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Wage Inequalities and Firm-Specific Compensation Policies in France

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ABSTRACT. – This paper examines the evolution of the wage structure in France after 1984. Our data come from two matched employer-employee wage surveys performed in 1986 and 1992. So, we have two cross-sections of establishments and individuals. A subsample of establishments present in both surveys allows us to analyse time-variations.

We find that the wage inequality increased between 1986 and 1992, which seems to be, in large part, explained by the evolution of employer-specific compensation policies. We analyse the role of employer characteristics in this evolution. We also show that between-plant specialization dramatically increased during the period in all dimensions. Finally, we observe that the evolutions of employer-specific wage policies are correlated with changes of the workforce in terms of experience and seniority.

Inégalités de salaire et politiques de rémunération des entreprises en France

RÉSUMÉ. – On observe entre 1986 et 1992 un léger accroissement des inégalités de salaire en France, lié à la divergence des politiques de rémunération des employeurs, alors que les caractéristiques des salariés – particulièrement l'ancienneté – voient leur poids diminuer. On explique les politiques de rémunération par les caractéristiques des employeurs ; leur évolution s'accompagne d'une modification des caractéristiques des salariés (âge moyen, ancienneté moyenne). Enfin, on observe une homogénéisation des qualifications au sein des entreprises.

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1 Introduction

This paper examines the role of employer-specific compensation policies in explaining the increasing wage inequalities observed in France after 1984. To accomplish this task, we use two matched employer-employee surveys, the Enquête sur la Structure des Salaires (Wage Structure Survey), realized in 1986 and 1992.

It is by now well documented that firm-specific compensation policies played a major role in the increase in wage inequalities observed in the US (LEVY-MURNANE [1992], pp. 1368-1369). According to GROSHEN [1989], plant-specific wage differentials explain 27 percent of the residual earnings variation in a cross-section ; DAVIS and HALTIWANGER [1991] find that 40 to 46 percent is due to plant-specific wage differentials. Furthermore, the latter study examines the growth in wage differentials between 1975 and 1986 and finds that half of it is due to the growth in differentials between plants of the manufacturing sector. Understanding the sources of such interfirm differentials is considered by LEVY and MURNANE [1992] as one of the most important directions of future research.

In order to analyze the role of employer-specific compensation policies in the wage formation process and eventually in the drift of wage inequality, we must be able to assess the respective impacts of worker characteristics and employer characteristics on wage levels. This may be achieved empirically by the estimation of a wage equation including both workers and employers variables. Precisely, our analysis relies on a wage equation of the following pattern :

$$wage_{it} = \gamma_{jt} + \alpha_i + x_{it}\beta + \epsilon_{it}$$

In this equation, the wage of a worker i at an employer j at time t depends on a vector of observed characteristics of the worker x_{it}, some unobserved characteristics of the worker summarized by the fixed-effect α_i and a time-varying employer-specific fixed-effect γ_{jt}. ϵ_{it} is a disturbance term.

In this equation, γ_{jt} represents a base level for wages at employer j at time t and thus, denotes whether workers at this employer are, at time t, more or less paid than at other employers.

To estimate the previous wage equation, the data requirements are stringent. To identify employer compensation policies from individual wage determinants, multiple types of data are necessary (see ABOWD, KRAMARZ and MARGOLIS [1994]). First, we need information on the workers to control for their observable characteristics (x_{it}). Second, we need longitudinal information on these workers to control for their unobserved characteristics (α_i). Third, we must be able to know the workers' employing firm ($j(i,t)$). Then, we must follow these employers on the period.

By surveying establishments of the manufacturing and construction industries and a representative sample of their workers, the Wage Structure Survey accomplishes the first and the third tasks both in 1986 and 1992 for a broad sample of establishments (respectively 10,000 and 4,000 units with respectively 320,000 and 43,000 workers). But we cannot match the two samples of workers since there is no common identification number. So we have two samples of workers, one at each date, and we estimate separately two wage equations (coefficient β is allowed to vary between the two dates). Besides, since we have cross-sections, we cannot disentangle the employer fixed-effect γ_{jt} from its workers' fixed-effects α_i. In particular, the estimation of γ_{jt} is not possible. We can only estimate a global fixed-effect for each employer : after estimating β_t by $\hat{\beta}_t$, the fixed-effect is calculated for employer j as the mean value of the difference $wage_{it} - x_{it}\hat{\beta}_t$, for all workers i at employer j. In fact, the fixed-effect for employer j estimates the sum of γ_{jt} and of the mean value of the individual fixed-effects α_i of the workers of that employer.

Hereafter, **these global fixed-effects will constitute our measures of employer-specific compensation policies**.

