows striking differences between public and private sector employees. Both males and females earn significantly more in public-sector jobs. Interestingly, the public-sector wage premium is higher for women who earn about 68% more in government jobs compared to females in private wage employment. Men in government jobs, on the other hand, earn 50% more than their counterparts in private jobs. Both men and women in public sector jobs are older than in private jobs and if age proxies for experience this suggests that government sector employees are marginally more experienced than private sector ones. Finally, for both genders the raw data shows that public-sector wage employees are more educated—a high proportion of men in government (versus private) jobs have at least 10 years of education (Matric). This is true for females as well suggesting that more educated females opt to seek employment in the public sector.

Finally, it is worth noting that we do not condition on occupation (unlike Nasir, 2000) because our primary objective in this study is to consistently estimate earnings functions by employment sector and gender to determine the degree of wage

Table 4
OLS earnings functions (males/females) aged 15–65

Variable	OLS		
	Pooled coefficient (S.E.) (1)	Public coefficient (S.E.) (2)	Private coefficient (S.E.) (3)
MALE	0.648 (0.043)***	0.190 (0.039)***	0.829 (0.047)***
AGEYRS	0.074 (0.004)***	0.064 (0.007)***	0.079 (0.004)***
AGE2	−0.001 (0.000)***	−0.001 (0.000)***	−0.001 (0.000)***
LESSPRIM	0.065 (0.035)*	−0.005 (0.058)	0.050 (0.037)
PRIMARY	0.111 (0.023)***	0.109 (0.048)**	0.085 (0.025)***
MIDDLE	0.193 (0.025)***	0.140 (0.039)***	0.176 (0.026)***
MATRIC	0.302 (0.027)***	0.300 (0.040)***	0.252 (0.025)***
INTER	0.432 (0.054)***	0.448 (0.042)***	0.356 (0.070)***
BACHELORS	0.621 (0.060)***	0.582 (0.045)***	0.612 (0.095)***
MA	1.001 (0.079)***	0.843 (0.056)***	1.299 (0.116)***
ODEGREES	1.081 (0.105)***	1.006 (0.076)***	1.138 (0.158)***
URBAN	0.188 (0.025)***	0.139 (0.023)***	0.195 (0.028)***
SINDH	0.053 (0.074)	−0.089 (0.044)**	0.089 (0.080)
NWFP	−0.058 (0.037)	−0.108 (0.034)***	−0.077 (0.041)*
BALOCHISTAN	0.163 (0.046)***	0.063 (0.028)**	0.198 (0.061)***
PUBLIC	0.320 (0.054)***	–	–
CONSTANT	5.657 (0.082)***	6.494 (0.153)***	5.442 (0.081)***
R^2	0.455	0.3444	0.354
N	10,728	2942	7786

Note: *, ** and *** denote significance at the 10, 5 and 1% levels, respectively. The dependent variable is natural log of monthly earnings (Rupees). Robust S.E.s are reported in parentheses. (–) Denotes not applicable. NO_EDUCATION and PUNJAB are the reference categories for education splines and province, respectively.

differentials. As occupational choice is likely to be determined by education, conditioning on occupation would change the interpretation of school effects. However, Table 3 shows some interesting descriptive statistics disaggregating occupational choice by gender and sector of employment. For instance, the highest proportion of men employed in the public sector work in services, shop and sales-related occupations (26%) followed by almost 18% in elementary occupations. Men in the private sector are also invariably concentrated in these two employment sectors (roughly 34% in each of the occupation categories). It is also apparent from the raw data that almost invariably (with the exception of professionals) wages are substantially lower in the private sector. For women, among those working in the public sector almost 40% are professionals while 27% are concentrated in services, shop and sales-related occupations. In the private sector, on the other hand, wage-working women are concentrated in low-skilled occupations such as services, agriculture and elementary occupations. This is hardly surprising given that raw data in Table 2 has shown that men and women in the public sector have attained higher education levels compared to their counterparts in the private sector.

4. Results

This section starts by presenting the earnings functions estimates followed by an Oaxaca decomposition of private–public wage differentials in the Pakistani labour market. The analysis begins with estimating 'single equation' OLS models on pooled samples of male and female wage employees by incorporating a dummy variable (PUBLIC) representing the employment sector in which the individual works. The results are presented in column (1) of Table 4 and are not surprising—male wage workers earn significantly more than females and the age-profile has the standard shape. The coefficients on various levels of schooling increase with higher levels of education, suggesting a convex education-earnings profile, a finding consistent with recent work in Pakistan (Aslam, 2007a,b; Kingdon & Soderbom, 2007). For the purpose of this study, the coefficient of interest is that on the PUBLIC dummy—it is large and significantly positive suggesting a wage premium to public-sector employees.

However, as mentioned in Section 2, the specification in column (1) has several drawbacks, one being that it constrains equality in the vector of coefficients across the two sectors. Columns (2) and (3) in Table 4 re-estimate earnings functions separately for the public and private sector wage employees to overcome this restriction. There are some interesting differences in earnings function estimates across the two sectors. For instance, the gender gap in earnings among public-sector workers is significantly smaller than in the private sector. This is unsurprising as government pay scales are compressed for men and women. The existence of a gender gap *within* the government sector can be attributed to different occupational choices between men and women. Table 3 has revealed that within the government sector, the highest proportion of women is in 'professional' occupations. However, even *within* this broad occupation category (such as 'professional'), women earn significantly less than men. This could be because among professional occupations, women opt for the more flexible jobs with lower working hours. Because data on hours worked is not available, this explanation cannot be further tested.

