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ABSTRACT

A New Method for Estimating Public Sector Pay Premia: Evidence from Britain in the 1990s*

This Paper looks at public sector pay in Britain. We present a novel instrument that exploits the variation in public sector status across individuals arising from the privatisation programme of the 1990s. We show formally that results that are estimated may thereby be robust to self-selection and measurement error, both of which are important and which bias cross-sectional estimates. Our results suggest that women in the public sector earn more then they could elsewhere, whilst, on average, men do not.

JEL Classification: J31
Keywords: measurement error, pay determination and selection effects

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1 Introduction

Many studies using cross section household data suggest that, on average, public sector workers earn more than private sector workers,even controlling for many observable characteristics.[1] The raw differential in Britain is about 17% for men and 28% for women. Controlling for age and education brings it down to 5% for men and 17% for women but these coefficients are still highly significant and robust[2]. The only groups which appear to do worse in the public sector are educated men. In addition researchers have found that pay distribution in the public sector is more compressed than that of the private sector, both within and between educational groups[3]. All comparisons of this sort are, however, unlikely to pick up the true effect of public sector status on pay given that there will be sorting of workers on the basis of tastes and productivities (see Nickell and Quintini, 2002; Borjas, 2002). The aim of this paper, therefore, is to find estimates of the public sector wage premium or penalty that are robust to self selection.

Typically, studies that have addressed the selection issue in this context (and, in similar vein, that of the union "mark-up") have either used selectivity correction or instrumental variable methods on a single cross-section, or exploited variation over time within individual's work histories. The novel feature of the present paper is that it uses both methods. It exploits the privatisations that have taken place over the 1990s in the UK as a source of exogenous variation in public sector status and uses longitudinal data to control for individual unobserved differences that do not change over time. This potentially allows us both to exploit the strengths and control for the weaknesses of both approaches and

[1] Representative studies for the US include Smith (1976, 1977), Moulton (1990) and Hundley (1997); for the UK, Gregory (1990), Disney and Gosling (1998), Blackaby, Murphy and O'Leary (1999), Elliott and Bender (1999), and Henley and Thomas (2001); for Canada, Gunderson (1979) and Gunderson, Hyatt and Riddell (2000). A large number of studies are surveyed in Bender (1998)

[2] source authors; own calculations from the British Household Panel Survey.

[3] Studies that examine pay compression in the public sector typically disaggregate by education and/or use quantile regression methods. They tell a similar story across widely different institutional environments: See Poterba and Rueben (1994) for the US, Disney and Gosling (1998) for Britain, Melly (1998) for Canada, Mueller (2002) for Switzerland, and Nielsen and Rosholm (2001) for Zambia.

so obtain robust estimates of public sector wage premia..

Our findings suggest that, on average, in Britain in the 1990s, men in the public sector earned slightly more than they could in the private sector, but that this difference was not statistically significant. The significance of the levels estimates disappears once we control for selection, endogenous job switching and measurement error. In contrast, women appear to earn significantly more in the public sector Our results show important differences across education groups. For college educated men (with degrees), it is shown that the public sector pay penalty in cross-section estimates may be driven entirely by selection. In first differences, there is a highly significant premium in the public sector. Measurement error drives down the measured premium but endogenous job mobility biases the estimate upwards. However, our findings on college educated women are inconclusive. At the other end of the skill spectrum, however, there is no evidence of a significant public sector pay premium for the least qualified men whatever estimation method is used. Given that the 1990s saw increased emphasis on competitive tendering for public services such as sanitation and various ancillary jobs, this last finding should not be too surprising.

2 Privatisation in the UK

1991/2 23% of working men reported themselves to be working in the public sector, but by 1999/2000 only 17% did so (source BHPS data). These changes came about not through the wholesale movement of nationalised industries into the private sector as most of those privatisations occurred in the 1980s (Green and Haskel, forthcoming) but through smaller (but arguably more far-reaching) changes. Competitive tendering (CCT) and contracting out created a distinction between provision and finance as the state bought in services from the private sector. Many workers, such as cleaners in hospitals and schools, or those working in the civil service became private sector workers even though the state purchased the services they produced. Changes in pay of these workers, relative to the workforce as a whole will give us an robust estimate of the effect of public sector status on wages.

There have been relatively few studies that exploit the quasi-experimental nature of the privatisation programmes of the 1980s and 1990s to look at the effect of ownership on labour market outcomes. Exceptions are Haskel and Szymanski (1993) who use company accounts data to look at the effect of privatisation on wages and employment, and Card (1986) who looked at the effect of deregulation on wages in the airline industry. The only study that we know of that uses micro data for the UK is unpublished work by Dickens and Machin (1998) who examine the big privatisations of the 1980s (gas, telecommunications etc.), and whose results on sectoral average wage growth before and after privatisation are inconclusive. The reason for this lack of analysis of individual workers is that individual level data does not contain direct information on transfer of ownership[4]. Although someone remaining in the same workplace but observed to leave the public sector is very likely to have been affected by a privatisation, they may possibly have mis-recorded their status in either or both time periods. False negatives are also possible as some people observed to remain in the public sector may in fact have left (or never joined) and some will lose/quit their jobs following a transfer of ownership. The presence of these false negatives and positives will bias the results. The results we present strongly indicate the importance of this feature of the data.

In this paper, we argue that changes in the proportion of workers working for the public sector at the occupational level can be used as alternative indicator of changes in public sector status and hence in whether an individual is affected by a privatisation. As the "noise" in this indicator is in part constructed and in part assumed to be orthogonal to the difference between reported and actual changes, it is possible to use an instrumental variable procedure to obtain results that are robust both to self selection and to measurement error. Occupational level changes can be used to infer the likelihood of a change in ownership.

This paper has also implications for the argument, made in the case of the

[4] This is not a problem for the Dickens and Machin study as ownership status(and its change) can be directly infered from narrow industry codes.

union mark up, that cross sectional estimates will be biased upwards due to selection and panel estimates will be biased downwards due to measurement error and endogenous job changes (see for example Freeman 1984 or more recently Swaffield 2000). It is thus argued that longitudinal estimates should be viewed as providing a lower bound to the parameter of interest and cross sectional estimates an upper bound. We suggest that this is not necessarily the case. The bias to the cross section need not be positive and the bias driven by endogenous job changes will only be negative under a certain set of assumptions. The IV procedure we adopt allows us to examine this directly.

3 Modelling the determination of public sector status and wages

Consider the following two equations.

$$w_{it}^{P} = \gamma_i \delta_t^{P} + x_{it}\alpha_t^{P} + \beta_t + \varepsilon_{it}^{P} \tag{1}$$

$$w_{it}^{NP} = \gamma_i \delta_t^{NP} + x_{it}\alpha_t^{NP} + \varepsilon_{it}^{NP} \tag{2}$$

In equations 1 and 2 w_i^J refers to the wage received by individual i if she worked in sector j, γ_i is a set of unobserved wage determining characteristics (e.g. ability) which have prices $\delta^{P(NP)}$ in the public(private) sector and is normalised to have mean zero, x_i are a set of observed wage determining characteristics (e.g. age, education) which have prices α. The εs are idiosyncratic shocks to wages constructed to be uncorrelated with x, or which sector is worked in. β is the average difference in pay across sectors controlling for the xs and the γs. As can be seen all parameters are allowed to vary over time and differ between the two sectors. This very general model is presented to show the restrictions necessary to estimate the parameters of interest (β and the αs)

Cross sectional differences (e.g. those estimated by OLS) yield:

$$E(w^P - w^{NP}|x) = E(\gamma|P)\delta^P - E(\gamma|P)\delta^{NP} + x(\alpha^P - \alpha^{NP}) + \beta$$

so as the γs are not observed the restriction of OLS must be that:

$$E(\gamma|P, x) = E(\gamma|NP, x)$$

Why might this restriction not be valid? For someone to work in the public sector two conditions have to hold:

$$P^*_{it} = g(w^P_{it} - w^{NP}_{it}, Z_{it}) \qquad \text{wants a public sector job} \tag{3}$$

Here Z denotes a set of variables influencing whether someone would chose to work in the public sector even if she was offered the same wage in the private sector.

$$P^{**}_{it} = \eta(w^P_{it} - \bar{w}_{it}, V_{it}) \qquad \text{is offered a public sector job} \tag{4}$$

In equation 4 $\bar{w}$ denotes the productivity of each worker in the public sector and V denotes a set of variables, such as public sector budgetary constraints and privatisation policies, that determine the position of the public sector labour demand curve.