From our samples of 1986 and 1992, we can estimate the fixed-effects of the establishments with at least two workers sampled. At each date, we try to explain these two values by establishment-level variables. Moreover, we analyze the variation of this fixed-effect from a subsample of 618 establishments present at both surveys. Eventually, to gain even more insight into the wage determination process, we also perform within-establishment regressions at both dates for 132 establishments.

Our findings can be summarized as follows. Inequality has increased in France between 1986 and 1992 in the manufacturing sector which confirms BAYET-CASES [1994] and KRAMARZ-MARGOLIS [1994] ' findings from a different data source. Employer-specific compensation policies are responsible for a large part of this observed increase. Furthermore, we show that workers' specialization between firms has increased drastically during the period (*i.e.* workers with identical observed characteristics are employed more and more at the same firms). We also show that seniority is less important in 1992 than in 1986, the relative weight of experience among wage determinants increased, when compared to tenure. Finally, between-establishment growth in wage differentials is mostly explained by the growth in employment, the signature of a firm-level agreement, the increases in the shares of the most-skilled employees (in particular, engineers, professionals and managers), and production technology (work in shifts).

Among the stylized facts that we mentioned, one of the most striking is the increase of specialization of workers within establishments. This trend, joined with a strong relation between the distributions of skills in the workforce and employer fixed-effects, may explain the divergence of employer-specific compensation policies. However, since our study remains at a descriptive level, we do not try to explain why we observe these facts. We can only suggest some explanations. For instance, the manpower specialization may be due to technological reasons, with employers concentrating on their most profitable activities or increasing in sub-contracting for non-core activities. Or, concerning the increase of wage inequality, the small increase in the minimum wage during the period, relatively to the mean wage, may have helped employers pay low wages to unskilled workers, thus inducing a widening of the bottom of the wage distribution, as compared to previous periods. However, our analysis will remain purely descriptive.

In the second section, we present our datasets. The next section examines patterns of evolution in wage inequality and specialization. The fourth section contains our estimates of the wage equations. The fifth section examines the employer-specific differences and their evolution. Finally, the sixth section concentrates on the respective roles of experience and seniority in the employers' compensation policies.

2 The Data

The Wage Structure Surveys (Enquête sur la Structure des Salaires, ESS hereafter) performed by the French National Statistical Institute (INSEE) in 1986 and 1992 were initiated in 1966 by the European Statistical Office (ESO). However, after the 1966, 1972 and 1978 surveys, the ESS program has been abandoned by the ESO. INSEE decided to resume this survey because of the amount of information collected at each time. The ESS collects individual wages within a sample of establishments of the manufacturing industries and of firms in the construction and in some service sectors. The agriculture, transportation, telecommunication and services supplied to individuals sectors are excluded from the scope of the ESS. But insurance companies, banks and the sector of services supplied to firms are included.

The sampling frame has two levels : at the first level, production units are sampled ; at the second level, individuals employed at these sampled units are also sampled. First, concerning production units, the universe to be sampled includes all establishments with at least ten employees.

The universe is known from the SIRENE system, a French database that includes some information on all existing firms. In particular, it issues a firm identification number (SIREN) used in all French statistical sources. The sampling rate is stratified according to the sector, the region and the size of the unit ; it ranges from 1/48 for units between 10 and 20 employees to 1 for the units above 500 employees.

The sampling structure for the employees belonging to sampled units is based on employee's year and month of birth. The sample is exhaustive in small units and the sampling rate is 1/24 in the largest (above 5,000 employees).

In the sample used in our analysis, we have 10,719 production units and 318,332 employees in 1986. Due to budget restrictions, the sample size is much smaller in 1992 ; it includes 3,803 establishments or firms and 42,783 employed individuals.

In the 1986 version, annual as well as October compensation are available for each sampled employee. October compensation (salaire brut) of each employee includes all employee-paid benefits but excludes employer-paid benefits and non-wage benefits. It can be decomposed into total wage, overtime compensation and October-specific bonuses. The total annual compensation includes all non-monthly benefits and bonuses. In 1992, total annual compensation decomposed as above is available.

Furthermore, we know for each worker its occupation, firm-specific seniority, age, country of origin, work schedule (number of hours and shifts), days of absence. In addition to this individual-level information, the unit survey gives the following information : total employment, existence of shifts and night work, existence of a firm-level agreement, of a branch-level agreement... The questions in the two versions of the ESS that we use in this analysis were not always formulated identically. The main differences will be mentioned in the text if important.

More detailed technical information on the 1986 version of the ESS is available in ROTBART [1991]. The technical report on the 1992 version of the ESS is not yet available.