The age-profiles also differ by employment sector—while earnings peak earlier for government sector workers (32 years), they peak much later (almost 40 years) for private sector employees, a gap of almost 8 years. The pattern on the education-

Table 5
OLS and selectivity-corrected earnings functions, males (15–65)—public and private

Variable	OLS		Selectivity corrected	
	Public (1) coefficient (Robust S.E.)	Private (2) coefficient (Robust S.E.)	Public (3) coefficient (Robust S.E.)	Private (4) coefficient (Robust S.E.)
AGEYRS	0.063 (0.008)***	0.088 (0.004)***	0.037 (0.033)	0.089 (0.005)***
AGE2	−0.001 (0.000)***	−0.001 (0.000)***	0.000 (0.000)	−0.001 (0.000)***
LESSPRIM	0.038 (0.046)	0.042 (0.036)	0.001 (0.062)	0.042 (0.036)
PRIMARY	0.116 (0.047)**	0.083 (0.024)***	0.056 (0.090)	0.082 (0.024)***
MIDDLE	0.161 (0.038)***	0.166 (0.024)***	0.099 (0.093)	0.162 (0.026)***
MATRIC	0.310 (0.037)***	0.254 (0.024)***	0.193 (0.156)	0.247 (0.030)***
INTER	0.461 (0.045)***	0.319 (0.062)***	0.316 (0.182)*	0.310 (0.066)***
BACHELORS	0.590 (0.044)***	0.567 (0.089)***	0.422 (0.214)**	0.559 (0.092)***
MA	0.833 (0.056)***	1.152 (0.122)***	0.631 (0.264)***	1.144 (0.123)***
ODEGREES	0.999 (0.075)***	1.036 (0.153)***	0.831 (0.233)***	1.029 (0.155)***
URBAN	0.141 (0.023)***	0.205 (0.028)***	0.142 (0.023)***	0.208 (0.028)***
SINDH	−0.097 (0.045)**	0.068 (0.078)	−0.114 (0.046)**	0.067 (0.078)
NWFP	−0.142 (0.037)***	−0.094 (0.042)**	−0.168 (0.040)***	−0.095 (0.042)
BALOCHISTAN	0.053 (0.028)*	0.170 (0.058)***	−0.019 (0.094)	0.166 (0.060)**
Λ	–	–	−0.132 (0.162)	0.017 (0.043)
CONSTANT	6.699 (0.172)***	6.120 (0.081)***	7.549 (1.057)***	6.090 (0.110)***
R^2	0.345	0.274	0.345	0.274
N	2594	6921	2594	6921
Mean (Dep. Var.)				

Note: *, ** and *** denote significance at the 10, 5 and 1% levels, respectively. The dependent variable is natural log of monthly earnings (Rupees). Robust S.E.s are reported in parentheses. (–) Denotes not applicable. NO_EDUCATION and PUNJAB are the reference categories for education splines and province, respectively.

Table 6
OLS and selectivity-corrected earnings functions, females (15–65)—public and private

Variable	OLS		Selectivity corrected	
	Public (1) coefficient (Robust S.E.)	Private (2) coefficient (Robust S.E.)	Public (3) coefficient (Robust S.E.)	Private (4) coefficient (Robust S.E.)
AGEYRS	0.079 (0.019)***	0.019 (0.010)*	0.205 (0.096)**	0.018 (0.010)*
AGE2	−0.001 (0.000)***	0.000 (0.000)	−0.002 (0.001)*	0.000 (0.000)
LESSPRIM	−2.188 (0.124)***	0.026 (0.144)	−2.211 (0.126)***	0.075 (0.142)
PRIMARY	–	−0.020 (0.140)	–	0.066 (0.163)
MIDDLE	−0.463 (0.171)***	0.322 (0.128)**	−0.088 (0.311)	0.427 (0.148)***
MATRIC	0.158 (0.164)	0.056 (0.102)	1.277 (0.786)	0.132 (0.106)
INTER	0.300 (0.153)*	0.453 (0.180)*	1.606 (0.916)*	0.493 (0.167)***
BACHELORS	0.485 (0.155)***	0.743 (0.195)***	1.963 (1.044)*	0.724 (0.195)***
MA	0.864 (0.174)***	1.677 (0.221)***	2.752 (1.348)**	1.582 (0.245)***
ODEGREES	1.021 (0.197)***	1.552 (0.276)***	2.772 (1.195)**	1.400 (0.307)***
URBAN	0.047 (0.071)	0.093 (0.076)	−0.294 (0.241)	0.078 (0.074)
SINDH	−0.075 (0.084)	0.300 (0.131)**	−0.096 (0.086)	0.373 (0.147)**
NWFP	0.079 (0.087)	0.077 (0.108)	0.387 (0.225)*	0.223 (0.185)
BALOCHISTAN	0.113 (0.084)	0.644 (0.171)***	0.454 (0.243)*	0.827 (0.240)***
Λ	–	–	0.765 (0.541)	−0.216 (0.184)
CONSTANT	6.308 (0.329)***	6.503 (0.208)***	1.360 (3.586)	6.920 (0.322)***
R^2	0.353	0.303	0.357	0.304
N	348	865	348	865

Note: *, ** and *** denote significance at the 10, 5 and 1% levels, respectively. The dependent variable is natural log of monthly earnings (Rupees). Robust S.E.s are reported in parentheses. (–) **Denotes not applicable. NO_EDUCATION and PUNJAB are the reference categories for education splines and province, respectively.

dummy coefficients shows coefficients increasing with increasing levels of education. The coefficients for different education levels are not significantly different between the two sectors.

