

The Public Sector Wage Gap : New Evidence from Panel Administrative Data*

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Abstract

With the increase in national debts, public sectors are under high pressure in many countries and the productivity and wage levels of civil servants are under scrutiny. The need for appropriate comparison with their private sector counterparts is stronger than ever. In this study, we suggest novel evidence for France by conducting a comprehensive assessment of the public sector wage gaps throughout the distribution and over a long period (1988-2010). We exploit a large panel of French salary workers drawn from administrative data. We estimate the premia/penalties of the public sector on the unconditional wage distribution while originally accounting for fixed effects and a jackknife correction for potential incidental parameter bias. Results point to a compressing effect of the public sector, i.e. larger public wage premia in the first half of the distribution, which partly reflects the efficient screening operated by public sector entry examinations. However, we find a gradual fall in the public wage gap since 1995, explained by a mix of political and business cycles. Critically, the positive selection into the public sector, particularly strong at lower quantiles, has faded away. A decline in the relative quality of the public workforce is possibly explained by the long-term degradation of the public wage gap itself and by the acceleration of job openings through less-selective recruitment schemes.

Key Words: public wage gap, unconditional quantile regression, fixed effects, incidental parameter bias, jackknife.

JEL Classification: J31, C14

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1 Introduction

European countries are strongly pressured to consolidate their public finances in the wake of the 2008 financial crisis. As a result, the productivity and wage levels of civil servants are under scrutiny. Real or even nominal wage cuts in the public sector have been observed in several countries (see Depalo et al., 2015). In France, in particular, public sector wages are nominally frozen since 2010. Arguably, excessive levels of public wages can hinder competitiveness (by crowding out employment from the private sector, increasing public budget deficits or diverting public spendings from productive services). Yet, inversely, large cuts in public wages may pose a threat to the quality of public services by making it difficult to retain skilled workers. In this context, it is highly relevant to precisely and comprehensively measure wage differentials between sectors when controlling for workers' observed and unobserved skills.

Admittedly, the public wage gap has already been widely investigated in the economic literature while controlling for observed characteristics of public versus private workers. Many studies examine the conditional wage gap at the mean or throughout the wage distribution using various techniques like quantile regressions.¹ Notwithstanding, two important limitations characterize the bulk of the literature and motivate our study. First, the presence of unobservables that affect both wage levels and selection into a particular sector require specific treatments. The IV approach applied to quantile regression shows some issues and faces the usual difficulty of finding exogenous instruments.² Panel estimations with fixed effect have seemed a promising approach, in particular their extension to distributional analyses.³ However, a well-known issue is the potential incidental para-

¹The public sector wage gap was estimated in various countries at the mean (Hartog and Oosterbeek, 1993, for the Netherlands, Dustmann and Van Soest, 1997, 1998, for Germany, Lassibille, 1998, for Spain, Borjas, 2002, for the United States, Glinskaya, 2005, for India, Heitmueller, 2005, for Scotland, Boyle et al., 2004, for Ireland, Imbert, 2015, for Vietnam, etc.) and more recently along the distribution using quantile regressions or other distributional analyses on cross-sectional data (for example Disney and Gosling, 1998, for the UK, Mueller, 1998, for Canada, Melly, 2005, for Germany, Lucifora and Meurs, 2006, for Italy, France and the UK, Campos and Pereira, 2008, for Portugal, Foley, 2009, for Ireland, Cai and Liu, 2011, for Australia, Maczulski, 2011, for Finland, Depalo et al., 2015, for euro-area countries).

²Several applications in the literature are based on the extension of quantile estimations to selection correction using IV, following Buchinsky (1998). Yet, Huber and Melly (2015) show that his method correctly works only if the coefficients of quantile regressions are constant across quantiles. Clearly, then, this method cannot be used to study heterogeneous effects. Other methods have been suggested by Abadie et al. (2002) and Chernozhukov and Hansen (2005). Regarding the difficulty to find instruments, note that some studies identify the sector choice using parental background, relying on the fact that workers with civil servant fathers are more likely to belong to the public sector (Bey, 2009, Maczulski, 2011). Yet it is hard to assume that these variables do not also affect potential wages.

³See applications in Campos and Centeno (2012) for 10 European countries using the ECHP data,

meter bias that characterizes short panel estimations of nonlinear models, such as fixed effect quantile regressions (Koenker, 2004, Canay, 2011).

Another difficulty pertains to the fact that distributional analyses most often rely on conditional quantile estimations, which are rather difficult to interpret. In particular, the presence of fixed effects in panel estimations make that estimated wage penalties or premia at different quantiles are to be understood as pay differentials conditional on individual observables and time-invariant unobservables. Recently, several contributions have suggested ways to estimate unconditional quantiles (Firpo et al., 2009, Chernozukhov et al., 2013). To our knowledge, there is hardly any application of these methods to characterize public sector pay gaps at different points of the unconditional wage distribution. Importantly, these approaches could allow including fixed effects in the wage equation without changing the interpretation of the public wage premia/penalties. Thus it would become possible to assess the public wage gap at different points of the unconditional wage distribution when unobserved characteristics are taken into account – and to compare to cases when they are not, in order to assess the contribution of unobservable skills to the raw wage differentials across sectors.⁴

In this paper, we suggest novel evidence for France while making methodological innovations that should address the concerns above. We conduct a comprehensive assessment of the public sector wage gaps on average and throughout the distribution, overall the long period and for detailed points in time. For that purpose, we exploit a large panel, 1/25 of all French salary workers, drawn from administrative data. We estimate unconditional quantile effects following Chernozukhov et al. (2013), accounting for fixed effects in panel estimations. The exceptionally long duration of the panel (22 years, from 1988 to 2010) tends to reduce the incidental parameter bias. More than this, we address this bias – one of the fundamental issues in the econometrics of nonlinear models – in the context of quantile estimations. Precisely, we develop a jackknife correction inspired by Dhaene and Jochmans (2011) that we apply to the estimation of unconditional quantile effects. In our application, we estimate public sector wage gaps overall and for sub-periods, which allows us to relate time trends to political and business cycles. We also analyse the extent to which the evolution of the public pay gap, and other factors pertaining to the selectivity of the public sector recruitment process, may affect the attractiveness of the public sector and the ‘quality’ of civil servants, as measured by the impact of unobservables on wage differentials.

Hospido and Moral-Benito (2016) for Spain, Siminski (2013) for Australia, or an early version of the present work using the French Labor Force Survey in Bargain and Melly (2008).

⁴An important exception to our statement is the study of Hospido and Moral-Benito (2016) who do estimate fixed effects unconditional quantile regressions using a panel for the years 2005-2012 in Spain. This paper does not address the incidental parameter bias.

Results of the fixed effects estimations of unconditional quantile effects point to a public wage premium at the bottom of the distribution and an insignificant gap at the top, i.e. a compressing effect of the public sector on the pooled wage distribution. This result tends to indicate that the French public sector has managed to attract ‘good’ workers in the lower part of the distribution, partly because of non-monetary gains (mission/motivation, job protection) and an efficient national recruitment process. Not correcting for incidental parameter bias tends to overstate the role of unobserved heterogeneity and to hide some of the compression effect. Regarding time changes, we find a slow but continuous fall in the public wage gap after 1995, partly explained by political cycles, specific public sector wage policies and economic downturns. Critically, we also reveal that the positive selection in the public sector has faded away over time. In particular, the large unobserved public premium characterizing the bottom of the distribution has gradually disappeared. This may be explained by several factors including the long-term decline in the wage gap itself and the opening of non-tenured positions and less selective recruitment schemes. We conclude on the fact that public pay policy should not concentrate solely on pay levels but also on the pay-quality mix that best reflects the public interest.

This paper is organised as follows. Section 2 presents the data and describes the raw public wage gap over time and across quantiles. Section 3 explains the empirical strategy and the estimators. Section 4 reports and analyses the results while section 5 concludes.

2 Data

2.1 Datasets

We use a detailed administrative data, the *Panel tous salariés* (PTS), recently provided through secured access by the French national statistic institute INSEE. The data is based on annual compulsory records of employees data (DADS) that all French companies have to fill each year. It is completed by wage records from the public sector. Compared to traditional survey data on wages, like the French Labor Force Survey (FLFS, *Enquêtes Emploi*), registered data basically avoids response errors. With a sampling rate of 1/25, it is also highly representative of all French salary workers (in comparison, the FLFS is sampled at 1/300). A third advantage is the panel dimension, with information on French employees from 1988 to 2010.⁵

A possible drawback often encountered with administrative data is the limited set of

⁵The sample comprises information on all French workers born in October of each even-numbered year over the period. An exception is the year 1990 when no data was collected for DADS due to the workload caused by the collection of the 1990 Census to national statistics services.

relevant variables. We avail of information on age, gender and occupation type. To the extent that workers' other characteristics are broadly time-invariant (like education levels), they should be picked in the fixed effects. Nonetheless, we shall provide additional checks using a match of the PTS with the *Echantillon Démographique Permanent* (EDP), a large survey data collected among a sample of French salary workers and containing information on education (highest diploma obtained), the number of children and the marital status of the employee. As shall be seen in robustness checks, our estimations on the subsample of PTS-EDP matched observations are similar to baseline results, especially when fixed effects are included.

2.2 Sample Selection, Wage and Statistics

Sample Selection, Background information and Wage Construction. PTS contains around 2 million individuals and about 18 million panel observations in total. Our selection goes as follows. The data focuses on employees only, which means that self-employed and farmers are not included. This is not a concern for our purpose since our wage gap measures should homogenously concentrate on salary workers in public versus private sectors. Additionally, we drop military and all persons not counted in the active population or not in work. We also drop workers present less than three times in the panel.⁶ To interpret our results, it is important to know that our baseline includes salary workers of all types of civil services: State/national (*fonction publique d'Etat*), regional/local (*fonction publique territoriale*) and health services (*fonction publique hospitalière*). The large majority, 73%, is composed of tenured servants (*fonctionnaires titulaires*), with unlimited contract and job protection. Non-tenured are recruited on permanent or time-limited contracts and constitute 17.2% of the public sector (15.8%, 19.4% and 16.5% in State, regional and health services respectively). A small fraction corresponds to *stagiaires* (employees in the first year internship following entry in the public sector and that precedes effective tenure).⁷

We construct hourly wages as follows. We focus on the main job (we ignore secondary activities and side-jobs). Since the employment duration variable available for the State civil service is in days (and not in hours), we also drop part-time jobs. We divide net annual earnings, including bonuses, by the number of annual working weeks. We then

⁶These are mainly near base and end year. This selection step does not affect the representativeness of our sample. Moreover, our main results being based on fixed effect estimations, those present only once in the panel are not used for identification of the public wage gap.