As

$$g'(w^P_{it} - w^{NP}_{it}) > 0, \eta'(w^P_{it} - \bar{w}_{it}) < 0$$

it is likely that $E(\gamma|P, x) \neq E(\gamma|NP, x)$.

Cross section selection models take equations 3 and 4 to construct a reduced form specification where public sector status is related to variables influencing wages, the wage differential in the two sector and to a set of variables in Z and in V. This specification is typically used in the first stage of an simple selection model (see Heckman 1976) or assuming $\alpha^{NP} = \alpha^P$ in a linear probability model in the first step in a two stage least squares instrumental variable procedure. The key problem is that consistent estimation involves finding a observed element of Z or V that is uncorrelated with wages but is also a significant determinant of public sector status. Suitable instruments are rarely available and the allocation of individuals into the public sector is hard to separate out from the allocation of workers into occupations and hence the determination of their wages. In addition the "double hurdle" nature of the model makes this last harder. For example the observed variation in Z may be across individuals of which none are offered a job in the public sector. It is for this reason that findings from existing studies have been inconclusive.[5]

If public sector status has positive premium for at least some people, and the private sector is competitive, then we should see workers queuing for jobs in the public sector (as in Abowd and Farber, 1982 model of union wages). The natural way to proceed then is to use the panel element to condition on $g(.)$ and exploit some exogenous time series variation in $\eta(.)$. This is the rationale for using panel data and exploiting the variation driven by privatisations

3.1 Longitudinal Analyses

Consider restricting equations 1 and 2 so that they collapse to

$$w_{it} = \gamma_i + x_{it}\alpha_t^P P_{it} + x_{it}\alpha_t^{NP}(1 - P_{it}) + \beta_t P_{it} + \nu_{it} \tag{5}$$

where P is a discrete variable taking the value 1 if the individual works in the public sector and $\nu_{it} = \varepsilon_{it}^P P + \varepsilon_{it}^{NP}(1 - P)$. Here the crucial restriction is the on the γs. Equation 5 implies that the effect of unobservables on wages does not change over time or vary between the public and private sectors. The residual ν will be orthogonal to the other terms due to the construction of the εs. We

[5] Studies for an exotic variety of countries correct for selection. They generally either identify off functional form, especially on the education variables, as in Belman and Heywood (1989) for the US, Stelener, van der Gaag and Vijverberg (1989) for Peru, Van der Gaag and Vijverberg (1988) for the Ivory Coast, and van Ophem (1993) for the Netherlands; or off background family characteristics, as in Terrell (1993) for Haiti, Borland, Hirschberg and Lye (1996) for Australia , and Dustman and van Soest (1998) for Germany. From a careful reading of their paper, it is not clear what is the identification strategy of Rees and Shah (1995) for the UK.

obtain:

$$
\begin{aligned}
E(w_{it} - w_{it-1}|x_{it}, x_{it-1}) &= x_{it}\alpha_t^P P_{it} - x_{it-1}\alpha_{t-1}^P P_{it-1} + \\
&\quad x_{it}\alpha_t^{NP}(1 - P_{it}) - x_{it-1}\alpha_{t-1}^{NP}(1 - P_{it-1}) \\
&\quad +\beta_t P_{it} - \beta_{t-1} P_{it-1}
\end{aligned}
$$

For simplicity of exposition, now assume that the αs, β and the xs are stable over time and that $\alpha^P = \alpha^{NP}$. Thus we have

$$
E(w_{it} - w_{it-1}|x_i) = \beta(P_{it} - P_{it-1}) \tag{6}
$$

Equation 6 demonstrates that with the restrictions implied by 5, a regression of changes in wages on changes in public sector status will yield consistent estimates of the βs. It is possible to adjust equation 6 to take account of time varying xs (such as age) or changes in the αs over time and across sectors[6]

There are two reasons why the restrictions implied by 5 may not be valid. First is the possibility that changes in P are endogenous, i.e. if the δs in equation 1 and 2 differ across sectors and vary over time, some of changes in public sector status may be caused by changes in the δs and hence wages. Second as is well known, the presence of measurement error on the changes in P will also introduce biases.

3.1.1 Endogenous switching

The starting point of the analysis is that people choose or are hired to work in the public sector based on characteristics that determine wages that are not observed by the econometrician. In this world a simple conditional correlation of changes in wages on changes in sectoral status will yield consistent estimates of the αs and βs only under the assumptions of equation 5. These restrictions are not testable or particularly plausible. In addition the direction of

[6] Few studies seem to use fixed effects estimators to examine public sector pay - representative studies include Disney and Gosling (1998) for the UK, Pedersen et al (1990) for Scandinavia, and Bales and Rama (2001) for Vietnam.

the bias is unknown. Thus, in order to address this issue, we need to consider why someone might leave or join the public sector in the first place (see Blank, 1985). First people change job and as a consequence move sector. Second the job/workplace/firm might move into (and usually out of) the public sector because of privatisation or competitive tendering. The first sort of moves are likely to result in a correlation between changes in sector and changes in unobservable factors determining wages (such as ability/labour quality). Relating wage changes just to the second sorts of moves (i.e. to those affected by privatisations) should, however, provide us with a cleaner estimate of the effect of sectoral status on wages. This is what we do.

As discussed in the introduction, in the absence of measurement error, a clean comparison would just look at status changes that are not associated with job changes. There are two problems with this strategy. First, there is the likelihood of measurement error on reported status. It is likely that many workers in the workplaces affected by contracting out arrangements will not know whether they are in the public sector or not (we have no external validation of 'true' status as in Card, 1996), or there may be simply random misreporting. Secondly, those who remain in the workplace following a privatisation will not necessarily be a random sample of the workforce. If wages are expected to fall, the better workers might leave to find another job. Moreover, if the public sector disproportionately protects jobs at the expense of productive efficiency, some workers might be sacked. Ideally, we would like to examine the evolution of wages of all workers who worked in the public sector establishment prior to privatisation, but this is not possible in our data, or indeed any other data set. Instead we use an instrumental variable procedure to "cleanse" the changes in public sector status from unobserved worker characteristics associated with moves across workplaces and jobs and so leave simply the changes caused by privitisations.

3.1.2 Measurement Error

In the presence of measurement error but assuming changes in status to be exogenous, the average change in wages amongst the group of workers reporting themselves to have left the public sector between time t and time $t-1$ will be determined by the proportion of people in this group who actually did leave the public sector and by the proportion of people who actually entered into the public sector even if they reported themselves to have left. More formally and generally:

$$\begin{aligned} E(w_{it} - w_{it-1}|P_{it} - P_{it-1} &= j) = \\ \beta[\Pr(P^*_{it} - P^*_{it-1} &= 1|P_{it} - P_{it-1} = j)] \\ -\beta[\Pr(P^*_{it} - P^*_{it-1} &= -1|P_{it} - P_{it-1} = j)] \end{aligned} \tag{7}$$

Where j takes the values -1, 0 or 1 defined respectively as exit from, staying in, and entry to the public sector, and $P_{it}(P^*_{it})$ is reported (actual) status in time t. For $j = -1$, for example, we will only get a consistent estimate of $-\beta$ by looking at the changes in their wages when $\Pr(P^*_{it} - P^*_{it-1} = -1|P_{it} - P_{it-1} = -1) = 1$. Consider now an extreme case where reported status conveys no useful information about actual status (i.e. $\Pr(P^*_{it} - P^*_{it-1} = j|P_{it} - P_{it-1} = j) = \Pr(P^*_{it} - P^*_{it-1} = j)$, then equation 7 will become

$$\begin{aligned} E(w_{it} - w_{it-1}|P_{it} - P_{it-1} &= j) = \\ \beta[\Pr(P^*_{it} - P^*_{it-1} &= 1) - \Pr(P^*_{it} - P^*_{it-1} = -1)] \end{aligned} \tag{8}$$

Each group will then experience the same change in wages and the estimate of β will be zero. The presence of measurement error, so long as it does not introduce a negative correlation between reported and actual status will bias our estimates of β towards zero. It is for this reason that Freeman (1984) and others have argue that first difference estimates of the union mark up provide a lower bound to the true parameter.