Wage determinants may differ across the gender domain and constraining the vector of coefficients in the public and private sector by introducing a gender-dummy may be too restrictive. Tables 5 and 6 re-estimate OLS earnings functions separately for males and females in public and private sector employment. The results are reported in columns (1) and (2) for males in Table 5 and similarly for females in Table 6. Columns (3) and (4) of the aforementioned tables show findings from the selection-corrected estimates which include a correction-term generated from first-step multinomial logit estimates (reported in Appendix Tables A1 for males and A2 for females).

Before discussing the results of main interest in Tables 5 and 6, focus briefly on the sectoral choice equations reported in Tables A1 and A2. The dependent variable is SECTOR and equals 1 if individual is either a non-labour force participant, unemployed, unpaid family worker or self-employed (in agriculture or otherwise) and this category is defined as ‘Other’, 2 if a wage employee in the private sector (‘private’) and 3 if employed in the public sector and paid regular wages (‘public’). The excluded category is ‘Other’. The multinomial estimates are important for the generation of the selection-terms to be included in the earnings functions and in determining whether sample selection is important. Because the objective of the first-step estimation is to control for sample selection (if any), we discuss only the signs and significance of the exclusion restrictions reported in Tables A1 and A2. As mentioned before, we use three exclusion restrictions: CHILD 5, ADULT70 and ONALAND. The exclusion restrictions were pared down to a more parsimonious set after experimentation[11]. All three are individually significant for males in both the private and public employment sector. While the demographic variables have the expected negative sign, the sign on land-ownership is positive in both sectors suggesting that owning non-agricultural land increases likelihood of being employed on private or public sectors rather than being in the ‘Other’ category. For the sample of women (Table A2), we notice two things: first, none of the exclusion restrictions is significant for the public sector employees and secondly, having a child aged less than 5 significantly reduces the probability of employment in the private sector compared to ‘Other’. The exclusion restrictions are jointly significant at the 0.1% level for males in private and public sector jobs and for females in the private sector (p-value of the F-tests are 0.000) though this is not true for women in the public sector.

Tables 5 and 6 reports the earnings function estimates for males (females) in public and private sector employment without (columns 1 and 2) and with (columns 3 and 4) sample-selection correction. Focus first on the selection–correction terms in columns (3) and (4) of the two tables. Clearly, the selection–correction terms are small and insignificant for all. It

[11] To begin with, MARRIED and ONALAND were also included in the set of exclusion restrictions on the premise that being married and owning non-agricultural land may influence sectoral choice but should not directly impact earnings. However, MARRIED and ONALAND were significant in estimated earnings functions invalidating them as credible exclusion restrictions. They are, however, not included in the final versions of earnings functions as we are interested in estimating as close to a standard-Mincerian earnings function as possible (rather than an extended one). The validity of the three remaining exclusion restrictions (CHILD5, ADULT70 and OALAND) was also tested. In no instance (male and female earnings functions in private and public sectors) was either of these exclusion restrictions independently significant. F-tests also revealed the exclusion restrictions to be jointly insignificant in all four specifications (p-values of the F-tests for the joint significance of CHILD5, ADULT70 and OALAND were: 0.63, 0.15, 0.26 and 0.42 in the earnings functions for males in private and public-sector jobs and females in private and public-sector jobs, respectively).

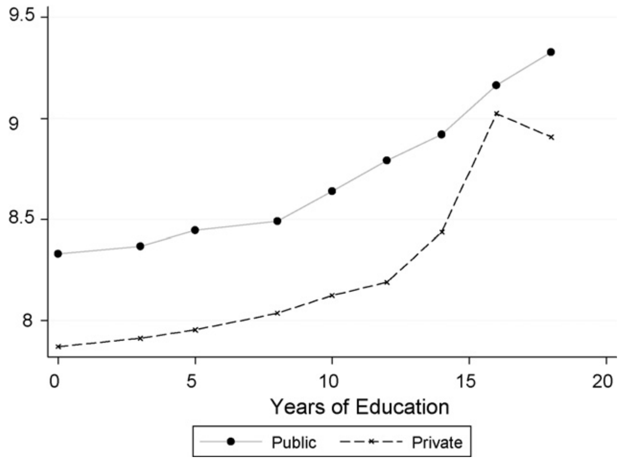

Fig. 1. Predicted earnings, males (public and private sector).