⁷Note also that in future work, we shall estimate public wage gap when comparing private workers with either the sole tenured civil servant of the national/state administration or with only the other groups (regional, local and health administrations with various types of contracts).

divide by the statutory work duration for a full-time (35 hours per week).⁸ We exclude individuals with wages in the top and bottom 0.1% in order to avoid outliers and make the comparison across sector more consistent (i.e. avoid categories like traders or highly paid CEOs).

Public Sector Characteristics. After selection, our sample based on PTS data contains around 1.3 million individuals and about 12.2 million panel observations. The PTS-EDP matched sample represents around a sixth of it. Table 1 presents an overview of the characteristics of public and private sector workers in our selected sample. The first row allows comparing log wages across sectors overall or for base and end year data. We observe raw wage gaps of around 14% over 1988-2010 according to PTS (10% according to the subsample of matched PTS+EDP data).

Next, public sector workers are on average 3.5 years older. The main explanation pertains to the fact that the majority of civil servants are tenured through a system of competitive examination at the national level (*concours de la fonction publique*). Many applicants work in the private sector while queuing for public sector tenured positions. Entering the civil sector as a tenured worker may take some years (or may never occurs), depending on the candidate’s performance, which usually improves with experience. Table 1 also shows that more than a half (a third) of the public (private) sector is composed of women. The gender difference is consistent with the occupational distribution within each sector.⁹ Occupation types are three hierarchical positions. They may not be extremely comparable across sectors, yet we observe a relatively larger share of executives and intermediate positions in the public sector. This is partly related to the age differences (public sector workers being older and with more work experience). It also reflects different hierarchical

⁸This statutory working time has changed from 39 to 35 hours starting 2000. For comparability, we use the new standard of 35 hours a week for all years. This simply means that the public wage gap that we measure is essentially the earnings gap between full-time employment in public versus private salary jobs. Using actual changes in work duration would require complicated adjustments given that not all private workers, nor not all public workers, have move to 35 hours at the same time, that the number of effective workweek hours has not necessarily changed given the use of over time in many occupations, etc. In further robustness checks, we will check our results with hourly wages computed with the administratively declared annual number of hours worked, which will restrict the sample to the sectors for which this variable is available (namely private sector and health, regional and local civil services).

⁹In particular, the national education system comprises a majority of female workers. Health and social services also account for a large part of the gender orientation of the public sector. Note that part of the public sector wage gap may somehow reflects the difference in gender wage gaps across sectors. More generally, further work should aim to replicate our estimations on an homogenous industry that comprises both public and private jobs with more comparable occupational structures, for instance the health sector.

structures across sectors: the public sector comprises only three main grades, known as categories A, B and C, which determine the level and the progression of wages in the administration.

As seen in Table 1, the EDP data broadly confirms the sector differences in age, gender and occupation compositions. It additionally brings some information about education, experience and family status. Experience gaps follow very closely the age gap between sectors. As expected, the public sector is characterized by much higher education levels. This pertains to the type of jobs often encountered in the administration (education and health professionals, civil engineering, etc.) and is consistent with the lower rate of blue collars. It is also related to the *concours* system to enter the tenured civil sector, as indicated above. Indeed, eligibility to this examination is granted on the basis of minimum requirements regarding highschool diploma and university degrees. Finally, marriage is more frequent among public sector workers, which is mechanically related to the age difference.

The lower part of Table 1 shows that public sector employees represent around a quarter of our samples. In the larger PTS data, we follow individuals for almost 13 years on average. Importantly for fixed effects estimations, which rely on transitions across sectors, we find a substantial number of movers in the data. Around 5% of all the individuals present in the panel, around 66,000 persons, have transited across sectors at least once. The fraction of moves from private to public sectors is larger than the reverse transition, yet both are represented and can contribute to identify the public sector wage gaps and, in extension, the gap associated with a particular direction of the move. We investigate this point in the section dedicated to robustness checks.

Raw Wage Gaps. The upper graph of Figure 1 depicts the time change in hourly wage in both sectors and at different percentiles (10th, 50th and 90th) of the pooled wage distribution. There is a clear average raw gap in favor of the public sector at all levels and all points in time. The trends are relatively parallel for low wages while the raw public premium tends to increase then decrease at higher levels. The lower graph confirms these observations by showing the raw difference in log wages across sectors. The raw gap is very low for top workers (between 2% and 9% over the period), more substantial at median and low wages (it oscillates between 12% and 20%). The evolution consists of a rising raw gap until the mid-1990s for all wage levels, followed by a decline for the 10th percentile versus a plateau then a decline starting in the early 2000s for the median and upper wages (the decline is almost continuous until 2010, with the exception of a temporary rebound coinciding with the 2008-09 crisis for the low wages). These trends are partly consistent with the public wage policies of the period and, particularly, with three policy

Table 1: Descriptive Statistics : PTS and PTS+EDP

	Panel Tous Salariés (PTS)						PTS + Echantillon Démographique Permanent (EDP)					
	Base year 1988		End year 2010		All years		Base year 1988		End year 2010		All years	
	Public	Private	Public	Private	Public	Private	Public	Private	Public	Private	Public	Private
Log wage rate (1)	2.12 (0.22)	2.00 (0.27)	2.19 (0.23)	2.08 (0.26)	2.18 (0.24)	2.04 (0.27)	2.45 (0.21)	2.35 (0.27)	2.46 (0.21)	2.44 (0.26)	2.50 (0.23)	2.40 (0.27)
<i>Demographics</i>												
Age	38.57	35.72	41.93	38.70	40.67	37.18	38.39	35.65	44.14	41.23	41.51	38.45
Female	53%	32%	61%	38%	59%	36%	54%	33%	61%	37%	60%	35%
Married							47%	42%	46%	38%	50%	42%
Married with children							39%	34%	39%	32%	42%	35%
<i>Experience and Occupation</i>												
Potential Experience (2)							19.55	16.13	25.66	22.85	22.64	19.69
Executive (3)	11%	5%	15%	8%	20%	7%	9%	4%	21%	8%	18%	7%
Intermediate professions (4)	44%	23%	31%	23%	37%	24%	43%	22%	27%	24%	33%	25%
Employees and workers (5)	45%	71%	54%	69%	48%	68%	48%	73%	51%	68%	49%	67%
<i>Education</i>												
Secondary education							10%	7%	7%	7%	8%	7%
Vocational training							20%	36%	26%	34%	25%	34%
High school							21%	11%	17%	19%	17%	16%
University first degree							18%	8%	18%	14%	19%	12%
University upper degrees							11%	4%	19%	7%	15%	7%
# panel observations	12,185,621						2,050,464					
# workers	1,328,835						248,186					
share of public sector	24%						23%					
average # period observed by workers	12.97						8.26					
% of individuals with at least one move overall	5.02%						3.79%					
average % moving per year	0.73%						0.59%					
average % moving Pub -> Priv per period	0.27%						0.2%					
average % moving Priv -> Pub per period	0.46%						0.39%					

Statistics in this table describe our selection of salary workers in the private and public sectors that are present at least in 4 waves of the panel over 1988-2010.

(1) PTS and PTS+EDP data on French salary workers' wages in 2002 euros, include bonuses and premium; standard errors in brackets.

(2) Potential experience, in # of years, is defined as the difference between present age and age at last diploma graduation.

(3) Includes administrative, commercial or technical executives, professors and 'higher intellectual professions'.

(4) Includes intermediary positions in commercial, technical and administrative sectors, health services, teachers, technicians.

(5) Commercial, technical and administrative employees and clerks

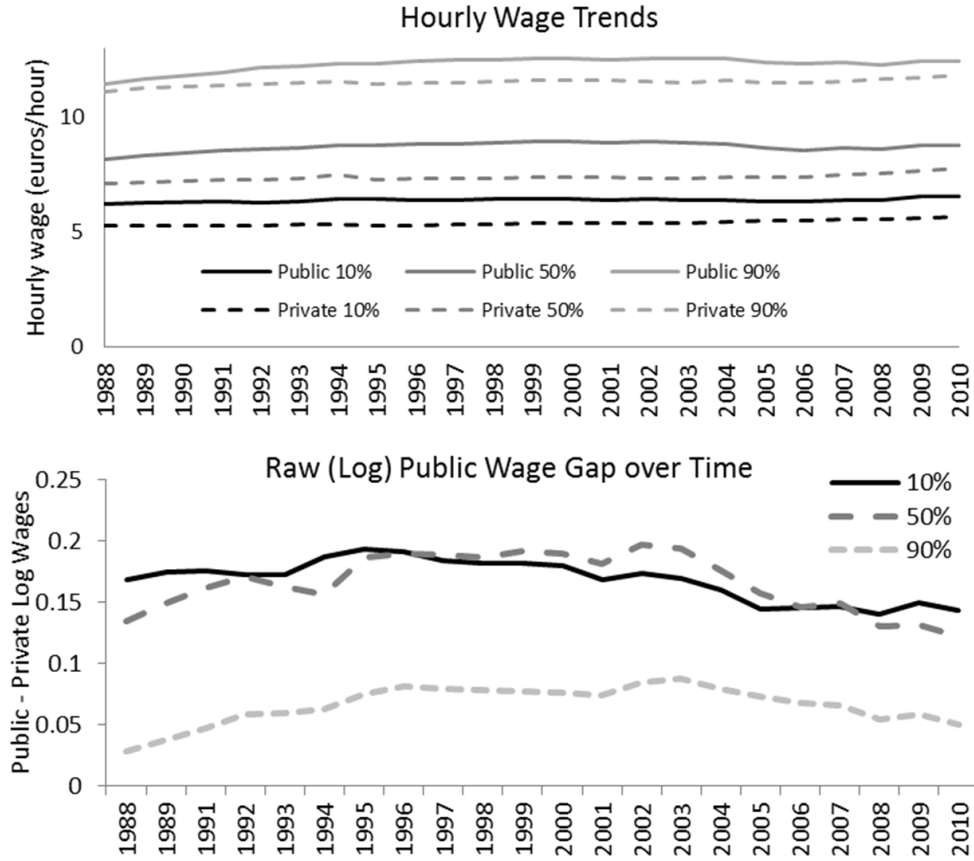


Figure 1: Evolution of the Raw Public Wage Gap by Quantile

plans (1990, 1993 and 1995) that aimed to boost public sector remuneration (plans that terminated in 2000). Better correlations with actual policy measures are expected to be seen once wage gaps are cleaned from essential workers' differences across sectors, as we shall suggest hereafter.