An approach which some (such as Swaffield 2000) have used in this context is simply to focus on changes in status where it is believed measurement error problems will be small. In our case this would be people who change job as well as status. The problem with this strategy is that this is the group for which the biases caused by endogenous changes in expected to be largest. When focussing on what happens to a person's wages when they change status without changing job, we get closest to the experiment we wish to conduct i.e. what happens to people's wages when they have been affected by a privatisation but we magnify the importance of measurement error. When we focus purely on those who change job, we minimise measurement error but maximise the potential biases caused by endogeneity of job moves. The aim of this paper is to obtain results that are robust to both problems. This is best done using an instrumental variable procedure, as is now demonstrated.

Equation 7 can be rewritten as

$$E(\Delta w|\Delta P) = \beta[E(\Delta P^*|\Delta P)] \tag{9}$$

If $E(\Delta P^*|\Delta P)$ can be substituted for ΔP in a regression of changes in wages on changes in status, therefore it is possible to obtain a consistent estimate of β that is robust to measurement error.

Now assume there are two indicators of the change in public sector status ΔP^1 and ΔP^2 where

$$\Delta P_i^1 = \Delta P_i^* + \nu_i^1 \text{ and } \Delta P_i^2 = \Delta P_i^* + \nu_i^2 \text{ and } cov(\nu^1, \nu^2) = 0$$

Then we must have

$$E(\Delta P^1|\Delta P^2) = E(\Delta P^*|\Delta P^2) \text{ and } E(\Delta P^2|\Delta P^1) = E(\Delta P^*|\Delta P^1) \tag{10}$$

using the identities in equation 10 we can therefore obtain a consistent estimate of $E(\Delta P^*|\Delta P)$. Using instrumental variables will thus result in estimates that are robust both to measurement error and, providing P^2 is unrelated to job changes, to endogenous switching.

3.1.3 The instruments used to control for measurement error and self selection

The instruments we use are based on occupation and are as follows. First

$$I^1_{it} = \Pr(P_{it} = 1|O_{it}) - \Pr(P_{it-1} = 1|O_{it-1}) \tag{11}$$

where I^1 is the constructed instrument, and O is the current occupation of the worker. Data on $\Pr(P_{it} = 1|O_{it})$ is obtained by merging into the BHPS another data source (the New Earnings Survey), in which information on public sector status is solicited from employers and not employees as is given in the BHPS. Thus noise in I^1 is by factors other than occupation that determine changes in public sector status. Both sources of noise can be assumed to be orthogonal to the errors driven by people simply getting their status wrong. The predicted change in public sector status given I^1 will therefore be consistent. Using this prediction will thus result in estimates of the public sector premium that are robust to measurement error.

The differences in I^1 across individuals are determined by two things. First differences change in relation to the proportion in each occupation that are in the public sector. We argue that this variation is driven mainly by privatisation. Second differences may change when people moving jobs between sectors change occupations, for example I^1 would be positive if someone left a clerical job in the private sector in order to become a public teacher. Estimates using I^1 are therefore not robust to endogenous job changes, unlike measurement error.

To control for endogeneity of job changes, we adopt two strategies. We first restrict our sample simply to those who remain in the same job in each year-to-year comparison. The problem with this is those people whose pay or working conditions are disproportionately affected by privatisation are likely to leave, so that the sample of those who remain in the same job may not be representative.

Our second strategy, therefore, to get as close as we can to looking at all those affected by a privatisation, is simply to condition on occupation in time $t-1$ and ignore any difference in occupation at t.

$$I_{it}^2 = \Pr(P_{it} = 1|O_{it-1}) - \Pr(P_{it-1} = 1|O_{it-1}) \tag{12}$$

I^2 is now orthogonal to changes caused by changes across jobs. By using it as an instrument, we can obtain results that are robust both to measurement error and to self selection. This technique will only work, however, if there is enough variation in I^2 in the data. This is examined in the next section. A comparison of results using first differences, IV using I^1 and IV using I^2 will indicate the relative magnitude and sign of the biases caused by measurement error and self-selection. We can check the result by our alternative comparison of using I^1 on a restricted sample of those people who do not change workplace[7].

Of course any shift in the structure of demand for occupational level skills may also affect I^2 as public sector employment may react slower or differently to technological changes. These shifts may also affect wages. We thus need to make the additional assumption that there has been no change over time in the demand for these skills, *conditional on education*, so that we can interpret all changes in the I as being driven by privatisation.

4 Occupational effects on Public Sector Status

This section demonstrates the role of occupation in determining the extent of, and changes in, public sector status. For many workers, the choice of whether to work in the public sector and the choice of occupation are closely related. Some occupations are completely dominated by the public sector (civil servants, teachers and nurses, for example) and some are mostly in the private sector (such as sales staff). Indeed part of the problem faced by those setting public sector pay, and by those researchers wishing to measure it, is that there are

[7] Although we expect these results to be biased as those who remain in the same job post privatisation are not a random sample of those who were in the workplace in time $t-1$.

very few occupations which are proportionately represented in both sectors. It is therefore hard to compare like with like. This is borne out by the BHPS data, 41% of the variance across individuals in public sector status can be attributed simply to what occupation they are in whereas only 11% which can be explained by age, gender, education and interactions thereof[8].

Table 1: Occupational effects on public sector status

	Variance	Standard Deviation	Prop of Total
Total	0.193	0.440	100.00%
Explained by:			
Differences Across Occupations	0.080	0.282	41.25%
Differences Over Time	0.000	0.019	0.19%
Different Trends of each Occupation	0.003	0.054	1.52%
All Occupational and Time effects	0.083	0.288	42.95%

Notes: source ANOVA decomposition on BHPS data 1991-9, sample is all workers 20-60

Table 1 an analysis of variance (anova) results on BHPS data to demonstrate the role attributable to occupation in determining public sector status. There are two things to note from this table: first the large role played simply by cross sectional differences in occupational status and second,that the role played by different trends across occupations is significant, albeit small. This last finding implies that, in the data, there is variation across occupations in the degree to which they are affected by privatisation at any point in time.

Figure 1 shows the distributions of I^1(instrument 1) and I^2 (instrument 2) by reported change in public sector status. As expected the correlation between I^1 and the reported change in occupational 'public sectorness' is strong. To illustrate with an example, public sector leavers in general, whether or not they change occupation, on average shift to 'less public intensive' occupations, whilst joiners exhibit the reverse. For the second instrument,I^2, (which conditions solely on the initial occupation) the correlation is less marked. For those leaving