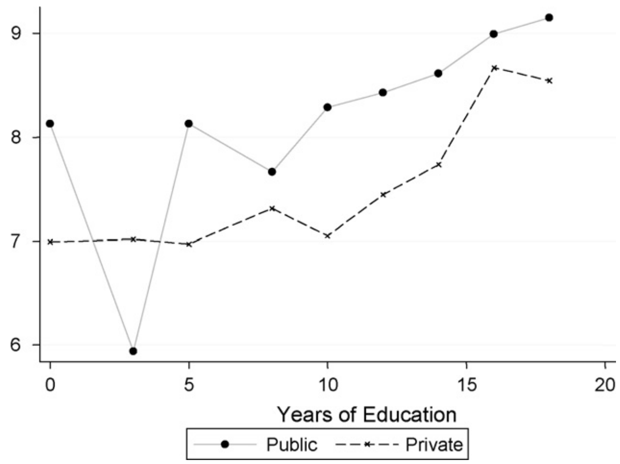

Fig. 2. Predicted earnings, females (public and private sector).

appears that the estimated effects corresponding to the selection terms are small. This finding, especially for women may be attributable, given the relatively small reported value for the constant term, to the fact that there is limited variation in the selection term across the female sub-sample. This, in turn, could be because the variables used to capture the observable characteristics of women who are selecting into the public sector are fairly similar. Thus, the selection term is competing for the role of the constant and the estimated effect of the constant gets washed-out. In this case, the relevant selection effects have potentially not been correctly identified here. However, we draw comfort from two recent studies from Pakistan and Bangladesh that report similar findings. In the first, Aslam, 2007a,b, using data from 2001 from Pakistan, estimates selection-corrected earnings functions for women wage earners and finds a similar result though the estimates for men show strong negative selection into wage employment (though those estimates do not differentiate by sector of employment). In Bangladesh, Asadullah (2005) also notes that the selection-terms for males and female are insignificant despite strongly significant exclusion restrictions in the selection equations. We use OLS estimates as our preferred estimates which are now discussed[12].

[12] However, the assignment of all self-employed individuals into 'other' may not be sensible as the process determining self-employment activity possibly differs substantially from that determining unemployment or being out of the labour force. We tried an alternative way of estimating the MNL model by allowing for four categories: out of the labour force/unemployed, self employed (agriculture and non-agriculture), wage employed in private sector and wage employed in the public sector. Selection-terms computed from this alternative model were incorporated in earnings functions but they remained insignificant and made no significant difference to the results. Moreover, Bourguignon et al. (2007) argue that the Dubin and McFadden (1984) procedure for selection bias correction is to be preferred to the commonly used Lee (1983) approach (used here). Moreover, they suggest that even if the Independence of Irrelevant Alternatives (IIA) is rejected, the multinomial logit model still provides fairly good correction for the outcome equation. The IIA was tested using the three category and four category MNL estimated in this paper (using the small-Hsiao test). The findings show that while the IIA is rejected in several cases in the four category MNL, in no instance is it rejected in the three category MNL model estimated in this study.

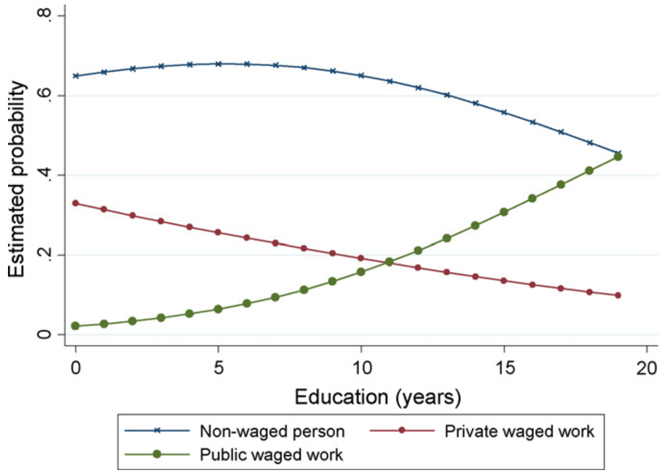

Fig. 3. Males: estimated probability of sectoral choice and education.

For males and females (Tables 5 and 6), note that the explanatory power (R^2) is higher for public-sector workers. This is because of stricter salary scales in this sector. For males (Table 5), the OLS estimates reveal standard age-earnings profiles in both sectors though earnings peak earlier (age 31.5) for public-sector employees and much later (age 44) for private-sector workers. The coefficients on the education splines are indicative of convex education-earning profiles in both sectors and also show that they do not significantly differ across the two sectors (except for INTER where the coefficient is of a greater magnitude in the public sector and MA where the coefficient is significantly greater for males in private-sector employment). Table 6 reveals a different picture for women. For instance, while women's age-earning profile has the standard shape in the public sector this is not the case in the private sector. Also note that women's earnings in the public sector peak aged almost 40, unlike public-sector males whose earnings peak almost 9 years earlier. The coefficients on the education dummies also show different profiles—coefficients are positive and increasing successively in both sectors only after INTER. Also note that while the private sector positively rewards women with at least MIDDLE-level schooling this is not the case in the public sector. This is possibly because a large majority of women opt for teaching jobs and while the private sector often hires women who have acquired at least middle-level schooling, the public sector imposes more stringent requirements and women can only enter teaching professions after completing at least MATRIC (Andrabi et al., 2006; Aslam, 2007b). The education–earnings profiles for men and women in the public and private sectors are plotted in Figs. 1 and 2. It is clear that public sector wages are invariably higher than private sector wages at all levels of education, even though the public–private wage gap narrows considerably at higher levels of education. For women the second data point (showing extremely low wages at 3 years of education) is driven by very few observations and is not statistically significant. Figs. 3 and 4 depict the estimated probability of sectoral choice by education. There are striking differences by gender: while for men the probability of private sector waged work steadily declines with education and the probability of public sector waged work steadily increases with education throughout the full range of education levels, for women there is a positive relationship between