3 Econometric Methods

3.1 Fixed Effects Quantile Regression

We solve the potential endogeneity problem of the employment sector using a fixed effects approach. Since we are interested in the difference between the whole public and private sector wage distributions, we use quantile regression methods. Recently, there has been an active literature about the estimation of quantile models in the presence of fixed effects. We succinctly summarize the different approaches and explain why we chose one in particular.

One of the main issue of fixed effects quantile regression is the incidental parameter problem first discussed in Neyman and Scott (1948). There is no transformation of the data that can remove the dependence on the fixed effects. First-differencing or de-meaning works for linear mean regression due to the linearity of the expectation operator but does not work for linear quantile regression. Thus, any estimator will be a function of the estimated fixed effects, which cannot be estimated consistently when the number of periods is finite. This is a serious issue because, in most cases, the number of periods is limited while the number of units is large. In our application, with more than 1 millions individuals observed on average during 13 periods with a maximum of 22 periods, we must and can take this potential bias into account.

Let Y_{it} denote the outcome (log wage) for observations $i \in \{1, 2, \dots, n\}$ in period $t \in \{1, 2, \dots, T\}$. We also observe a vector of regressors X_{it} and the public sector indicator variable S_{it} . Several fixed effects quantile models and estimators have been suggested. Koenker (2004) assumes that the individual fixed effects α_i only shift the conditional distribution of the outcome without changing its shape:

$$Q_{Y_{it}}(\theta | X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i) = X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta) + \alpha_i, \quad (1)$$

where $Q_Y(\theta | X)$ is the θ^{th} conditional quantile of Y given X for some $0 < \theta < 1$. This is a linear quantile regression model as introduced by Koenker and Bassett Jr (1978) with individual fixed effects as additional regressors. In model (1), the fixed effects are treated differently from the other regressors: they are constrained to be the same at all quantiles. This may seem to be an unnatural assumption in a quantile regression setting where the goal is often to analyze the heterogeneity of the effects. On the other hand, it considerably reduces the dimension of the problem since we have to estimate only one fixed effect for all quantiles instead of a whole distribution of fixed effects per individual.

Koenker (2004) suggests to impose the cross-quantile restrictions on the fixed effects by estimating jointly several quantile regressions.¹⁰ This very large problem can nevertheless been solved in a reasonable amount of time by exploiting the sparse structure of the matrix of regressors. Canay (2011) suggests an alternative 2-step estimator. He notes that assuming (1) for all $\theta \in (0, 1)$ implies :

$$\begin{aligned} E[Y_{it} | X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i] &= \int_0^1 (X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta) + \alpha_i) d\theta \\ &= X'_{it} \int_0^1 \beta(\theta) d\theta + S_{it} \cdot \int_0^1 \gamma(\theta) d\theta + \alpha_i \\ &\equiv X'_{it}\bar{\beta} + S_{it} \cdot \bar{\gamma} + \alpha_i \end{aligned}$$

¹⁰He also considers shrinking the individual effects toward a common value to reduce the dimensionality of the problem and the variance of the estimates. We do not pursue this approach here and prioritize the reduction of the bias resulting from endogenous sector choice.

which is a linear fixed effect model for the mean. Thus, he suggests to compute in the first step the traditional within-estimate of the fixed effects $\hat{\alpha}_i$. In the second step, each quantile function can be estimated by a standard linear quantile regression because

$$Q_{Y_{it}-\alpha_i}(\theta|X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i) = X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta).$$

This simplifies the computation of the estimates compared to the joint estimator of Koenker (2004). However, this estimator is also inconsistent with a finite number of periods because $\hat{\alpha}_i$ suffers from the incidental parameter bias.

Kato et al. (2012) and Kato and Galvao (2016) consider quantile regression models with individual quantile-specific fixed effects:

$$Q_{Y_{it}}(\theta|X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i) = X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta) + \alpha_i(\theta). \quad (2)$$

The individual effects are only allowed to shift the distribution in (1) while in (2), they can affect the whole distribution of the outcome. While this added flexibility may certainly be useful to accomodate more complex patterns in the data, it comes at a cost – the necessity to estimate a whole function for each individual. In practice, the authors estimate the parameters by running separate quantile regressions for each quantile of interest. Obviously, even without covariates, only an approximation of $\alpha_i(\cdot)$ consisting of T different values can be estimated.

In our application this approach is not computationally feasible due to the large number of observations, the large number of quantile regressions needed to obtain the unconditional effects as explained in Section (3.2), and the number of bootstrap replications needed to estimate the variance. In addition, our Monte Carlo simulations show a relatively large incidental parameter bias for this estimator, which is probably due to the functional individual effects $\alpha_i(\cdot)$.¹¹ On the other hand, the location shift model for the individual effects in (1) seems not natural and is rejected by the data in our application. For this reason, we suggest an intermediate model with interacted fixed effects:

$$Q_{Y_{it}}(\theta|X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i) = X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta) + \alpha_i \cdot \delta(\theta). \quad (3)$$

This model treats the observed (X_{it} and S_{it}) and unobserved (α_i) regressors symmetrically by keeping them constant over the distribution but allowing them to have a different effect at each quantile. We could imagine further extensions of this model by including several individual fixed effects with different coefficients at each quantile. Ultimately, with T

¹¹Monte Carlo simulations mentioned throughout the paper are not reported but are available from the authors upon request.

different individual fixed effects, we would be back to the completely flexible model (2). For computational reasons we use the simple interacted model (3) in our application.

Assuming (3) for all $\theta \in (0, 1)$ implies

$$\begin{aligned} E[Y_{it}|X_{i1}, \dots, X_{iT}, S_{i1}, \dots, S_{iT}, \alpha_i] &= \int_0^1 (X'_{it}\beta(\theta) + S_{it} \cdot \gamma(\theta) + \alpha_i \cdot \delta(\theta)) d\theta \\ &= X'_{it} \int_0^1 \beta(\theta) d\theta + S_{it} \cdot \int_0^1 \gamma(\theta) d\theta + \alpha_i \cdot \int_0^1 \delta(\theta) d\theta \\ &\equiv X'_{it}\bar{\beta} + S_{it} \cdot \bar{\gamma} + \alpha_i\bar{\delta} \end{aligned}$$

which is a linear fixed effect model for the mean. Without loss of generality, we normalize $\bar{\delta} = 1$. Similarly to Canay (2011), we compute in the first step the traditional within-estimate of the fixed effects $\hat{\alpha}_i$. In the second step, we regress Y_{it} on X_{it} , S_{it} , and $\hat{\alpha}_i$ via traditional quantile regression. The coefficients on $\hat{\alpha}_i$ allow us to test the location shift model (1). Of course, since $\hat{\alpha}_i$ is consistent for α_i only at the \sqrt{T} rate, this estimator will also suffer from the incidental parameter bias. As explained in Section (3.3) below, we reduce the bias using the jackknife.

3.2 Unconditional Quantile Effects

The results of conditional quantile regression models with fixed effects must be interpreted carefully. It is tempting to interpret the results at low quantiles as the effect for low earners and the effect at high quantiles as the effect for high earners. This is only correct conditionally on the covariates, among which there are individual fixed effects. This means that, for high (low) values of θ , $\gamma(\theta)$ provides the effect of being employed in the public sector during the periods with high (low) wages. In other words, the inter-personal differences are captured by the fixed effects and the variation of γ over the distribution capture the differences over time.

Policy makers are certainly also interested in knowing the public sector effect on the *unconditional* wage distribution. It is clearly much easier to interpret. Moreover, issues about income inequality, for instance, are always stated in absolute terms and not conditionally on the individual unobserved ability. For this reason, we shall estimate the public sector effect on the unconditional wage distribution. We follow the procedure suggested by Chernozhukov et al. (2013) with the difference that one of the regressor (the individual fixed effect) has been previously estimated. We now precisely describe the algorithm that we use:

Algorithm 1 1. Using a standard fixed effects estimators for the mean, we obtain the estimated individual fixed effects $\hat{\alpha}_i$.

2. We estimate 100 quantile regression of Y_{it} on X_{it} , S_{it} and $\hat{\alpha}_i$ on a regular grid of 100 θ_q quantiles. For $q = 1, \dots, 100$ we obtain the estimates $\hat{\beta}(\theta_q)$, $\hat{\gamma}(\theta_q)$ and $\hat{\delta}(\theta_q)$.
3. The estimate of the counterfactual unconditional distribution in the private and public sector during the last period T take respectively the following forms:

$$\begin{aligned}\hat{F}_{Y(0)}(y) &= \frac{1}{100 \cdot n} \sum_{i=1}^n \sum_{q=1}^{100} 1 \left(X'_{iT} \hat{\beta}(\theta_q) + \hat{\alpha}_i \cdot \hat{\delta}(\theta_q) \leq y \right) \\ \hat{F}_{Y(1)}(y) &= \frac{1}{100 \cdot n} \sum_{i=1}^n \sum_{q=1}^{100} 1 \left(X'_{iT} \hat{\beta}(\theta_q) + \hat{\gamma}(\theta_q) + \hat{\alpha}_i \cdot \hat{\delta}(\theta_q) \leq y \right)\end{aligned}$$

4. We report the unconditional quantile public sector effects

$$\hat{\Delta}(\tau) = \hat{F}_{Y(1)}^{-1}(\tau) - \hat{F}_{Y(0)}^{-1}(\tau)$$

for a grid of quantiles τ .

The estimated parameter is the difference between the τ quantile of the unconditional distribution that we would observe during the last period if everybody was employed in the public sector and the τ quantile of the distribution that we would observe if everyone was employed in the private sector. These unconditional distributions are obtained by integrating the conditional distributions over the distribution of the covariates, including the estimated fixed effects, during the last period. The conditional distribution functions are approximated using 100 quantile regressions defined in Section (3.1).

3.3 Incidental Parameter Bias Correction

All the estimators discussed in Sections (3.1) and (3.2) suffer from the incidental parameter bias. Even if the number of individuals is very large, these estimator will be biased when the number of periods is finite. Arellano and Weidner (2015) characterize the bias of the estimator of model (2). They show that when the number of periods is moderate, the fixed effects estimators will underestimate the heterogeneity along the distribution by averaging the quantile coefficients around the quantile of interest. In the extreme case when $T = 2$, the estimated coefficients will be constant as a function of the quantile index θ . Thus, naively applying fixed effects quantile regression to short panels may give the impression that unobserved heterogeneity is explaining the variation along the distribution while this is only the consequence of the incidental parameter bias.