[8] source BHPS data

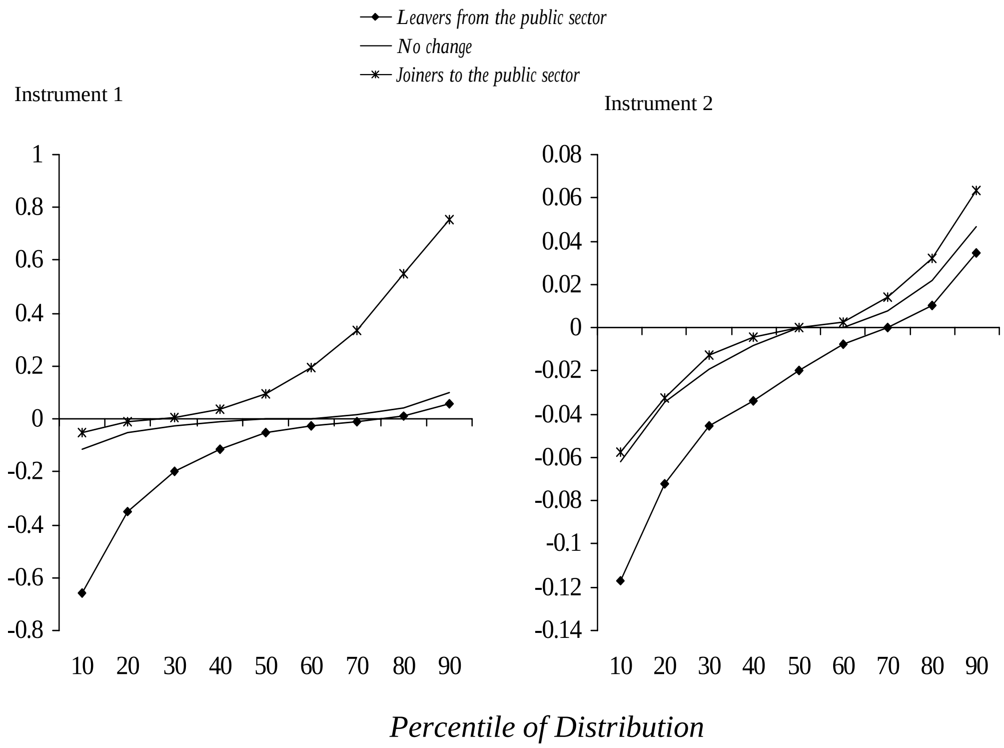
the public sector, there is still a correlation with the change, the distribution of those who remain in the same sector statistically dominates the distribution of those who leave. The difference is less clear cut for those who join compared to those who stay in the public or private sectors. This is as expected, since we expect privatisation to be able to explain why some one might leave the public sector but not why someone might join it[9].

This potential asymmetry between joiners and leavers with I^2 does have two implications for our results. First it suggests that simple IV (which imposes a linear relationship between the instrument and the endogenous independent variable) might be inappropriate [10]. Thus we present results using an ordered probit in the first step as well (i.e. where we define as -1, 0, 1 as the actions of leaving, staying in and joining the public sector). Second it suggests that the return we will estimate will be for those who have chosen (or are selected to) work in the public sector. In the language of the treatment effects literature we will estimate the "effect of treatment on the treated"rather than the average effect (if the δs in equations 1 and 2 are the same, this will of course not matter). We now move on to present our empirical results

[9] This is encouraging in view of the necessary identification assumption that all changes in I are driven by privatisation.

[10] in fact the discrete nature of this variable would suggest that two stage least squares (2SLS) may be inconsistent anyway

Figure 1 Percentiles of instruments 1 and 2 by reported change in sector



Instrument 1 is $\Pr(P_t=1|\text{occupation in time t})- \Pr(P_{t-1}=1|\text{occupation in time t-1})$
Instrument 2 is $\Pr(P_t=1|\text{occupation in time t-1})- \Pr(P_{t-1}=1|\text{occupation in time t-1})$
Source NES and BHPS data

5 Empirical Results

The data we use come from waves 1 to 9 of the British Household Panel Survey. We select a sample of workers aged 20-60. The panel is unbalanced. Information on whether someone changed workplace between waves is not available in the cross section so we used the work history files to construct this measure.

The results we present are based on four analyses

1. Simple "levels" estimation of the public sector wage premium.

2. Regressions of changes in wages on changes in public sector status (first difference results). These would be consistent estimates in the absence of measurement error and endogenous job switching.

3. Two stage least squares instrumental variable estimation of the public sector wage premium. Here the reported change in public sector status is regressed against I^1 and I^2 (plus a set of control variables). The predicted value is then used as a proxy for the actual change.

4. Exogeniety tests. Here ordered probits of the change in public sector status using the instruments are estimated and then are used to construct a generalised residual (a measure of the unobservables in the determination of public sector status). These residuals are then included in the regression of changes in wages. This is a simple exogeneity test; under the null that the unobservables determining changes in public sector status do not affect wages, the coefficient on this residual will be zero. As we demonstrate both the IV and the exogeneity test results are very similar.

The detailed regression results are reported in the appendix; the text focuses on the results of general interest.

5.1 The determination of changes in public sector status

Table 2 below demonstrates that for all groups: men overall women overall and each education group, there is a strong association between each of the instruments and the reported change in public sector status. The only coefficient which is not significant at the 5% level is on instrument 1 for the sample of men with no qualifications who remain in the same job in the linear probability case. Instrument 2 is significant for men with no qualifications, however, which suggests that there is some bias to the results caused by unqualified men disproportionately affected by privatisation leaving their jobs.

5.2 The estimated effect of public sector status on pay

Table 3 shows the estimates of the public sector pay premia from the each of the four estimation methods. Looking at men first, it appears that the 5% mark up becomes insignificant once we control for individual characteristics that do not change over time (the first difference results). Interestingly it is the standard error that changes more than the coefficient suggesting that measurement error and/or small sample problems might be important. The third row presents the IV results using instrument 1; these results are robust to measurement error. As expected the coefficient increases more than the standard error. This suggests that men in the public sector earn more than they could elsewhere. Amongst men who remain in the same job but switch sector, however, the estimated return is still not significant, providing evidence that job switching is endogenous to relative pay. The results using instrument 2 (which should generate estimates that are robust both to measurement error and to endogenous job switching) shows the public sector effect to be insignificant. Results using both stayers and instrument 2 therefore suggest that the bias induced by measurement error is cancelled out by the bias caused by endogenous job switching.

The results using the generalised residuals from the ordered probit in the first step are similar, suggesting that the findings are not simply an artefact of imposing a linear model in the first step. The coefficient on the generalised residual using instrument 1 is negative and significant suggesting that measurement error biases the results downwards. The coefficient using instrument 2 is however insignificant suggesting that the bias caused by endogenous moves across jobs is positive and cancels out the bias caused by measurement error. All these findings together suggest that the preferred model is that in first dif-

Table 2: Relationship between instruments and observed change in public sector status;

MEN

	Linear Probability Model			Ordered Probit Model		
	Instrument 1		*Instrument 2*	*Instrument 1*		*Instrument 2*
	All	Stayers	All	All	Stayers	All
All education groups	0.222	0.050	0.228	1.688	0.948	2.392
	(0.028)	(0.025)	(0.075)	(0.274)	(0.365)	(0.702)
Highest Qualification						
Degree or Above	0.227	0.058	0.223	1.908	1.078	2.515
	(0.010)	(0.009)	(0.028)	(0.112)	(0.157)	(0.321)
'A' level type	0.187	0.030	0.240	1.605	0.753	2.667
	(0.017)	(0.015)	(0.051)	(0.213)	(0.319)	(0.556)
'O' level type	0.288	0.094	0.203	2.193	1.299	2.145
	(0.017)	(0.016)	(0.048)	(0.173)	(0.224)	(0.524)
None	0.151	0.044	0.210	2.085	1.616	3.172
	(0.022)	(0.024)	(0.057)	(0.292)	(0.390)	(0.922)

WOMEN

	Linear Probability Model			Ordered Probit Model		
	Instrument 1		*Instrument 2*	*Instrument 1*		*Instrument 2*
	All	Stayers	All	All	Stayers	All
All education groups	0.399	0.070	0.172	2.138	0.668	1.339
	(0.034)	(0.035)	(0.080)	(0.277)	(0.378)	(0.411)
Highest Qualification						
Degree or Above	0.270	0.071	0.230	1.841	0.949	1.545
	(0.010)	(0.010)	(0.033)	(0.092)	(0.140)	(0.248)
'A' level type	0.337	0.081	0.251	1.844	0.800	1.477
	(0.024)	(0.023)	(0.071)	(0.200)	(0.363)	(0.417)
'O' level type	0.228	0.067	0.211	1.691	1.012	1.291
	(0.015)	(0.013)	(0.055)	(0.124)	(0.172)	(0.468)
None	0.247	0.071	0.272	1.980	1.124	2.195
	(0.023	(0.022)	(0.069)	(0.233)	(0.330)	(0.683)

Notes coefficients (standard errors) on instruments are reported, full results given in the appendix ferences. Men in the public sector earn slightly more than they could in the private sector but this difference is not significant.