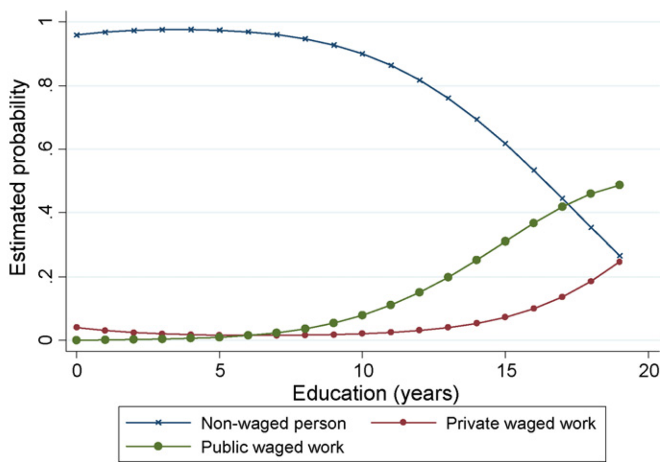

Fig. 4. Females: estimated probability of sectoral choice and education. *Note*: These predictions are based on the multinomial logits reported in Appendix A1 and A2 with education specified as continuous and non-linear rather than in splines.

Table 7A
Public–private wage decomposition—males (15–65), OLS

Based on OLS estimates	Males					
	Standardising by public means			Standardising by private means		
	Coefficients (*D*)	Characteristics (*E*)	Combined (*T*)	Coefficients (*D*)	Characteristics (*E*)	Combined **(*T*)
INTERCEPT	0.578	0.000	0.578	0.578	0.000	0.578
AGEYRS	−0.973	0.628	−0.345	−0.794	0.450	−0.345
AGE2	0.754	−0.448	0.306	0.543	−0.238	0.306
LESSPRIM	0.000	−0.002	−0.002	0.000	−0.002	−0.002
PRIMARY	0.002	−0.004	−0.002	0.004	−0.006	−0.002
MIDDLE	0.000	−0.012	−0.013	−0.001	−0.012	−0.013
MATRIC	0.014	0.026	0.040	0.008	0.032	0.040
INTER	0.017	0.026	0.042	0.005	0.037	0.042
BACHELORS	0.005	0.088	0.093	0.001	0.092	0.093
MA	−0.032	0.102	0.070	−0.004	0.074	0.070
ODEGREES	−0.002	0.038	0.036	0.000	0.037	0.036
URBAN	−0.036	0.017	−0.019	−0.030	0.011	−0.019
SINDH	−0.042	0.000	−0.042	−0.041	−0.001	−0.042
NWFP	−0.010	−0.001	−0.012	−0.010	−0.002	−0.012
BALOCHISTAN	−0.028	0.021	−0.008	−0.014	0.006	−0.008
Total	0.246	0.478	0.724	0.246	0.478	0.724
Unexplained gap (%)		33.9			33.9	

Note: Estimates based on parameters from columns (1) and (2) in Table 6.

education and chances of public waged work only after 8 years of schooling and between education and chances of private waged work after 15 years of schooling.

4.1. *Public–private wage differentials*

In this section, we examine public–private wage differentials separately for male and female wage employees. The total difference in earnings between public and private sector employees (male and female) is decomposed into the portion 'explained' by differences in characteristics between workers in the two sectors and the portion left 'unexplained' by differences in characteristics. This is done using Oaxaca's methodology (Oaxaca, 1973) detailed in Section 2. The preferred (OLS) estimates reported in columns (2) and (3) in Tables 5 and 6 are used. The results of the decomposition exercise are presented in Tables 7A for males and 7B for females.

For men, expressed in natural logs, the total difference in earnings between public and private sector employees is 0.724 both when standardising by public or by private sector means. Of this total, reported in the last row in Table 7A, 0.478 is explained by differences in the characteristics of individuals and 0.246 (i.e. about 34%) of this raw difference in earnings

Table 7B
Public–private wage decomposition—females (15–65), OLS

Based on OLS estimates	Females					
	Standardising by public means			Standardising by private means		
	Coefficients (*D*)	Characteristics (*E*)	Combined (*T*)	Coefficients (*D*)	Characteristics (*E*)	Combined (*T*)
INTERCEPT	−0.196	0.000	−0.196	−0.196	0.000	−0.196
AGEYRS	2.038	0.054	2.092	1.869	0.223	2.092
AGE2	−0.663	−0.025	−0.688	−0.599	−0.089	−0.688
LESSPRIM	−0.006	−0.001	−0.007	−0.056	0.049	−0.007
PRIMARY	0.000	0.001	0.001	0.001	0.000	0.001
MIDDLE	−0.014	−0.006	−0.019	−0.027	0.008	−0.019
MATRIC	0.027	0.010	0.037	0.008	0.028	0.037
INTER	−0.024	0.049	0.026	−0.007	0.033	0.026
BACHELORS	−0.065	0.139	0.074	−0.016	0.090	0.074
MA	−0.140	0.233	0.093	−0.027	0.120	0.093
ODEGREES	−0.031	0.062	0.032	−0.009	0.041	0.032
URBAN	−0.028	0.013	−0.015	−0.022	0.007	−0.015
SINDH	−0.068	0.000	−0.068	−0.068	0.000	−0.068
NWFP	0.001	0.017	0.018	0.000	0.018	0.018
BALOCHISTAN	−0.059	0.054	−0.006	−0.015	0.009	−0.006
Total	0.772	0.600	1.372	0.836	0.537	1.372
Unexplained gap (%)		56.3			60.9	

Note: Estimates based on parameters from columns (1) and (2) in Table 7.

remains ‘unexplained’. In other words, the relatively higher wages of public-sector male employees are not fully explained by superior experience, education, urban residence or provincial variations.