We apply the half-panel jackknife correction suggested by Dhaene and Jochmans (2015). Suppose that the number of periods T is even. Let $\hat{\gamma}(\theta)$ be the estimate based on the

whole panel. We also compute the estimates based on the first $T/2$ periods and the last $T/2$ periods, which we respectively denote by $\hat{\gamma}_1(\theta)$ and $\hat{\gamma}_2(\theta)$. The bias corrected estimator is given by

$$\begin{aligned}\hat{\gamma}_{BC}(\theta) &= \hat{\gamma}(\theta) - [0.5 \cdot (\hat{\gamma}_1(\theta) + \hat{\gamma}_2(\theta)) - \hat{\gamma}(\theta)] \\ &= 2 \cdot \hat{\gamma}(\theta) - 0.5 \cdot (\hat{\gamma}_1(\theta) + \hat{\gamma}_2(\theta)).\end{aligned}$$

The intuition is very simple: Since the incidental parameter bias is proportional to $\frac{1}{T}$, the bias of $0.5 \cdot (\hat{\gamma}_1(\theta) + \hat{\gamma}_2(\theta))$ is twice as large as the bias of $\hat{\gamma}(\theta)$. Thus, the difference between these estimates provides an estimate of the bias. We subtract this estimated bias from the original estimate.

We did numerous simulations that confirm the theoretical results and show a very significant reduction of the bias, yet at the price of seriously increasing the variance of the estimator. We could reduce the variance of the jackknife bias correction by incorporating the information about the mean coefficients. We know that the traditional fixed effect estimator, denoted by $\hat{\gamma}$, is unbiased even when T is as low as 2. At the same time, model (3) for all $\theta \in (0, 1)$ implies that $\gamma = \int_0^1 \gamma(\theta) d\theta$. Thus, our final estimator of $\gamma(\theta)$ is the recentered bias corrected estimator

$$\hat{\gamma}_{RBC}(\theta) = \hat{\gamma}_{BC}(\theta) + \hat{\gamma} - \int_0^1 \hat{\gamma}_{BC}(\theta) d\theta.$$

In simulations, the variance of this estimator is much lower than the variance of $\hat{\gamma}_{BC}(\theta)$ and only marginally larger than the variance of the uncorrected estimator $\hat{\gamma}(\theta)$.

We also use the half-panel bias correction for the estimator of the unconditional effects defined in Section (3). We correct both the first-stage quantile regression coefficients $\hat{\beta}(\theta)$, $\hat{\gamma}(\theta)$ and $\hat{\delta}(\theta)$ and the second stage counterfactual quantile functions $\hat{F}_{Y(1)}^{-1}(\tau)$ and $\hat{F}_{Y(0)}^{-1}(\tau)$.

4 Results

4.1 The Compressing Effect of the Public Sector

Raw Gap and Unconditional Quantile Effects. We first focus on the public sector wage gap over the pooled years, using the various estimation methods outlined above. Note that the unconditional quantile effects are evaluated at characteristics of the end year 2010. Figure 2 summarizes our results. We confirm some of the previous observations based on the raw wage differentials across sectors, notably the fact that the gap is larger in the first half of the distribution but much lower for top earners. Then, it is interesting to

compare the raw gap levels with the estimates of the unconditional quantile effects (UQE hereafter) on pooled data (i.e. without fixed effects). The difference is very large: the raw gap of 14% (14%, 6%) at the 10th (50th, 90th) percentile falls to 10% (6%, 2%) with UQE. This drop essentially reflects the fact that civil servants have ‘better’ observables, at least that the few covariates used for estimations on PTS data make a difference. This is notably the case of age (reflecting the higher experience of public workers) and gender (possibly lower gender discrimination in the public sector where women represent a larger share of the workforce). Note that UQE estimates still point to larger wage gaps in lower quantiles, i.e. a *compressing effect* of the public sector on the pooled wage distribution. Such a compressing effect due to positive selection on observables in the public sector is found in many studies (for instance in Mueller, 1998, Melly, 2005, Lucifora and Meurs, 2006, or Cai and Liu, 2008). That the public sector tends to reduce wage inequalities compared to private wage setting seems reasonable. Governments are indeed supposed to ensure a good pay for lower-skilled workers while the public opinion may constrain them to moderate earnings of high-skilled civil servants.

Adding Fixed Effects. Next, the comparison between UQE with and without fixed effects in Figure 2 indicates that an additional part of the raw wage gap at low quantiles is also explained by the selection of workers with better (time-invariant) unobservable characteristics in the public sector. At the 10th percentile, this represents more than a third of the remaining public sector premium. This result tends to indicate that the French public sector succeeds in attracting ‘good’ workers in the lower part of the distribution possibly because of amenities (including job protection) or non-monetary gains (notably the key element of intrinsic motivation related to the public sector ‘mission’, cf. Besley and Ghatak, 2005). It may also simply denote the relative efficiency of the national examination process in selecting talented ones in the pool of applicants, notably among young and low-wage candidates. In contrast, fixed effects regressions (FE-UQE) are doubled compare to UQE at the top of the distribution. With such a negative selection, it seems that the public sector fails to retain the most productive ones at the top.

Is the Compression Effect really gone? Overall, we find that the compressing effect of the public sector tends to disappear when fixed effects are introduced in UQE. This result is shared with several studies based on (conditional) quantile effects estimated on panel data (for instance, Bargain and Kwenda, 2012, or Nordman and Roubaud, 2014, on the informal sector wage gap). As indicated in section 3.2, however, applying fixed effects quantile regression to short panels may attribute an excessive role to unobserved heterogeneity. The incidental parameter bias is likely to explain why recent quantile regression estimates always point to a flattening of the public sector compression effect when fixed

effects are taken into account, for instance in Siminski (2013), Hospido and Moral-Benito (2016) or Bargain and Melly (2008). According to these analyses, the compression effect can be partly or fully explained by self-selection of better-skilled workers among civil servant at the bottom of the wage distribution. The best illustration of our point is provided by Campos and Centeno (2011, table 5.5), who show that this flattening – and the quasi-disappearance of the public wage gap – occurs systematically across a wide range of European countries.

Jackknife Corrections. One of our main contributions is therefore to suggest a correction of the incidental parameter bias. Figure 2 shows that the jackknife correction almost restores part of the compression effect found with UQE on pooled data (i.e. without fixed effects). For instance, the difference between the 10th and 90th percentile is doubled with the correction (it is 3.5 percentage points without correction, 6 points with correction, and 8 points with UQE on pooled data). This show that the hypothesis of a compressed wage profile of the public sector due to non-competitive wage settlements cannot be completely overruled.

Nonetheless, a part of the compression effect can still be partly explained by a selection of workers on unobserved characteristics. This result concerns lower quantiles and consolidates our previous discussion on the likely positive selection of public workers in lower quantiles. Since the entry in the public sector is conditioned to passing competitive exams, it may be that public workers at the lower end of the distribution are also those who have succeeded and are likely characterized by higher unobserved skills than private sector workers with similar observables. Besides, jackknife-corrected results do no longer show negative selection into the high-earning public sector. This is a reassuring thought for those concerned by the efficiency of top management in the public sector and the lack of performance-based bonuses in the administration.

4.2 Time Trends in the Public Sector Wage Gap

Few studies look at the evolution of the public wage differential over time, especially when unobserved skills are taken into account. Yet, as noticed by Disney and Gosling (2007), this is the time-varying public wage differential that has the most relevant policy implications.¹² Our long panel can fruitfully be used for that purpose while controlling for fixed effects and incidental parameter correction. In particular, we study how the unobserved relative quality of the public sector workforce varies over time (overall and at

¹²Hospido and Moral-Benito (2016) for instance compares the wage compression between 2008 and 2012 and find that wage compression appears in 2012, after the crisis has affected private wages and led to public wage cuts, representing a threat to the recruitment of high-skilled public workers.

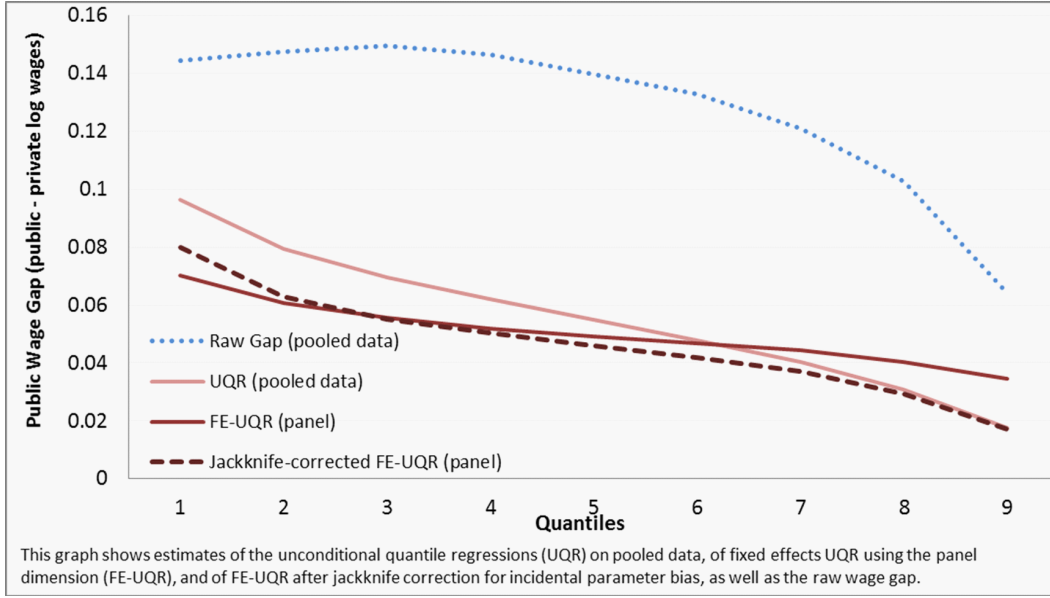


Figure 2: Public Wage Gap Estimations: Comparing Methods

different quantiles), and how it relates to changes in the public wage gap itself and, more generally, to the selectivity of the public recruitment process.

Evolution of the Average Gap. Figure 3 first shows the evolution of the average wage difference across sectors (raw gap) and of the estimated public sector wage gaps when controlling for basic covariates (OLS) and, additionally, for fixed effects (FE). The time change in raw gaps have been commented before at different quantiles (see Figure 1). Overall, we find again an increase in the public advantage until the mid-1990s, a plateau and a decline in the early 2000s. Comparing the raw gap to OLS, we derive similar interpretations as in Figure 2. That is, a large part of the observed differential is explained by simple covariates as gender, age and occupations (the gap oscillates between 1.5% and 7.5% over time, surrounding the median UQE estimate of 5% obtained on pooled years). As seen before, the pay gap further decrease when fixed effects are included.¹³ Arguably, the latter may capture missing observables in the PTS data (like education) or more generally ‘better’ unobservables among public sector workers relative to their private sector counterparts. As we shall see in robustness checks using matched PTS-EDP data, the former explanation plays a minor role while the latter, i.e. a positive selection into the civil sector, is the bulk of the story.