For women, the story is different. The cross sectional results suggest that women in the public sector earn significantly more than women in the private sector. The first difference results show this to be robust to unobservable differences across individuals that do not change over time - although the coefficient is reduced somewhat, it is still strongly significant. The IV results using instrument 1 suggest that the first difference results are biased downwards (although not significantly so). Similarly the coefficient on the generalised residual is negative but not significant. It appears then that measurement error is less important for women than for men. The IV results using instrument 2 are insignificant but, crucially, the coefficient on the residual using this instrument is both small is size and insignificant. This suggests that the preferred model for women is as just as it is for men, the first difference one. The difference is that women in the public sector, in contrast to men, do appear to earn more than they could in the private sector.

As suggested in the introduction, a number of studies suggest that public sector wage differentials vary across education groups. In levels, it is suggested, pay compression in the public sector tends to benefit less skilled groups. So the average results presented here could be masking differences across these groups. Although we did conduct the analysis for all four of the education groups and the results can be seen in the appendix, we restrict our analysis in the text to some of particular interest: first college educated men and women (with degrees) and secondly, men with no formal qualifications.

For men,with college education, the most striking finding in table 4 is that the cross section result is of an opposite sign to the first difference results. Men with degrees in the public sector earn significantly less then men with degrees in the private sector, but when they leave (join) the public sector they experience a significant wage cut (increase). Thus the apparent negative cross sectional premium is entirely driven by selection. The result using instrument 1 suggests that this positive differential is biased downwards by measurement error but

Table 3: Alternative Estimates of Public Sector Wage Differentials

	Men		Women	
	All workers	Stayers only	All workers	Stayers only
OLS	0.046 (0.016)		0.172 (0.013)	
First differences	0.040 (0.030)	-0.005 (0.026)	0.092 (0.025)	0.005 (0.021)
2 stage least squares on first differences				
instrument 1	0.363 (0.141)	0.585 (0.598)	0.178 (0.099)	0.059 (0.304)
instrument 2	0.267 (0.359)		0.127 (0.286)	
Exogeneity test using generalised residual from ordered probit				
with *instrument 1*				
β(public)	0.418 (0.127)	0.205 (0.271)	0.231 (0.100)	-0.104 (0.234)
β(residual)	-0.164 (0.053)	-0.080 (0.103)	-0.065 (0.043)	0.043 (0.091)
with *instrument 2*				
β(public)	0.357 (0.287)		0.058 (0.297)	
β(residual)	-0.129 (0.122)		0.015 (0.130)	

Notes dependent variable is the change in the log of the hourly wage, coefficient on level or change in public sector status is reported unless otherwise stated. OLS results control for education. See the appendix for the full estimation results. Standard errors in parenthesis

the result using instrument 2 implies that the bias caused by endogenous job moves is biasing the measured differential upwards. Again the first difference results appears to be the preferred ones. Encouragingly, for this interpretation, the coefficient on the generalised residual is small in both sign and significance. The large standard error on the public sector status variable using instrument 2 could thus be driven by the fact that a linear model was imposed in the first step. For women, selection effects seem to work the same way as for men. Measurement error is important. The results using instrument 2 are however inconclusive, suggesting perhaps that there is not enough variation in the data to identify the public sector pay premium for educated women.

The reversal of the implied relationship between wages and public sector status for college educated men implies that public sector workers are negatively selected. Perhaps surprisingly we have also shown that those affected by a privatisation receive higher wages in the public sector. This does not mean that a private sector man picked at random would experience a wage gain on moving into the public sector, however. A zero (or even negative) average premium that would explain our negative selection effects (see also Nickell and Quintini, 2002; Borjas, 2002) is thus not inconsistent with the data.

Table 4: Different Estimates of Public Sector Wage Differentials, those with a degree level qualification or above

	Men		Women	
	All workers	Stayers only	All workers	Stayers only
OLS	-0.070 (0.040)		0.019 (0.041)	
First differences	0.160 (0.064)	-0.025 (0.029)	0.203 (0.089)	0.061 (0.064)
2 stage least squares on first differences				
instrument 1	0.696 (0.312)	0.232 (1.215)	0.768 (0.296)	1.299 (1.168)
instrument 2	0.134 (1.155)		0.617 (0.648)	
Exogeniety test using generalised residual from ordered probit				
with *instrument 1*				
β(public)	0.887 (0.392)	-0.108 (0.409)	0.705 (0.345)	0.944 (0.664)
β(residual)	-0.323 (0.156)	0.033 (0.160)	-0.252 (0.144)	-0.372 (0.277)
with *instrument 2*				
β(public)	0.185 (0.652)		-0.509 (0.572)	
β(residual)	-0.010 (0.271)		0.325 (0.273)	

Notes dependent variable is the change in the log of the hourly wage, coefficient on level or change in public sector status is reported unless otherwise stated. See the appendix for the full estimation results. Standard errors in parenthesis

Turning now to men with no formal qualifications, it can be seen that none of the specifications result in a significant wage gain for working in the public sector. The OLS result is basically zero and each further specification suggests that this is not driven by selection and/or measurement error. This is of no surprise, male manual jobs and the skills required in the public sector are reasonably similar to those done in the private sector and so one would expect the market to equate prices. This finding that college-educated men may get a premium in the public sector whereas less skilled men receive none appears to contradict a number of previous findings on public sector pay.

Table 5: Different Estimates of Public Sector Wage Differentials, those with no formal qualifications

	Men		
	All workers	Stayers only	
OLS	-0.001 (0.038)		
First differences	-0.030 (0.082)	-0.009 (0.038)	
2 stage least squares on first differences			
instrument 1	-0.340 (0.389)	-0.151 (1.026)	
instrument 2	0.325 (0.707)		
Exogeniety test using generalised residual from ordered probit			
with *instrument 1*			
β(public)	-0.038 (0.262)	0.123 (0.306)	
β(residual)	0.003 (0.112)	-0.049 (0.114)	
with *instrument 2*			
β(public)	0.291 (0.410)		
β(residual)	-0.126 (0.158)		

Notes dependent variable is the change in the log of the hourly wage, coefficient on level or change in public sector status is reported unless otherwise stated. See the appendix for the full estimation results. Standard errors in parenthesis

6 Conclusions

In this paper we have presented estimates of the public sector pay premia that are robust both to self selection and measurement error. Overall we find that the biases caused by measurement error (when modeling changes in pay) are largely equal and opposite to those caused by endogenous changes. In the main it appears that longitudinal estimates of the public sector pay premia are consistent. We find that, once we allow for self-selection, there is no evidence that men in the public sector earn more than men in the private sector overall. This aggregate zero differential hides the fact that men with degrees do appear to receive higher wages than they would elsewhere. The negative differential found in the cross section appears to arise from selection. In contrast, men with no educational qualifications do not appear to gain a significant premium in the public sector. On average women in the public sector earn more than they could do in the public sector, but there is no clear result for college-educated women.

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Appendix: Detailed Regression Results

Table a: Regression of Wage Levels

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	7.750	3.970	13.354	12.731	9.146	6.929	7.175	2.971	4.605	1.800
	0.418	*0.398*	*1.595*	*1.614*	*0.908*	*1.089*	*0.720*	*0.601*	*0.816*	*0.766*
Age Squared|	-0.819	-0.450	-1.504	-1.565	-0.998	-0.819	-0.735	-0.341	-0.452	-0.174
	0.056	*0.053*	*0.211*	*0.208*	*0.122*	*0.149*	*0.103*	*0.083*	*0.101*	*0.095*
Degree or Above		0.653								
		0.026								
A level Type Qual	-0.243	0.401								
	0.024	*0.022*								
O level Type Qual	-0.374	0.200								
	0.024	*0.017*								
No Qual	-0.622									
	0.027									
Works in Public Sector	0.046	0.172	-0.070	0.019	0.069	0.212	0.098	0.199	-0.001	0.143
	0.016	*0.013*	*0.040*	*0.041*	*0.028*	*0.029*	*0.026*	*0.020*	*0.038*	*0.026*
Constant	0.744	0.687	-0.276	-0.181	0.242	0.551	0.500	1.071	0.763	1.161
	0.078	*0.072*	*0.288*	*0.278*	*0.159*	*0.181*	*0.118*	*0.101*	*0.156*	*0.144*
Year Dummmies included	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
R squared	0.2497	0.2059	0.1411	0.0943	0.1631	0.126	0.1792	0.0678	0.0688	0.0292
Number of individuals	3885	4190	501	456	1077	902	1442	1763	865	1069
Number of Obs (NxT)	17049	18409	2337	1910	5101	3945	6135	8002	3476	4552