Table 7A also shows that much of the pro-public wage gap for men is explained by the superior characteristics of workers in public sector jobs rather than higher returns by the government for possessing these characteristics. For instance, for all education levels above (and including MATRIC), the proportions attributable to characteristics (E) are higher than those to the rewards to possessing those characteristics (D). Also note that of the total wage gap (0.724), education explains about 39 (35)% of the total wage gap when standardising my public (private) sector means.

Table 7B reports the OLS-based decomposition results for wage-working women. Firstly, notice that the raw gap in monthly earnings expressed in natural logs between the public and private sectors is greater for women at 1.372 compared to that for men (0.724). The larger public-sector wage premium for women (compared to men) can be seen as lesser wage discrimination in the public sector, similar to the argument proposed by Smith (1976). Note also that a smaller proportion of this wage-gap is ‘explained’ by difference in characteristics—about 60% of the sectoral wage gap is not explained by differences in observed characteristics between women in the two sectors. One explanation for this could be that unobserved or unmeasured variables are more important for women for pay determination compared to men and our study has not adequately quantified them. It may even be that the observed variables used to proxy ‘experience’ are less well measured for women compared to men. Even though women, on record, may have the same years of experience as men but due to family commitments they may take time off more often or work fewer hours in which case learning-by-doing would be less than for men. Alternatively, the larger ‘unexplained’ proportion for women compared to men could be because of a greater extent of ‘discrimination’ against them in the private sector. This is also borne out by the results in Table 4. Comparing the coefficients on MALE in columns (2) and (3), clearly, the coefficient is significantly larger in the private-sector compared to the public sector and shows that while males earn 19% more than women in the public sector, they earn more than four times as much in the private sector indicating a greater degree of differential treatment against women in the private compared to the public sector.

However, as with men, the pro-public gap in earnings is explained by women in government jobs possessing superior characteristics rather than the government sector rewarding these characteristics more (the larger number of negative signs in ‘rewards’ under column D is outweighed by the larger coefficients and positive signs in the ‘characteristics’ under column E). If anything, public sector jobs’ returns to almost all characteristics are less than in the private sector (negative signs) except AGEYRS which is substantially highly rewarded in government employment. Finally, we note that education explains about 35 (27)% of the total sectoral gap in log earnings (1.372) when standardising by public (private) sector means.

It is worth noting, however, that as the earnings functions do not control for hours worked (due to lack of data), some of the variation in earnings is attributable to hours worked. Thus, if private-sector wage earners work, on average, longer hours per month than individuals in the public sector, the public–private wage gap may be much narrower than its hourly equivalent. Some of this variation in hours worked may be captured through occupation controls and the decomposition exercise were repeated using the following one-digit occupation dummies: managers (equals 1 if manager, 0 otherwise and so on), professionals, associate professionals, clerks, services, agriculture, crafts and machine operators. Those in elementary occupations constituted the omitted category. The decomposition exercise (using OLS) shows that between 28 and 33% of the public–private wage gap for men is unexplained while for women between 52 and 65% of the wage-gap is now unexplained (when standardising by public and private sector means, respectively). Thus, while the upper-bound estimate of the unexplained wage-gap remains largely unchanged for men, the incorporation of occupation dummies to control (albeit partially) for hours worked causes the upper-bound estimate to increase for women suggesting that part of the story indeed rests on the difference in hours worked[13].

Summarising, the results of the decomposition exercise reveal that some 34 (60%) of the public–private wage difference is not explained by public sector men (women) possessing superior characteristics compared to their counterparts in the private sector. Among the most common explanations proposed in extant literature for the existence of the ‘unexplained’ portion of the wage gap is the public-sector ‘economic rents’ argument. Alternatively, however, if there are important differences in the unobserved or unmeasured characteristics of workers in the two sectors, the ‘unexplained’ portion could be capturing these. Moreover, the decomposition results are based on OLS estimates but part of the problem with these is that the education dummies in earning function estimates are potentially endogenous. This endogeneity arises if unobserved variables such as ability increase individual earnings and also increase schooling (for instance if more able individuals also acquire more schooling). As ability is often unobserved, it resides in the error term of the earnings function. This will bias the OLS parameter estimates upward.

One way of addressing the aforementioned concerns is to use household fixed-effects (FE) on sub-samples of at least two wage workers of each gender in a given household. This provides a cleaner test since unobserved or unmeasured characteristics differences such as ability between males (or females) *within* the family are likely to be much lower than across families. Ideally, one should estimate earnings functions using household fixed-effects for both males and females. However, the sample of female wage earners in each sector is relatively small to begin with and would be considerably reduced by restricting it to at least two female wage earners in a given household to make meaningful inferences. However, there are slightly more than 2000 households with at least two male-wage earners aged 15–65 on whom this approach can

[13] The detailed results of earnings functions incorporating occupation dummies and the consequent decomposition exercise are available from the authors on request. We are grateful to a referee for suggesting this point.