Note that given the size of the PTS, estimations are extremely precise (standard errors

¹³Note that the corrected FE-UQR yields a gap of 4.5% at the median, which is higher than the FE estimate over all the years. This difference remains to be explained.

are reported in Table A.2 in the Appendix). Notice also that the difference between OLS and FE estimates of the public wage gap decreases steadily until 2008. As we shall see in robustness checks using matched PTS-EDP data, this is not due to a changing role of covariates not present in the PTS, for instance a decline in the returns to education for civil servants. This pattern rather means that unobserved skills in the public sector become less and less favorable relatively to the private sector (an exception is the bounce in 2009-10). That positive selection in the public sector tends to fade away over time must pertain to the pay gap itself and to the selectivity in public jobs, which we will investigate a bit later.

Interpretation of the Time Trend: Policy and Business Cycles. For now, let us provide interpretations of the time trends in the wage gap. We focus on FE estimates as our best measure of the difference in pay, all things equal, between the two sectors. We actually see that the public sector wage gap basically oscillates around zero over the long run. This is not really evidence of competitive markets: the period is long and wage equalization never really occurs (even if average wage gaps are never large). Interestingly, the pattern we observe is explained by a combination of policy and business cycles. Policy cycles can be decomposed in presidential terms: Mitterrand from 1988 to 1995, Chirac I from 1995 to 2002, Chirac II from 2002 to 2007 (beginning of 5 rather than 7-year terms), Sarkozy from 2007 to 2012. In a traditionally two-party system, the party of the president in power is a good indicator of the general policy governing the public sector in France.¹⁴ Looking at Figure 3, we see a sharp public sector penalty characterizing the late 1980s, which reflects the turn to rigor after 1983 in France, accompanied by wage restraint (de-indexation of wages on prices) and the prioritizing of public-debt reduction (implying a decrease in public wage costs). Better conditions under Mitterrand's (second) term have allowed the socialists to conduct a compensating policy. The period 1988-1995 is indeed marked by measures precisely allowing for "catch-up" pay increases in the public sector. Several such measures (*plans de réformes catégorielles*) have started in 1989 (Jospin), 1990 (Durafour) and 1993 (Lang). The first half of the nineties is also a period of slack in the labor market with moderate increases in private sector wages. These combined trends

¹⁴The Fifth Republic regime in France has given a much prominent role to the president, legitimated by the fact that he is elected by direct popular vote. Moreover, a president usually benefits from a majority at the parliament (*Assemblée Nationale*), as the latter is renewed two months after the president's election. Exceptions include periods of so-called "cohabitation", during which the president has lost this majority and must nominate a prime minister from the opposition. The duality of the executive power has been experienced by Mitterrand (socialist) and Balladur (conservative) under the "velvet cohabitation" of 1993-1995. Inversely, Chirac had to rule with the socialist prime minister Jospin in 1997-2002. These exceptions do not invalidate our interpretations over the long period.

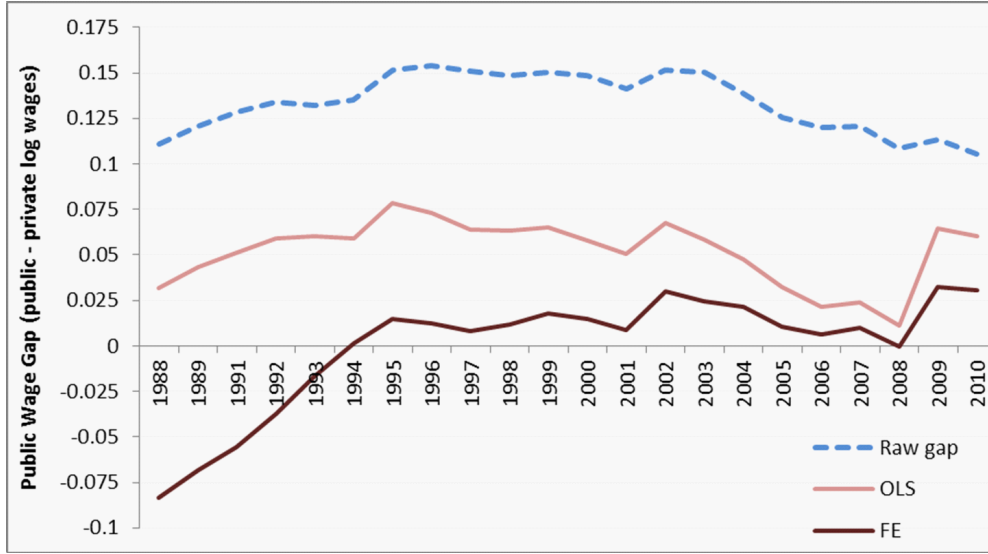


Figure 3: Raw Gap, OLS and FE Estimates over Years 1988-2010

fully explain our results for 1988-1995.

The relative wage progression in the public sector has stopped in the following period, with a relative stagnation of the public wage gap. This is possibly the result of two opposite forces: better economic conditions in the late 1990s (improving private wages) and the continuation of the *réformes catégorielles* until 2000. Business cycles play a salient role especially over short periods. The recession of 1993 has no visible impact but the economic slowdown of 2001-02 seems to affect particularly the private sector and to explain a temporary bounce in the public premium at this point in time. With his reelection in 2002, supported by a large parliamentary majority, Chirac implements a conservative program including tax cuts and a relative control of public wages. Yet this is more the economic upturn of the period that drives the gradual decline in the public sector premium. The rebound of our end period clearly pertains to the recent crisis, marked by a sudden drop in private wages in 2009.

Evolution of the Wage Gap for different Quantiles. Figure 4 reports estimates of the public wage gap over time using UQE and FE-UQE (jackknife corrected FE-UQE estimations will also be added in a revised version of this paper). Overall, time trends are very similar across unconditional quantiles, especially the sharp decline after 2002 and the bounce after 2009. With UQE, differences are reminiscent of our analysis of Figure 1. The trends in the raw wage gap actually showed similar patterns at different quantiles at most periods but a decline among low quantiles – rather than a plateau for median and higher quantiles – for the intermediary period 1995-2002. Similarly here, estimates

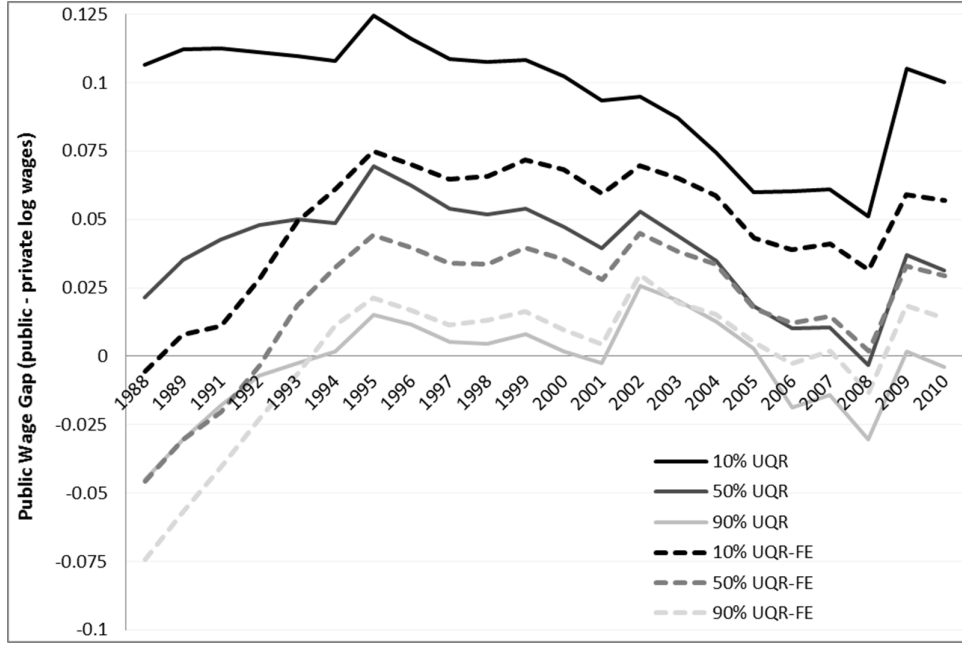


Figure 4: Unconditional Quantile Estimates Over Years 1988-2010

for the 10th percentile show a relatively stable gap in the first period and a decline in the intermediary period. The 50th and 90th percentiles are more in line with the time pattern at the mean.

If we turn to FE-UQE estimates, we observe much less differences across quantiles. We see the rather regular ordering of wage gaps across quantiles, as discussed before, and notice that this pattern is relatively stable over time. Precisely, with the exception of the catch-up period before 1995, the public sector premium oscillates around 5 – 6% for the 10th quantile, around 2 – 3% for the median and around 1% for the 90th. Rather parallel curves for 10th and 90th percentiles mean that the compressing effect of the public sector is a relatively constant feature of the French wage setting. Next, we see that time changes actually follow very closely the pattern already discussed for the mean gap estimated by FE: a rise until 1995, then a relative stagnation until 2002 and, finally, a sharp decline until the recent crisis. The bounce in 2009 is similar across quantiles but the impact of the short-lived recession at the turn of 2001-2002 has affected private sector workers especially in the upper part of the distribution, which is consistent with the nature of the e-bubble crisis.

Evolution of the Positive Selection into the Public Sector. If we now consider the difference between UQE and FE-UQE, we observe a similar trend for the low quantiles as what we discussed before for the mean (i.e. for the difference between OLS and FE):

a gradual decline in the unobserved premium attached to the public sector. The same is true at the median, yet the potential effect of the public sector recruitment process is less visible, i.e. we start with smaller premia and they completely disappear in the early 2000s and until the end of the period under study. For high quantiles, the premium oscillates around zero throughout the period and even turns into a penalty in the last subperiod. As discussed above, these results may change with jackknife correction (to come). In particular, as seen in estimations over all years (Figure 2), the negative selection at the top of the distribution disappears when the incidental parameter bias is removed.

Yet, this correction may not change the overall picture discussed with FE estimation: the public sector unobserved premium tends to fade away. It is tempting to relate it to the attractiveness of the public sector. We do so in the Appendix A.1 (Figure A.1) where we compare the "unobserved public premium" (gap between UQE and FE-UQE) at 50% and the selectivity rate for intermediary positions in the public sector (category B). This comparison is limited, however, since the premium is the result of estimations based on the whole public sector in France, while selectivity rates are available only for entry in the tenured national civil service. In Appendix A.2 (Figure A.2), we attempt to explain the trend in selectivity using two usual determinants: the public wage gap itself and unemployment. We show that selectivity in tenured civil jobs depends mainly on unemployed for low and median wage workers while it rather depends on the public wage gap for high wage workers.