Notes: Dependent Variable is the log of the hourly wage Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up

Table b: Regression of wage changes

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	-1.100	-1.208	-1.363	-2.590	-1.511	-1.697	-1.205	-1.126	-0.899	-1.936
	0.215	*0.205*	*0.784*	*1.033*	*0.322*	*0.461*	*0.403*	*0.296*	*0.548*	*0.527*
Age Squared	0.104	0.134	0.140	0.294	0.156	0.201	0.118	0.135	0.071	0.198
	0.027	*0.026*	*0.097*	*0.136*	*0.041*	*0.062*	*0.054*	*0.038*	*0.063*	*0.061*
P_t-$P_{(t-1)}$	0.040	0.092	0.160	0.203	0.060	0.103	-0.003	0.047	-0.030	0.107
	0.030	*0.025*	*0.064*	*0.089*	*0.065*	*0.048*	*0.042*	*0.039*	*0.082*	*0.036*
Constant	0.269	0.305	0.362	0.625	0.345	0.359	0.254	0.261	0.268	0.500
	0.041	*0.039*	*0.151*	*0.188*	*0.061*	*0.082*	*0.071*	*0.057*	*0.124*	*0.108*
Year Dummmies included	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
R squared	0.010	0.008	0.014	0.027	0.014	0.013	0.014	0.006	0.009	0.009
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in log of the hourly wage Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. P_t is public sector status in time t

Table c: Regression of wage changes restricting the sample to those who remain in their job

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	-0.906	-1.143	-1.065	-3.160	-0.895	-1.683	-1.252	-0.668	-0.576	-2.258
	0.212	*0.225*	*0.863*	*1.168*	*0.289*	*0.527*	*0.346*	*0.283*	*0.559*	*0.687*
Age Squared	0.094	0.127	0.114	0.379	0.090	0.194	0.141	0.073	0.053	0.239
	0.026	*0.028*	*0.105*	*0.153*	*0.035*	*0.069*	*0.043*	*0.035*	*0.064*	*0.077*
P_t-$P_{(t-1)}$	-0.005	0.005	-0.025	0.061	0.047	0.027	-0.034	-0.046	-0.009	0.049
	0.026	*0.021*	*0.029*	*0.064*	*0.062*	*0.046*	*0.043*	*0.032*	*0.038*	*0.030*
Constant	0.224	0.286	0.287	0.739	0.225	0.381	0.254	0.175	0.195	0.549
	0.042	*0.043*	*0.171*	*0.239*	*0.058*	*0.099*	*0.066*	*0.055*	*0.120*	*0.148*
Year Dummmies included	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
R squared	0.005	0.005	0.004	0.018	0.011	0.011	0.009	0.006	0.005	0.006
Number of individuals	2762	2969	364	292	793	641	993	1286	600	740
Number of Obs (NxT)	11376	11700	1642	1187	3494	2434	3962	5070	2237	2966

Notes: Dependent Variable is the change in log of the hourly wage Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. P_t is public sector status in time t

Table d: Regression of changes in public sector status

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	0.060	-0.075	0.406	0.291	0.154	-0.145	0.121	0.092	-0.046	-0.093
	0.112	*0.136*	*0.488*	*0.682*	*0.232*	*0.356*	*0.187*	*0.202*	*0.205*	*0.263*
Age Squared	-0.009	0.005	-0.059	-0.057	-0.023	0.010	-0.016	-0.015	0.007	0.012
	0.014	*0.017*	*0.061*	*0.085*	*0.030*	*0.047*	*0.025*	*0.027*	*0.025*	*0.031*
Instrument 1	0.227	0.270	0.222	0.339	0.187	0.337	0.288	0.228	0.151	0.247
	0.010	*0.010*	*0.028*	*0.034*	*0.017*	*0.024*	*0.017*	*0.015*	*0.022*	*0.023*
Constant	-0.008	0.020	-0.096	-0.048	-0.015	0.038	-0.006	-0.012	0.003	0.021
	0.021	*0.026*	*0.095*	*0.132*	*0.044*	*0.064*	*0.034*	*0.037*	*0.041*	*0.054*
Year Dummmies included	yes	yes	yes	yes	yes	Yes	yes	yes	yes	yes
R squared	0.039	0.046	0.039	0.076	0.030	0.063	0.063	0.039	0.021	0.035
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in public sector status. OLS results reported. Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 1 expected change in public sector status given occupation in t and t-1

Table e: IV estimates of Wage Changes using instrument 1

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.363	0.178	0.696	0.768	0.591	0.176	0.268	0.068	-0.340	-0.080
	0.141	*0.099*	*0.312*	*0.296*	*0.343*	*0.151*	*0.176*	*0.141*	*0.389*	*0.320*
Age	-1.124	-1.200	-1.662	-2.595	-1.611	-1.685	-1.232	-1.128	-0.905	-1.943
	0.218	*0.205*	*0.844*	*1.114*	*0.345*	*0.460*	*0.409*	*0.295*	*0.549*	*0.529*
Age Squared	0.011	0.013	0.018	0.031	0.017	0.020	0.012	0.014	0.007	0.020
	0.003	*0.003*	*0.011*	*0.015*	*0.004*	*0.006*	*0.006*	*0.004*	*0.006*	*0.006*
Constant	0.273	0.303	0.429	0.616	0.356	0.356	0.327	0.211	0.267	0.485
	0.042	*0.039*	*0.164*	*0.204*	*0.064*	*0.082*	*0.076*	*0.059*	*0.125*	*0.111*
Year Dummmies included	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in log of the hourly wage Standard errors adjusted for repeated sampling of the same individual and for the inclusion of a generated regressor in italics. Estimates are from the BHPS. Age and Age squared coefficients scaled up. P_t-$P_{(t-1)}$ is instrumented by instrument 1 (expected change in public sector status given occupation in t and t-1)

Table f: Regression of changes in public sector status restricting the sample to those who remain in their job

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	0.090	-0.048	0.382	0.710	0.153	-0.005	0.083	-0.056	0.058	-0.038
	0.097	*0.116*	*0.403*	*0.620*	*0.191*	*0.301*	*0.168*	*0.171*	*0.188*	*0.229*
Age Squared	-0.013	0.006	-0.052	-0.092	-0.022	-0.003	-0.011	0.006	-0.006	0.007
	0.012	*0.015*	*0.050*	*0.077*	*0.024*	*0.039*	*0.022*	*0.022*	*0.022*	*0.027*
Instrument 1	0.058	0.071	0.050	0.070	0.030	0.081	0.094	0.067	0.044	0.071
	0.009	*0.010*	*0.025*	*0.035*	*0.015*	*0.023*	*0.016*	*0.013*	*0.024*	*0.022*
Constant	-0.014	0.010	-0.087	-0.163	-0.024	-0.005	-0.002	0.013	-0.012	0.002
	0.019	*0.022*	*0.079*	*0.123*	*0.036*	*0.056*	*0.031*	*0.032*	*0.038*	*0.048*
Year Dummmies included	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
R squared	0.004	0.005	0.006	0.012	0.003	0.007	0.011	0.006	0.004	0.005
Number of individuals	2762	2969	364	292	793	641	993	1286	600	740
Number of Obs (NxT)	11376	11700	1642	1187	3494	2434	3962	5070	2237	2966

Notes: Dependent Variable is the change in public sector status. OLS results reported. Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 1 expected change in public sector status given occupation in t and t-1

Table g: IV estimates of Wage Changes using instrument 1 but restricting the sample to those who remain in their job