Table 8
Public–private wage decomposition—males (15–65), household fixed-effects

Based on FE estimates	Males					
	Standardising by public means			Standardising by private means		
	Coefficients (*D*)	Characteristics (*E*)	Combined (*T*)	Coefficients (*D*)	Characteristics (*E*)	Combined (*T*)
INTERCEPT	0.459	0.000	0.459	0.459	0.000	0.459
AGEYRS	−0.596	0.335	−0.262	−0.498	0.236	−0.262
AGE2	0.522	−0.215	0.307	0.395	−0.088	0.307
LESSPRIM	−0.002	0.004	0.002	−0.007	0.010	0.002
PRIMARY	0.003	−0.003	−0.001	0.008	−0.008	−0.001
MIDDLE	0.000	−0.006	−0.007	−0.001	−0.006	−0.007
MATRIC	0.032	0.023	0.055	0.014	0.040	0.055
INTER	0.027	0.008	0.034	0.007	0.027	0.034
BACHELORS	0.055	0.026	0.081	0.009	0.072	0.081
MA	0.023	0.040	0.063	0.002	0.061	0.063
ODEGREES	−0.002	0.025	0.023	0.000	0.023	0.023
Total	0.519	0.236	0.755	0.388	0.367	0.755
Unexplained gap (%)		68.7			51.4	

Note: Estimates based on household fixed-effects results not reported but available from authors on request.

be applied. As we are interested in investigating public–private differentials, earnings functions are estimated using household fixed-effects on separate sub-samples of households with at least two male members in a given sector. All male members (related/unrelated directly) and recorded as household residents are initially included in the sample. As a robustness check, the decomposition exercise is done on the sample of only directly 'related' males (exclude non-relatives, fathers-in-law and sons-in-law of the household head). Two caveats must be noted: clearly, the household fixed-effects approach outlined above cannot net out ability differences. Also, imposing constraints on the samples introduces further selectivity problems. However, the fixed-effects approach is still a useful procedure as it is a powerful way of netting out at least environmental effects and thus helps in putting a tighter bound on the true effects.

The decomposition results based on the fixed-effects estimates are presented in Table 8. The fixed-effects findings show that an even larger proportion of the log wage-gap (0.755) is now unexplained compared to OLS estimates in 7A—about 69 and 51% when standardising by public and private means, respectively. These estimates rise even more when done on the sample of 'related' males—to 77 and 64%, respectively, when standardised by public and private sector means. If unobserved factors such as ability contribute towards the 'unexplained' portion rather than rents, we would expect the unexplained portion to *fall* after controlling for HH fixed effects. The rise in the 'unexplained' portion suggests that the wage differentials are arising because of rents or some other explanation rather than household-level unobserved factors controlled for in household fixed-effects estimation[14].

Our results partially corroborate and partially diverge from Nasir's (2000) findings. Contrary to our estimates, the author using Labour Force Survey data from 1997 for men only finds a relatively small public–private wage gap (0.05) in favour of the public sector of which a larger proportion is explained by superior characteristics of public-sector employees (0.12). However, the author notes that the rate of compensation of private sector workers is higher compared to government sector counterparts which is consistent with our finding that much of the gap in wages is due to superior public-sector characteristics rather than higher public sector rewards (for both men and women). It is important to note that Nasir distinguishes between the private formal and informal sectors and the aforementioned results are for the comparison between the public and private formal sectors[15]. However, the public–private informal sector findings are even more consistent with ours—of the wage-gap (0.526), almost 37% remains unexplained (not different from our findings for men in 8A). Our findings are also not very different from Hyder and Reilly's (2005) estimates based on a sample of men and women in waged employment in public, private and 'state-owned' enterprises (SOE). Using the 2002 Labour Force Survey from Pakistan, the authors find an average gap of 0.685 (between the public and private sectors) for a pooled sample of men and women, of which roughly 40% is explained by differences in characteristics (compared with 66% in our sample of men and 40% among women). The authors do not report, nor discuss, the differential 'returns/rewards' in the two sectors.

However, it is important to note that our findings are not directly comparable with the two previous studies because: (1) our findings are based on a household survey rather than a labour-force survey; (2) we use an all-encompassing definition of 'private'; (3) to our knowledge, ours is the only study in Pakistan that estimates public–private wage differentials separately for men *and* women; (4) this study attempts to highlight sample-selection using multinomial logits[16]; and (5) uses household fixed-effects to address the problem of unobservables that may generate biases in earnings function estimates.

[14] Summary statistics of wage earning males in the fixed-effects sample and earnings function estimates are available from the authors on request.

[15] The author's exposition of what constitutes the 'formal' and 'informal' private sectors is very unclear.

[16] While Hyder and Reilly (2005) address sample-selectivity in sectoral choice using Heckman's approach, they do not address the second cause of selection bias, i.e. the choice between labour force participation and non-participation. Nasir (2000) study uses simple OLS to estimate earnings functions and no mention is made of sample selection as a potential problem in his study.

5. Conclusion

The objective of this study was to examine wage differentials between public and private sector jobs for males and females in Pakistan. The raw public–private differential in gross wages was more than 1.5 times as high for men and more than three times as high for women. While conditioning on education and experience reduced the size of the public–private wage differential for both genders, the standardised differential remained very substantial.

Robust estimates of wage functions (and hence of wage differentials) were obtained using three methodologies: Ordinary Least Squares, Sample-selectivity-corrected and household fixed-effects. The parameters of the estimated wage functions were then used to decompose the public–private wage differential into the portion 'explained' by public–private differences in workers' mean characteristics and the 'unexplained' portion. The latest, nationally representative data (the Pakistan Living Standards Measurement Survey, PSLM, 2005) were used. To our knowledge, this is the first study to estimate *ceteris paribus* public–private wage differentials separately for males and females.