In Figure 5, we represent the trend in the unobserved public premium for the three points of the distribution. Critically, the overall decreasing pattern might be explained, at least partly and after 1995, by the long-term decline in the public wage gap itself (as measured by FE-UQE). Other studies have actually pointed to the quality of the public workforce being endogenous to the perceived wage differential (Nickell and Quintini, 2002, Disney and Gosling, 2008). In our case, the relationship is not clear for high wages. However, for low and median wages, the fall of the unobserved public premium lines up quite well with the almost continuous decline in the public wage gap after 1995. The last puzzle is what happened before 1995. This subperiod is an exception in the sense that the catch-up pay policies have not managed to improve the quality of the public sector. The reason is that increased financial incentives to join the civil sector have indeed boosted the number of applicants, but many of them did not enter the public sector through the selective national examination procedure, rather as non-tenured public employees.¹⁵ The same reasoning also contributes to explain the declining unobserved public premium in more

¹⁵The "protocole Durafour" plan launched in 1990 not only deeply reformed the public wages structure but also tried to compensate the increasing number of retirements by more vacancies and in particular by accelerating entries in the public sector without competitive exams.

recent years.¹⁶

4.3 Robustness Checks and Additional Results

Movers. The identification of the public wage gap in panel estimations requires that there is enough transitions across sectors. In the data section, we have provided general statistics that convey that it is the case. Using administrative data has the advantage of dealing with a very large number of observations, so that the wage gap can be precisely estimated, even if the movers represent only a fraction of the workforce (0.73% of the workforce per year, on average, and 5% of all the workers over the course of the panel, see Table 1). We now complete this information with additional statistics. Figure 6 presents overall transitions across sectors, calculated as the proportion of movers (over a year's workforce) by quintile (defined as the wage quintile of destination) averaged over the years 1988-2010. It conveys that moves are not concentrated at particular points of the wage distribution. We only notice more frequent transitions at the second quintile, and more frequent moves from private to public sector than in the other direction (overall, but especially in the third quintile).

In addition, Figure 6 shows the time trend in the number of transitions/movers per year and per quintile. Observations go as follows. *First*, there is a substantial number of moves per year and at different quintiles for a reasonable identification of our most detailed estimates (FE-UQE with time interactions). There is roughly around 500-2500 observed moves per year and quintile. The exception is a more frequent transition towards the public sector (upper graph) among quintiles 2 and 3, as remarked above, which seems to characterize the whole period. We also notice an acceleration of entries in the public sector at the lowest quintile during the recent crisis, which coincides with specific policy measures (creation of temporary public jobs for low-skilled workers). *Second*, if we specifically look at moves towards the private sector (lower graph), there seems to be a monotonic pattern with wage levels: the higher the quintile, the lower the frequency of a move to the private sector. This is likely due to the fact that civil servants with a higher pay are those with more experience in the public sector (hence less probability of changing) and with more chances of being tenured (which also considerably reduce the chance of transiting, given the cost of passing entry examination for tenured public jobs). *Third*, time trends are

¹⁶For instance in 2001, a new law has extended competitive hires for tenured position to workers coming from the private sector (*troisième concours*, see Bounakhla, 2015). While quotas reserved for this new scheme have attracted new applicants (and contributed to the increase in selectivity seen in Appendix Figures A.1 and A.2), they may have also led to the absorption of employees selected on their professional experience and not on their education level or ability to pass selective examinations as in the traditional system, possibly contributing to degrade the positive selection in the civil sector.

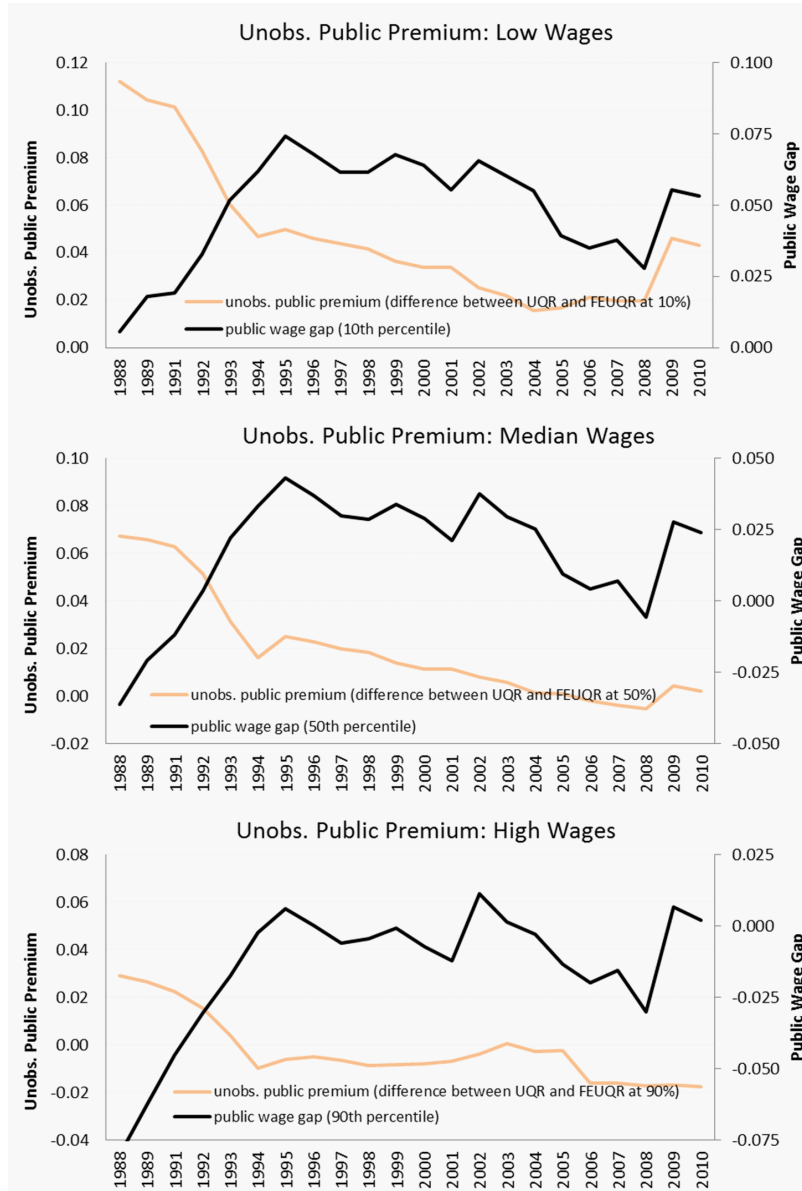


Figure 5: The Declining Unobserved Public Premium

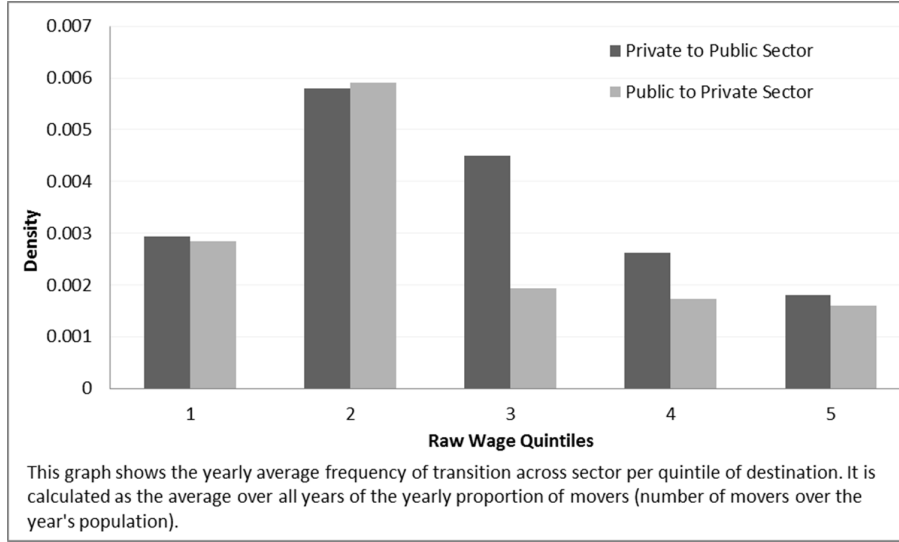


Figure 6: Proportion of Movers per Quintile

relatively comparable for both directions of the moves, with a slowdown in the early 1990s and a long-term increase in the number of moves after 1995. This similarity means that time trends in moves are not directly interpretable on the basis of changes in public sector policies – which would primarily affect transitions towards the public sector. Note also that the main time variations are jumps corresponding to economic downturns: 1991, 2002 and 2008-2010. Interestingly, transitions during these crisis go both ways (even if more marked for the private-to-public transitions, as expected).