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.585	0.060	0.232	1.299	2.297	-0.253	0.215	-0.135	-0.151	0.439
	0.598	*0.304*	*1.215*	*1.168*	*3.172*	*0.765*	*0.434*	*0.316*	*1.026*	*1.106*
Age	-0.963	-1.140	-1.172	-3.998	-1.250	-1.679	-1.272	-0.673	-0.566	-2.246
	0.224	*0.225*	*1.008*	*1.668*	*0.683*	*0.521*	*0.347*	*0.280*	*0.571*	*0.691*
Age Squared	0.102	0.127	0.128	0.487	0.141	0.193	0.144	0.073	0.052	0.236
	0.028	*0.028*	*0.127*	*0.216*	*0.094*	*0.068*	*0.043*	*0.035*	*0.065*	*0.078*
Constant	0.234	0.285	0.312	0.938	0.289	0.410	0.254	0.176	0.193	0.548
	0.044	*0.043*	*0.203*	*0.366*	*0.118*	*0.093*	*0.066*	*0.055*	*0.122*	*0.148*
Year Dummmies included	yes	yes	yes	Yes	yes	yes	yes	yes	yes	yes
R squared										
Number of individuals	2762	2969	364	292	793	641	993	1286	600	740
Number of Obs (NxT)	11376	11700	1642	1187	3494	2434	3962	5070	2237	2966

Notes: Dependent Variable is the change in log of the hourly wage Standard errors adjusted for repeated sampling of the same individual and for the inclusion of a generated regressor in italics. Estimates from the BHPS. Age and Age squared coefficients scaled up. P_t-$P_{(t-1)}$ is instrumented by instrument 1 (expected change in public sector status given occupation in t and t-1)

Table h Regression of changes in public sector status

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	0.070	-0.092	0.549	0.021	0.207	-0.177	0.089	0.105	-0.049	-0.041
	0.114	*0.139*	*0.495*	*0.704*	*0.235*	*0.367*	*0.192*	*0.206*	*0.207*	*0.267*
Age Squared	-0.010	0.006	-0.077	-0.025	-0.030	0.013	-0.012	-0.017	0.007	0.007
	0.015	*0.018*	*0.062*	*0.088*	*0.030*	*0.048*	*0.026*	*0.027*	*0.025*	*0.031*
Instrument 2	0.223	0.230	0.228	0.172	0.240	0.251	0.203	0.211	0.210	0.272
	0.028	*0.033*	*0.075*	*0.080*	*0.051*	*0.071*	*0.048*	*0.055*	*0.057*	*0.069*
Constant	-0.010	0.025	-0.122	0.012	-0.025	0.050	-0.001	-0.014	0.006	0.011
	0.022	*0.026*	*0.096*	*0.136*	*0.044*	*0.066*	*0.035*	*0.038*	*0.041*	*0.055*
Year Dummmies included	yes	yes	yes	Yes	Yes	yes	yes	yes	yes	yes
R squared	0.005	0.004	0.011	0.014	0.007	0.006	0.007	0.003	0.009	0.007
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in public sector status. OLS results reported. Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 2 is expected change in public sector status given occupation in t-1

Table i: IV estimates of Wage Changes using instrument 2

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.267	0.127	0.134	0.617	0.424	0.533	0.220	-0.340	0.325	-0.020
	0.359	*0.286*	*1.155*	*0.648*	*0.572*	*0.425*	*0.650*	*0.611*	*0.707*	*0.740*
Age	-1.117	-1.205	-1.349	-2.594	-1.579	-1.623	-1.227	-1.084	-0.892	-1.941
	0.217	*0.206*	*1.023*	*1.086*	*0.358*	*0.468*	*0.402*	*0.307*	*0.551*	*0.530*
Age Squared	0.107	0.133	0.138	0.304	0.166	0.196	0.121	0.128	0.070	0.199
	0.027	*0.026*	*0.135*	*0.149*	*0.047*	*0.063*	*0.054*	*0.040*	*0.063*	*0.062*
Constant	0.272	0.304	0.359	0.618	0.352	0.339	0.326	0.207	0.269	0.485
	0.041	*0.039*	*0.205*	*0.195*	*0.063*	*0.085*	*0.074*	*0.061*	*0.126*	*0.111*
Year Dummmies included	yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
R squared										
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in log of the hourly wage Standard errors adjusted for repeated sampling of the same individual in italics and for the inclusion of a generated regressor estimates from the BHPS. Age and Age squared coefficients scaled up. P_t-$P_{(t-1)}$ is instrumented by instrument 2 (expected change in public sector status given occupation in t –1)

Table j: Ordered Probit Esimates of Changes in Public Sector Status

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Instrument 1	1.908	1.841	1.688	2.138	1.605	1.844	2.193	1.691	2.085	1.980
	0.112	*0.092*	*0.274*	*0.277*	*0.213*	*0.200*	*0.173*	*0.124*	*0.292*	*0.233*
Age	-0.229	-1.091	2.760	-2.310	0.427	-0.083	-0.537	-0.635	1.309	-1.005
	0.947	*0.764*	*3.654*	*3.685*	*1.787*	*1.689*	*1.554*	*1.175*	*2.349*	*1.578*
Age Squared	0.001	0.094	-0.369	0.165	-0.103	-0.052	0.046	0.034	-0.157	0.116
	0.122	*0.097*	*0.448*	*0.469*	*0.223*	*0.225*	*0.209*	*0.152*	*0.293*	*0.193*
Threshold 1	-2.326	-2.314	-1.788	-2.522	-2.176	-2.009	-2.551	-2.245	-1.972	-2.419
	0.184	*0.148*	*0.705*	*0.740*	*0.345*	*0.315*	*0.286*	*0.226*	*0.463*	*0.321*
Threshold 2	2.165	1.754	2.543	1.489	2.267	1.890	1.964	1.835	2.800	1.869
	0.185	*0.147*	*0.733*	*0.741*	*0.340*	*0.313*	*0.286*	*0.226*	*0.482*	*0.327*
Year Dummmies included	Yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
Log Likelihood	-2393.2	-3806.5	-384.5	-435.3	-720.6	-926.4	-887.0	-1652.5	-372.5	-756.8
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in public sector status. (-1 = leave, 0 = stay in public or stay in private sectors 1 = join public sector) Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 1 expected change in public sector status given occupation in t and t-1

Table k: Exogeniety Test Results using instrument 1

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.418	0.231	0.887	0.705	0.646	0.274	0.222	0.092	-0.038	-0.056
	0.127	*0.100*	*0.392*	*0.345*	*0.243*	*0.135*	*0.154*	*0.142*	*0.262*	*0.208*
Generalised Residual	-0.164	-0.065	-0.323	-0.252	-0.249	-0.083	-0.103	-0.021	0.003	0.072
	0.053	*0.043*	*0.156*	*0.144*	*0.110*	*0.063*	*0.063*	*0.061*	*0.112*	*0.091*
Age	-1.095	-1.185	-1.722	-2.276	-1.511	-1.679	-1.186	-1.123	-0.898	-1.942
	0.214	*0.206*	*0.816*	*1.047*	*0.321*	*0.463*	*0.402*	*0.297*	*0.553*	*0.527*
Age Squared	0.105	0.132	0.188	0.264	0.159	0.200	0.116	0.135	0.071	0.199
	0.027	*0.026*	*0.102*	*0.137*	*0.041*	*0.062*	*0.054*	*0.038*	*0.063*	*0.061*
Constant	0.267	0.299	0.444	0.558	0.337	0.354	0.247	0.260	0.268	0.503
	0.041	*0.039*	*0.159*	*0.192*	*0.060*	*0.083*	*0.071*	*0.057*	*0.126*	*0.108*
Year Dummmies included	yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
R squared	0.012	0.008	0.022	0.038	0.019	0.014	0.015	0.006	0.009	0.009
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in log of the hourly wage. Under the null that the unobservables in the ordered probit do not effect wage changes the coefficient on the generalised residual should be zero. Standard errors adjusted for repeated sampling of the same individual but not for the inclusion of a generated regressor in italics. Eestimates are from the BHPS. Age and Age squared coefficients scaled up.