3 Wage Inequalities and Worker Specialization Increased between 1986 and 1992

Wage inequalities decreased in France until 1984 (see BAYET-CASES [1994]). This decrease is observed in the whole population as well within most subgroups defined by sex or education (BAYET-CASES [1994], see also KRAMARZ-MARGOLIS [1994]). The driving force of this decrease in inequality lies in the strong increases of the minimum wage (the SMIC). However, between 1986 and 1992, the situation changed, as described in LOLLIVIER [1994]. The share earned by the less paid workers has decreased for the first time since 1972 ; wage increases are most important in the neighborhood of the median value of the distribution. But, the share earned by the best paid workers is still decreasing. Lorenz curves are difficult to compare because they have an intersection. However, since they only intersect once and the coefficient of variation is greater in 1992 than in 1986, the inequality increased between the two dates (SHORROCKS, FOSTER [1987]). Other indicators, such as the ratio of the ninth to the first decile of the wage distribution, lead to the same conclusion.

This increase in wage inequalities is accompanied by some transformations of the population structure (Table 1). As before 1986, the proportion of engineers, managers and professionals continues to increase. Meanwhile, the proportion of unskilled blue-collar workers strongly decreases while that of skilled blue-collar workers increases. This evolution is new, it differs from past tendencies. Moreover, the population becomes older between 1986 and 1992 since the mean age increases from 38.6 to 39.1. But the mean seniority does not increase, as it was the case before 1986. The shares of workers with seniority above 15 years or below 5 years both increase. In-between, the proportion of workers with 10 years of tenure strongly

TABLE 1

Composition of the Sample by Sex and Qualification in 1986 and 1992

	1986			1992		
	%	mean age	mean seniority	%	mean age	mean seniority
Unskilled blue collars	33.3	36.5	10.4	17.7	37.1	10.5
Men	22.3	36.1	10.1	11.1	36.7	9.6
Women	11.0	37.3	11.2	6.6	37.8	11.9
Skilled blue collars	24.5	38.7	12.7	41.2	38.8	12.8
Men	23.1	38.7	12.6	36.1	38.7	12.6
Women	1.4	39.0	12.9	5.1	39.3	14.1
Foremen	9.7	42.0	16.4	6.3	41.8	15.8
Men	8.4	42.2	16.6	5.4	42.0	16.2
Women	1.3	41.1	14.9	0.8	40.9	13.4
Clerks	11.4	37.1	11.3	12.0	38.4	11.7
Men	3.8	38.2	12.0	4.3	39.2	12.8
Women	7.6	36.5	11.0	7.7	37.9	11.1
Technicians	10.4	38.9	13.7	10.4	39.4	13.6
Men	8.6	38.8	13.8	9.0	39.8	14.0
Women	1.8	39.2	13.4	1.4	37.1	11.2
Eng, Man and Prof	10.7	43.2	13.8	12.4	42.3	12.9
Men	9.5	43.3	14.0	10.7	42.7	13.1
Women	1.2	41.7	13.3	1.7	39.4	11.2
All Qualifications	100.0	38.6	12.4	100.0	39.1	12.5
Men	75.7	38.9	12.6	76.7	39.4	12.7
Women	24.3	37.7	11.6	23.3	38.4	12.1

Data: Insee, ESS (Manufacturing Industries and Construction) 1986, 1992.

decreases. This fact could be explained by employers keeping their oldest workers and generating greater mobility among others.

An investigation of sub-populations (LOLLIVIER [1994]) shows that inequalities increased more strongly within young people, unskilled workers and women. Before 1986, inequalities between people of a given age were decreasing, except for older groups. So, across generations, young workers faced less and less inequalities. On the contrary, between 1986 and 1992, inequalities increased among young workers. This is also true for unskilled blue-collar workers. In the past, increases of the minimum wage reduced wage inequalities and this reverses between 1986 and 1992. First, the composition of this group has probably evolved ; furthermore, the minimum wage did not increase very strongly between the two surveys. Such evolutions suggest an increase of dualism in the labor market, with a primary market giving higher wages and careers and a secondary market with lower wages and a greater job instability. Women, young and unskilled workers are particularly concerned by this secondary market. Moreover, the share of the variance of wages explained by workers' characteristics decreased between 1986 and 1992 (Table 2). In particular, age and seniority are less important in 1992 than in 1986. On the other hand, employer effects become more important as if a trade-off between employer effects and individual effects was operating. This substitution reinforces the role played by labor market dualism and the importance of employer policies in wage-setting.

TABLE 2

Proportion of the Variance of Log(wages) explained (%)

	1986	1992
Sex	7	5
Sex qualification	66	57
Sex age	19	14
Sex seniority	15	10
Firm Effect	38	48

Data: Insee, ESS (Manufacturing Industries and Construction) 1986, 1992.