Adding Covariates. As discussed, OLS and UQE estimations using the PTS were based on a limited set of explanatory variables (age, gender and position). In Figure 8, we report time trends of the mean wage gap with OLS and FE estimations. We compare baseline estimates using the PTS with estimates obtained with the matched PTS-EDP data. The EDP contained usual variables for wage equations, as described in the data section, and notably the potential experience, education and marital status. As expected, OLS estimates decrease when using these variables, reflecting better endowments among public sector workers. Also expected is the fact that FE estimates decrease much less, since much of the information is captured by fixed effects (education does not change much over time). Importantly, our conclusions are not affected. Indeed, the time trend in the public sector wage gap is very similar to the one previously described and interpreted on the basis of PTS data. The fading away of the positive selection into the public sector also appears very clearly when comparing OLS and FE estimates with the PTS-EDP matched data. As for our baseline estimations, the coefficients estimated on PTS-EDP

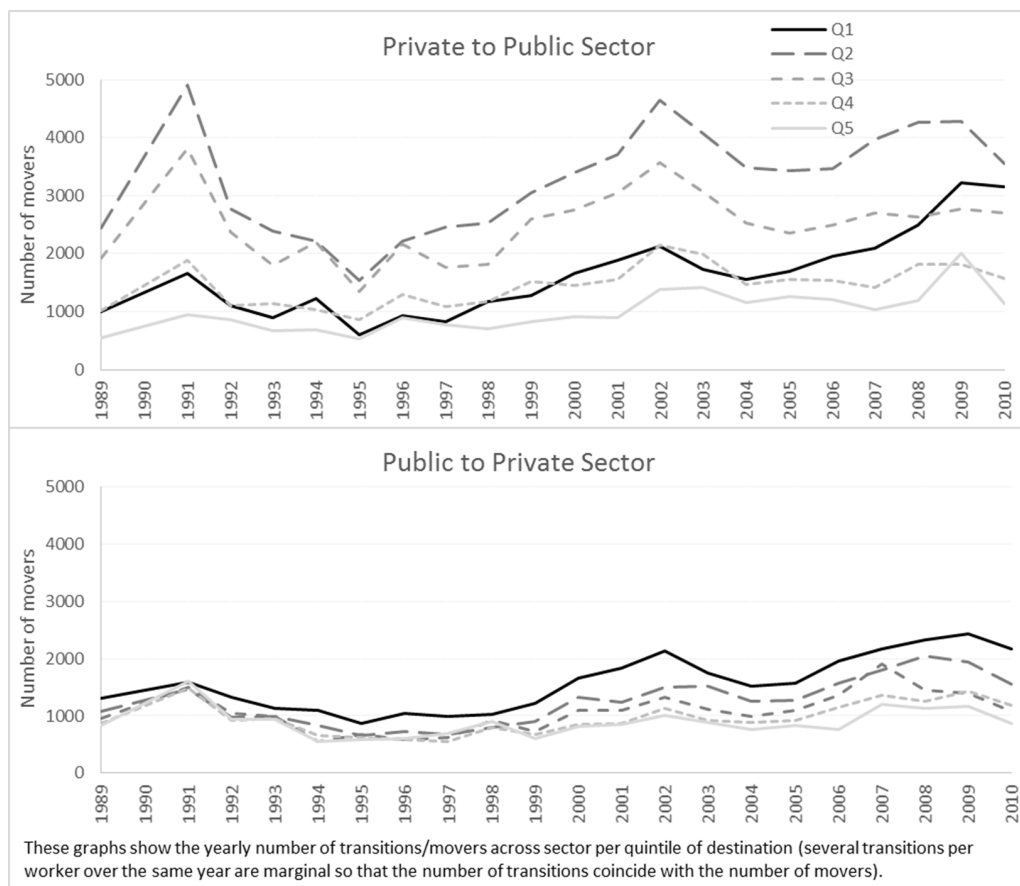


Figure 7: Evolution of the Moves across Sector by Quintile

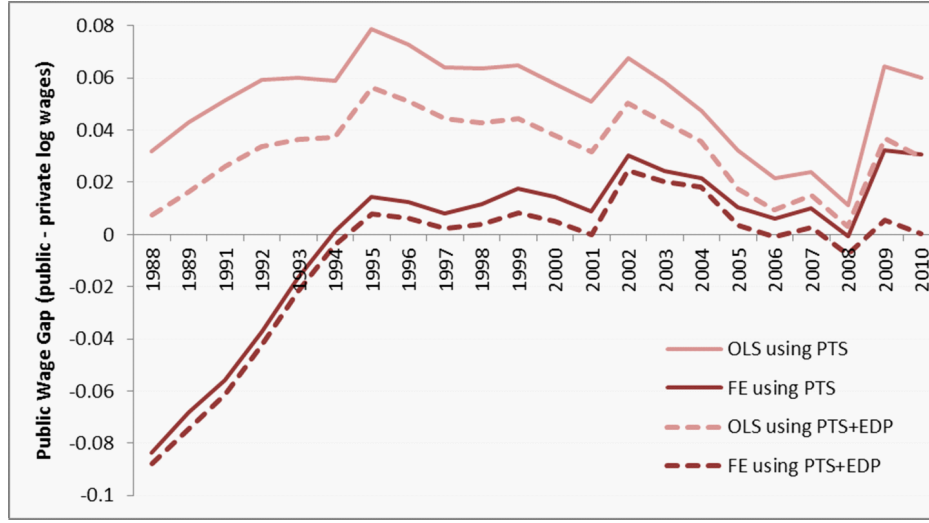


Figure 8: OLS and FE Estimates: Adding EDP Covariates

are very precise, as can be seen in Table A.2 in the Appendix.

5 Conclusion

This paper contributes to the literature in three ways. First, we estimate the public wage gap over 22 years of a large administrative panel, at the mean and along the unconditional wage distribution. As previously shown, monetary returns are higher in the public sector, but the public wage gap disappears progressively since 1995. There is still a compressing effect: the pay gap is sufficiently different between bottom and top quantiles to narrow a little the wage distribution. Second, the long duration of the panel and innovative jackknife methods allow correcting for the incidental parameter bias. The latter tends to underestimate the role of unobserved heterogeneity (and the compressing effect), here and possibly in most studies applying fixed effects quantile regressions. Third, we investigate the time trends over 1988-2010. We indicate that the decline in the public wage gap may have contributed to the gradual decrease, after 1995, in the quality of the public workforce, as measured by the unobserved public premium. This critical aspect calls for better wage-quality mix measures. Another factor is the diversification of recruitment methods, which tends to decrease the overall selectivity of the public sector.

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A Appendix

A.1 Selectivity and the Unobserved Public Premium

In Figure A.1, we compare selectivity and the public workforce quality measure based on the unobserved public premium. As said in the main text, this comparison is limited by the fact that the premium is the result of estimations based on the whole public sector in France, while selectivity rates are available only for entry in the tenured national civil service (around 30% of the public workforce in recent years). Nonetheless, we observe an overall declining trend in selectivity starting in 1995, with the exception of years 2004-2006. The bounce of 2009-2010 also lines up well with the trend in unobserved public premium. The major difference is found under Mitterrand's term. A possible explanation is that pay rise at that time (*réformes catégorielles*) have attracted new applicants – hence risen selectivity – but many of them have actually joined the public sector not through the selective process but through alternative recruitment processes, as explained in the main text. This has possibly led to a sharp decline in unobserved skills in the civil sector. Again, further robustness checks will be needed to shed light on these questions, notably a clean comparison of selectivity (for tenured civil jobs) with an unobserved public premium estimated on a subsample including only tenured civil employees.

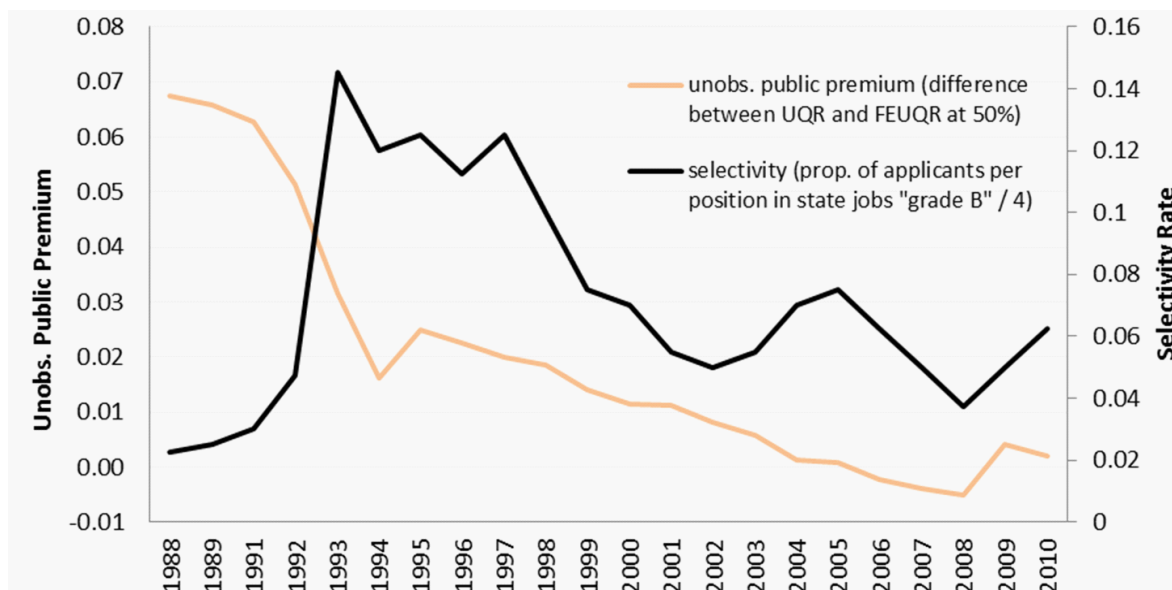


Figure A.1: Selectivity versus Unobs. Public Premium at Median Wage Levels

A.2 Selectivity, Unemployment and the Public Wage Gap

We can relate selectivity to some of its determinants: the public wage gap itself and the unemployment rate (see also Pouget, 2003, on an older period). Both affect the demand for public positions (the numerator of the attractivity rate) while vacancies (denominator) vary with other factors that are more difficult to control (e.g. conservative parties tend to reduce the public sector). We assume that the public wage gap measured using the whole public sector is a good proxy, even to explain selectivity in the tenured civil service.¹⁷ We shall nonetheless refine by looking at selectivity for civil position at the three grades of the administration (categories A, B and C). We associate to each of them the relevant unemployment rates (i.e. for high, median and low skills) and use the public wage gap at 90, 50 and 10%, respectively. Results are shown in Figure A.2. At the three wage levels, the attractiveness of tenured jobs in the civil sector fluctuates, with peaks around 1993-1998 and 2004-2007. This seems relatively well correlated with the variation in unemployment, especially for low-skilled workers. This is consistent with the fact that the latter are subject to higher risk of unemployment (as conveyed by the difference in unemployment rates across the three groups), and hence have a higher propension to try to obtain a tenured position for job security. The link between selectivity and the public sector wage gap is less marked. It is nonetheless quite visible for the high-skilled, which may be explained by a greater ability to transit across sectors. The correlation is negative only between 2002 and 2007: Chirac II's government did not only attempted to moderate public wage progression but also to reduce the size of the tenured civil sector while substituting with more flexible public sector jobs.¹⁸

To illustrate these results, we have run a regression of the selectivity rate overall or for each wage groups. Results are reports in Table A.1. We naturally refrain from any causal interpretation, since the vacancies for tenured positions (the denominator of the selectivity rate) depend on public sector policies that also affect the public wage gap. Nonetheless, we see that unemployment, the public wage gap and Chirac II's term "explain" 67% of the variance in the selectivity rate overall (unreported: unemployment and wage gaps alone explain 62%). Results for the low wages confirm that unemployment is the significant determinant (and explain around a quarter). For median wages as well, unemployment seems the main driver (and explains as much as 78% of the time variance in selectivity).

¹⁷There is no particular reason for the State to discriminate among its employees, tenured or not (salary scales are official and concern all the public sector workers within each administration).