The vector of coefficients differed quite substantially across males and female wage-functions in the public and private sectors, highlighting the importance of estimating wage functions separately by gender. The Oaxaca-decomposition revealed that while differences in public–private worker characteristics explained about 66% of the public–private difference in log wage for men, the corresponding figure for women was only 40%. While for men, household fixed-effects could be used in an attempt to control for workers' unobserved characteristics, this was not possible on the small sample of female wage-earners. While in OLS estimates, 66% of the public–private wage gap for men was explained by the differing observed characteristics of public and private employees, in the household fixed-effects estimates for men, the explained proportion fell to about 40%, i.e. the residual or unexplained portion was about 60%.

What might account for the residual element? For men, we can argue that it could be 'rents' accruing to public sector workers (since we have controlled, arguably with some success, for men's unobserved characteristics via household fixed-effects estimation). However, for women we cannot rule out that it is differences in unobservables (observed and hence rewarded by employers, but not observed by researchers) that drive the residual. What we can say is that public sector workers invariably enjoy large premiums compared to private sector workers and this could simply be because they are 'lucky' to be working in the government sector (as argued by Lokshin & Glinskaya, 2005 for India) or have the right social connections and networks to become a part of the coveted public sector.

Appendix A

See Tables A1 and A2.

Table A1
Multinomial logit estimates, omitted category 'Other'—males (aged 15–65)

VARIABLE	Private (1) *z*-value (S.E.)	Public (2) *z*-value (S.E.)
AGEYRS	0.170 (0.007)***	0.533 (0.015)***
AGE2	−0.002 (0.000)***	−0.006 (0.000)***
LESSPRIM	0.054 (0.064)	0.730 (0.152)***
PRIMARY	−0.053 (0.052)	1.085 (0.104)***
MIDDLE	−0.355 (0.050)***	1.025 (0.106)***
MATRIC	−0.681 (0.049)***	1.954 (0.086)***
INTER	−0.905 (0.081)***	2.416 (0.106)***
BACHELORS	−0.424 (0.079)***	2.962 (0.099)***
MA	−0.012 (0.150)	3.814 (0.136)***
ODEGREES	−0.417 (0.148)***	2.956 (0.146)***
CHILD5	−0.067 (0.014)***	−0.031 (0.018)*
ADULT70	−0.089 (0.046)*	−0.191 (0.068)***
OALAND	1.148 (0.045)***	0.539 (0.060)***
URBAN	0.091 (0.040)**	−0.096 (0.063)
SINDH	−0.146 (0.046)***	0.208 (0.069)***
NWFP	0.057 (0.049)	0.518 (0.072)***
BALOCHISTAN	−0.399 (0.057)***	1.203 (0.075)***
CONSTANT	−5.060 (0.143)***	−14.743 (0.317)***
Pseudo-$R2$	0.158	
N	28,179	
Log-likelihood	−19992.9	

Note: *, ** and *** denote significance at the 10, 5 and 1% levels respectively. The dependent variable is categorised as (1) = participation in 'Other' where other constitutes non-labour force participation, unemployment, livestock ownership or self-employment (agricultural/non-agricultural), (2) = participation in waged work in private sector and (3) = participation in waged work in public sector. Robust standard errors are reported in parentheses and are corrected for clustering at the household-level. (–) Denotes not applicable. NO_EDUCATION and PUNJAB are the reference categories for education splines and province, respectively.

Table A2
Multinomial logit estimates, omitted category 'Other'—females (aged 15–65)

	Private (1) *z*-value (S.E.)	Public (2) *z*-value (S.E.)
AGEYRS	0.028 (0.016)*	0.418 (0.045)***
AGE2	−0.001 (0.000)**	−0.005 (0.001)***
LESSPRIM	−0.565 (0.227)*	0.324 (1.026)
PRIMARY	−1.150 (0.186)***	–
MIDDLE	−1.380 (0.203)***	1.666 (0.475)***
MATRIC	−0.815 (0.157)***	4.320 (0.272)***
INTER	−0.272 (0.203)	4.887 (0.295)***
BACHELORS	0.366 (0.177)**	5.461 (0.288)***
MA	1.571 (0.259)***	6.841 (0.332)***
ODEGREES	1.985 (0.332)***	6.553 (0.392)***
CHILD5	−0.146 (0.038)***	0.022 (0.046)
ADULT70	−0.105 (0.111)	−0.011 (0.129)
OALAND	0.889 (0.117)***	−0.046 (0.152)
URBAN	−0.062 (0.099)	−1.010 (0.183)***
SINDH	−0.976 (0.112)***	−0.153 (0.189)
NWFP	−1.618 (0.149)***	0.874 (0.167)***
BALOCHISTAN	−2.353 (0.226)***	0.993 (0.224)***
CONSTANT	−4.082 (0.346)***	−14.663 (0.843)***
Pseudo-R^2	0.193	
N	27,544	
Log-likelihood	−4691.0	

Note: *, ** and *** denote significance at the 10, 5 and 1% levels respectively. The dependent variable is categorised as: (1) = participation in 'Other' where other constitutes non-labour force participation, unemployment, livestock ownership or self-employment (agricultural/non-agricultural), (2) = participation in waged work in private sector and (3) = participation in waged work in public sector. Robust standard errors are reported in parentheses and are corrected for clustering at the household-level. (–) Denotes not applicable. NO_EDUCATION and PUNJAB are the reference categories for education splines and province, respectively.

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