Table l: Ordered Probit Esimates of Changes in Public Sector Status

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Instrument 1	1.078	0.949	0.948	0.668	0.753	0.800	1.299	1.012	1.616	1.124
	0.157	*0.140*	*0.365*	*0.378*	*0.319*	*0.363*	*0.224*	*0.172*	*0.390*	*0.330*
Age	0.627	-0.778	0.739	3.704	-0.335	0.089	-0.006	-0.829	4.659	0.711
	1.342	*1.162*	*4.590*	*4.820*	*2.558*	*2.247*	*2.031*	*1.910*	*3.380*	*2.717*
Age Squared	-0.122	0.093	-0.119	-0.433	-0.034	-0.060	-0.024	0.085	-0.569	-0.038
	0.168	*0.144*	*0.561*	*0.611*	*0.309*	*0.285*	*0.263*	*0.244*	*0.422*	*0.307*
Threshold 1	-2.194	-2.315	-2.252	-1.341	-2.538	-2.071	-2.569	-2.432	-1.581	-2.118
	0.269	*0.234*	*0.895*	*0.898*	*0.519*	*0.431*	*0.385*	*0.366*	*0.630*	*0.612*
Threshold 2	2.539	2.110	2.390	2.851	2.206	2.238	2.131	2.036	3.443	2.526
	0.273	*0.233*	*0.910*	*0.920*	*0.502*	*0.434*	*0.388*	*0.362*	*0.654*	*0.624*
Year Dummmies included	Yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
Log Likelihood	-1443.432	-1989.007	-234.248	-260.452	-420.778	-461.922	-542.240	-837.910	-228.689	-408.834
Number of individuals	2762	2969	364	292	793	641	993	1286	600	740
Number of Obs (NxT)	11376	11700	1642	1187	3494	2434	3962	5070	2237	2966

Notes: Dependent Variable is the change in public sector status. (-1 = leave, 0 = stay in public or stay in private sectors 1 = join public sector) Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 1 expected change in public sector status given occupation in t and t-1

Table m: Exogeniety Test Results using instrument 1 but restricted sample to those who remain in the same job

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.205	-0.104	-0.108	0.944	0.741	-0.256	0.238	-0.287	0.123	0.614
	0.271	*0.234*	*0.409*	*0.664*	*0.843*	*0.440*	*0.246*	*0.245*	*0.306*	*0.555*
Generalised Residual	-0.080	0.043	0.033	-0.372	-0.261	0.116	-0.107	0.096	-0.049	-0.217
	0.103	*0.091*	*0.160*	*0.277*	*0.307*	*0.183*	*0.099*	*0.093*	*0.114*	*0.212*
Age	-0.913	-1.150	-1.052	-3.404	-0.867	-1.685	-1.251	-0.684	-0.599	-2.287
	0.212	*0.225*	*0.868*	*1.214*	*0.284*	*0.527*	*0.345*	*0.281*	*0.567*	*0.688*
Age Squared	0.095	0.128	0.112	0.407	0.089	0.193	0.142	0.074	0.056	0.241
	0.026	*0.028*	*0.105*	*0.158*	*0.035*	*0.069*	*0.043*	*0.035*	*0.065*	*0.077*
Constant	0.225	0.287	0.284	0.823	0.215	0.379	0.250	0.179	0.200	0.557
	0.042	*0.043*	*0.172*	*0.260*	*0.057*	*0.098*	*0.066*	*0.055*	*0.121*	*0.148*
Year Dummmies included	yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
R squared	0.005	0.005	0.004	0.020	0.011	0.011	0.010	0.006	0.005	0.007
Number of individuals	2762	2969	364	292	793	641	993	1286	600	740
Number of Obs (NxT)	11376	11700	1642	1187	3494	2434	3962	5070	2237	2966

Notes: Dependent Variable is the change in log of the hourly wage. Under the null that the unobservables in the ordered probit do not effect wage changes the coefficient on the generalised residual should be zero. Standard errors adjusted for repeated sampling of the same individual but not for the inclusion of a generated regressor in italics. Eestimates are from the BHPS. Age and Age squared coefficients scaled up.

Table n: Ordered Probit Esimates of Changes in Public Sector Status

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Instrument 2	2.515	1.545	2.392	1.339	2.667	1.477	2.145	1.291	3.172	2.195
	0.321	*0.248*	*0.702*	*0.411*	*0.556*	*0.417*	*0.524*	*0.468*	*0.922*	*0.683*
Age	-0.475	-1.616	4.776	-2.086	0.401	-0.541	-1.650	-1.271	1.728	-0.509
	0.942	*0.775*	*3.689*	*3.742*	*1.743*	*1.681*	*1.538*	*1.170*	*2.406*	*1.853*
Age Squared	0.025	0.150	-0.644	0.149	-0.105	-0.012	0.185	0.106	-0.208	0.065
	0.120	*0.097*	*0.459*	*0.467*	*0.217*	*0.220*	*0.207*	*0.150*	*0.295*	*0.220*
Threshold 1	-2.305	-2.332	-1.326	-2.316	-2.136	-2.048	-2.588	-2.271	-1.868	-2.252
	0.183	*0.152*	*0.714*	*0.757*	*0.332*	*0.320*	*0.276*	*0.226*	*0.481*	*0.381*
Threshold 2	2.005	1.533	2.867	1.405	2.183	1.648	1.655	1.615	2.776	1.871
	0.186	*0.150*	*0.737*	*0.756*	*0.335*	*0.314*	*0.280*	*0.224*	*0.487*	*0.381*
Year Dummmies included	Yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
Log Likelihood	-2599.111	-4144.864	-409.243	-490.610	-761.774	-1013.714	-1006.786	-1787.066	-392.900	-811.998
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in public sector status. (-1 = leave, 0 = stay in public or stay in private sectors 1 = join public sector) Standard errors adjusted for repeated sampling of the same individual in italics, estimates from the BHPS. Age and Age squared coefficients scaled up. Instrument 2 is expected change in public sector status given occupation in t-1

Table o: Exogeniety Test Results using instrument 2

	All Ed Groups		Degree or Above		A level type Qual		O level type Qual		No Qual	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
P_t-$P_{(t-1)}$	0.357	0.058	0.185	-0.509	0.573	0.876	0.405	-0.453	0.291	0.032
	0.287	*0.297*	*0.652*	*0.572*	*0.649*	*0.452*	*0.413*	*0.601*	*0.410*	*0.534*
Generalised Residual	-0.129	0.015	-0.010	0.325	-0.209	-0.349	-0.168	0.217	-0.126	0.032
	0.122	*0.130*	*0.271*	*0.273*	*0.282*	*0.203*	*0.173*	*0.258*	*0.158*	*0.224*
Age	-1.090	-1.215	-1.375	-2.812	-1.514	-1.630	-1.162	-1.197	-0.936	-1.940
	0.215	*0.211*	*0.855*	*0.987*	*0.322*	*0.461*	*0.410*	*0.308*	*0.556*	*0.529*
Age Squared	0.104	0.134	0.142	0.310	0.159	0.201	0.113	0.141	0.076	0.198
	0.027	*0.026*	*0.107*	*0.133*	*0.042*	*0.062*	*0.055*	*0.038*	*0.064*	*0.062*
Constant	0.266	0.307	0.365	0.675	0.339	0.337	0.240	0.280	0.279	0.502
	0.042	*0.041*	*0.166*	*0.179*	*0.061*	*0.083*	*0.074*	*0.061*	*0.126*	*0.109*
Year Dummmies included	yes	yes	yes	yes	Yes	yes	yes	yes	yes	yes
R squared	0.010	0.008	0.014	0.029	0.015	0.014	0.014	0.006	0.009	0.009
Number of individuals	2968	3203	382	313	831	676	1093	1403	650	800
Number of Obs (NxT)	13203	14255	1834	1453	4021	3041	4690	6234	2611	3477

Notes: Dependent Variable is the change in log of the hourly wage. Under the null that the unobservables in the ordered probit do not effect wage changes the coefficient on the generalised residual should be zero. Standard errors adjusted for repeated sampling of the same individual but not for the inclusion of a generated regressor in italics. Eestimates are from the BHPS. Age and Age squared coefficients scaled up.