¹⁸The workforce of State services, where tenured jobs are more represented, has decrease by 1% over these six years. The proportion of non-tenured in the whole public sector has increased from 14% to 16.5%. The small decline in public gap over 1999-2001 may have similar causes: the share of temporary contracts in the public (private) sector has increased (decrease).

For high wages, the public sector wage gaps is the significant factor. The model with lagged wage gap even improves the fit (up to 55% of the variance is explained).

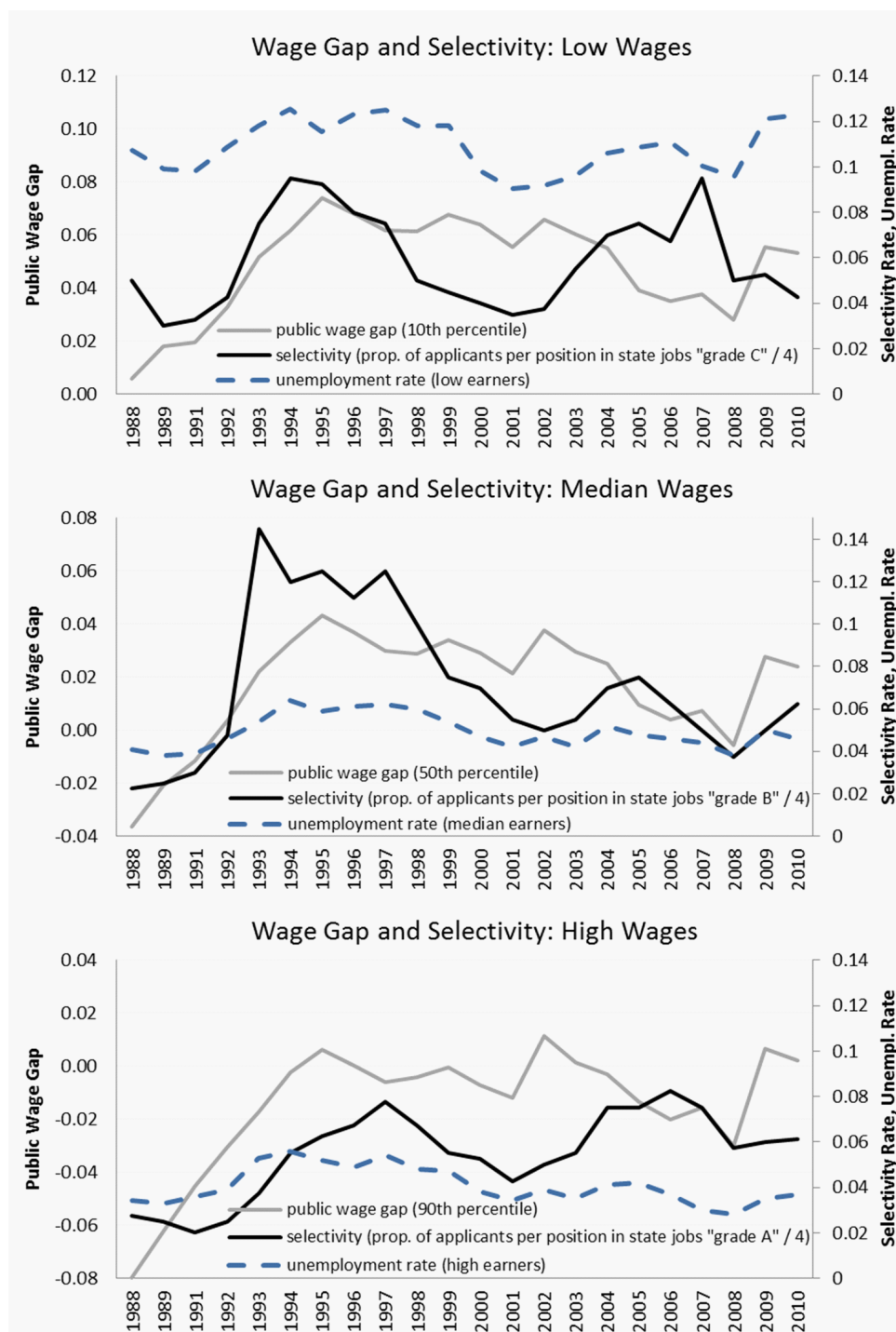


Figure A.2: Wage Gap, Unemployment and State Sector Selectivity

Table A.1: Correlates of Public Sector Attractiveness

	All groups	Low wages	Low wages	Median Wages	Median Wages	High wages	High wages	High wages
Public Wage Gap (a)	82.54** (36.82)	65.99 (92.92)	46.27 (94.55)	64.30 (100.9)	74.05 (108.6)	207.8*** (63.62)	178.7** (65.80)	
Public Wage Gap, lagged (a)								207.2*** (56.07)
Unemployment Rate (b)	5.679*** (1.198)	3.232** (1.519)	3.830** (1.619)	14.22*** (2.585)	13.96*** (2.794)	-0.352 (1.777)	0.294 (1.802)	
Chirac II (c)	2.969 (2.830)		4.265 (4.062)		-1.103 (3.773)		4.380 (3.216)	2.732 (2.893)
Constant	-25.86** (10.69)	-15.06 (15.61)	-21.60 (16.76)	-42.46*** (11.64)	-41.10*** (12.81)	26.39*** (7.785)	22.34** (8.177)	24.12*** (6.917)
Observations	22	22	22	22	22	22	22	21
R-squared	0.673	0.270	0.312	0.789	0.790	0.393	0.450	0.552

Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

(a) estimated by FE or FE-UQR, using 10, 50 and 90 percentiles for low, median and high wage groups respectively

(b) using rates corresponding to each wage group

(c) dummy for years 2002-2006

A.3 Public Wage Gap Estimates: Alternative Data

Table A.2: OLS and FE Estimates on PTS versus PTS + EDP

VARIABLES	PTS		PTS+EDP		VARIABLES	PTS		PTS+EDP	
	(1) OLS	(2) FE	(1) OLS	(2) FE		(1) OLS	(2) FE	(1) OLS	(2) FE
Public*yr1988	0.0318*** (0.000676)	-0.0835*** (0.000797)	0.00736*** (0.00186)	-0.0877*** (0.00217)	Public*yr2006	0.0216*** (0.000631)	0.00609*** (0.000740)	0.00949*** (0.00127)	-0.000798 (0.00186)
Public*yr1989	0.0432*** (0.000671)	-0.0681*** (0.000781)	0.0162*** (0.00184)	-0.0748*** (0.00211)	Public*yr2007	0.0241*** (0.000639)	0.00993*** (0.000749)	0.0152*** (0.00131)	0.00275 (0.00188)
Public*yr1991	0.0512*** (0.000666)	-0.0557*** (0.000759)	0.0260*** (0.00183)	-0.0613*** (0.00206)	Public*yr2008	0.0111*** (0.000637)	-0.000482 (0.000752)	0.00298*** (0.00131)	-0.00760*** (0.00189)
Public*yr1992	0.0593*** (0.000690)	-0.0373*** (0.000793)	0.0337*** (0.00189)	-0.0427*** (0.00214)	Public*yr2009	0.0644*** (0.000616)	0.0324*** (0.000757)	0.0367*** (0.00142)	0.00542*** (0.00193)
Public*yr1993	0.0601*** (0.000722)	-0.0167*** (0.000790)	0.0364*** (0.00196)	-0.0215*** (0.00215)	Public*yr2010	0.0602*** (0.000620)	0.0308*** (0.000764)	0.0299*** (0.00143)	0.000138 (0.00194)
Public*yr1994	0.0588*** (0.000732)	0.00125 (0.000784)	0.0373*** (0.00199)	-0.00395* (0.00211)	Vocational training			0.0415*** (0.00107)	
Public*yr1995	0.0787*** (0.000693)	0.0145*** (0.000766)	0.0562*** (0.00190)	0.00771*** (0.00206)	High School			0.0858*** (0.00129)	
Public*yr1996	0.0728*** (0.000702)	0.0123*** (0.000762)	0.0510*** (0.00192)	0.00627*** (0.00204)	University first degree			0.123*** (0.00140)	
Public*yr1997	0.0640*** (0.000698)	0.00794*** (0.000757)	0.0442*** (0.00191)	0.00224 (0.00204)	University upper degrees			0.104*** (0.00170)	
Public*yr1998	0.0635*** (0.000681)	0.0115*** (0.000748)	0.0428*** (0.00186)	0.00375* (0.00200)	Couple			0.0162*** (0.00146)	0.0179*** (0.00126)
Public*yr1999	0.0649*** (0.000682)	0.0178*** (0.000748)	0.0443*** (0.00187)	0.00846*** (0.00201)	Couple with children			0.00413*** (0.00153)	-0.00707*** (0.00130)
Public*yr2000	0.0578*** (0.000677)	0.0145*** (0.000749)	0.0380*** (0.00187)	0.00512** (0.00200)	Age	0.0263*** (8.65e-05)	0.0262*** (0.00200)	0.0262*** (0.000222)	0.0276*** (0.00260)
Public*yr2001	0.0508*** (0.000681)	0.00898*** (0.000751)	0.0317*** (0.00187)	-0.000168 (0.00199)	Female	-0.0635*** (0.000354)	0.0141 (0.0331)	-0.0788*** (0.000829)	-0.262*** (0.000803)
Public*yr2002	0.0675*** (0.000686)	0.0302*** (0.000752)	0.0505*** (0.00139)	0.0246*** (0.00188)	Intermediary Profess	0.216*** (0.000355)	0.0757*** (0.000281)	0.182*** (0.000886)	0.0700*** (0.000675)
Public*yr2003	0.0583*** (0.000705)	0.0244*** (0.000756)	0.0429*** (0.00140)	0.0203*** (0.00189)	Executive	0.375*** (0.000510)	0.179*** (0.000485)	0.333*** (0.00134)	0.168*** (0.00122)
Public*yr2004	0.0475*** (0.000692)	0.0216*** (0.000754)	0.0356*** (0.00138)	0.0181*** (0.00188)	Constant	1.367*** (0.00157)	1.434*** (0.0536)	1.668*** (0.00420)	1.827*** (0.0663)
Public*yr2005	0.0324*** (0.000676)	0.0104*** (0.000752)	0.0175*** (0.00135)	0.00350* (0.00188)	R-squared	0.384	0.284	0.392	0.252
					# observations	11,521,606		1,938,461	
					# workers	1,055,727		202,553	

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Notes: Public*yr1988 is the interaction of the public sector dummy with the 1988 time fixed effect. Individuals with less than 4 observations are dropped, which explains the slightly lower sample size (11.5 million individuals rather than 12.18 in the base selection).