To explore more specifically the issue of specialization and dualism, we compute measures of specialization proposed by KREMER and MASKIN [1994] for different subgroups. The question is the following: can we say that groups of workers, say engineers, professionals and managers, tend to work more and more in separate firms. The measures are

$$spe_k = \frac{\sum_j n_j (p_{jk} - p_k)^2}{n p_k (1 - p_k)}$$

where p_k is the proportion of skill k in the economy, p_{jk} is the proportion of skill k in employer j [1].

Results are presented in Table 3. They are striking. Specialization increased massively between 1986 and 1992. Unskilled blue-collar workers are more and more separated from other type of workers, and therefore,

TABLE 3

Specialization Index

	1986	1992
Unskilled blue-collar worker dummy	33.8	40.2
Skilled blue-collar worker dummy	26.0	32.6
Foreman dummy	11.1	18.2
Clerk dummy	11.3	22.2
Technician dummy	17.2	24.0
Engineers, professional or manager dummy	16.9	27.7
seniority	26.6	37.0
experience (age)	14.0	23.9

These values are the R-square of the regressions of each variable on firm dummies.
Data: Insee, ESS (Manufacturing Industries and Construction) 1986, 1992

1. Notice that this measure can also be computed as the R-square of the regression of skill dummies (1 if the worker is of skill k, 0 otherwise) on firm dummies . Then, one can also compute specialization measures with continuous variables ; for instance, specialization by age could be evaluated as the R-square of the regression of age on firm dummies.

work together in the same firms. This is true for each of our six categories of skills. The number even doubled for clerks.

Furthermore, age and seniority specialization increased. Young (resp. old, resp. low-seniority, resp. high-seniority) workers work more and more in the same firms. Within firm population is more and more homogeneous.

4 Wages in 1986 and in 1992

4.1. The Model

Our wage equations fully interact sex, skill with quadratic functions of experience (age here) and firm-specific seniority. The reason for such interactions is that, at least in France, compensation policies largely differ by gender and skill, the latter being due to the existence of collective agreements (both at the national and the branch levels) for each of the following groups : blue-collar workers (conventions ouvriers), clerks, foremen and technicians (Employés, techniciens et agents de maîtrise, ETAM) and engineers, professionals and managers (Cadres). Among the "ouvriers", such "conventions collectives" clearly distinguish between skilled and unskilled personnel. And, we will see that, for example, age is more important for workers with high skill levels as seniority matters more for workers with low levels of skills. In addition to these variables, we introduce the nationality dummies and a employer fixed-effect that will represent the employer-specific compensation policy. Hence, the model is:

$$y_{ij} = rn_i + \sum_k (a_{1k} age_{ik} + a_{2k} age_{ik}^2 + s_{1k} sen_{ik} + s_{2k} sen_{ik}^2) + u_j + v_{ij}$$

where y_{ij} is the log of annual compensation, age_{ik} (resp. sen_{ik}) is the age (resp. seniority in the employing firm j) of worker i and where k represents its sex$\times$skill group, where n_i is worker's nationality, u_j is a dummy for employer j and v_{ij} is the residual.

4.2. Econometric Results

Table 4 presents the regression results in 1986 and 1992. As usually observed in wage equations, blue-collar workers wage is more seniority-based than age-based. A usual explanation is that blue collar workers possess specific human capital whereas technicians or managers are more educated and, hence, have more non employer-specific (general) human capital.

More important, is the decrease of most seniority coefficients between the two surveys. In particular, this decrease is important for most of male skill-groups : unskilled blue-collar workers, clerks, technicians, engineers, managers and professionals. Furthermore, seniority effects vanish in 1992 for female technicians, and engineers, managers and professionals.

According to LEVY-MURNANE [1992], DAVIS-HALTIWANGER [1991] and others, firm-specific policies were a very likely candidate explanation to the observed increase in US wage inequalities. Our results corroborate this point. Inequality increased in France between 1986 and 1992. A large part of this increase is due to firm-effects. The standard error of u_j went from 0.18 in 1986 to 0.23 in 1992, an increase of almost 30%. Furthermore, such numbers mean that, depending on the employing firm, the wage of two workers with the same individual characteristics may differ from more than one third.

Moreover, the standard error of the adjustment increases with time. This standard error is an index of residual inequalities, once taken into account

TABLE 4

Wage Equations in 1986 and 1992 (Log of annual compensation)

	1986	1992
French	0 (*)	0 (*)
EC (6 countries)	0.01 (4.41)	0.00 (0.15)
Spanish, Portuguese	-0.01 (-2.60)	-0.02 (-3.77)
Other Europe	-0.01 (-3.22)	-0.05 (-4.25)
Algeria, Morocco, Tunisia	-0.07 (-28.40)	-0.09 (-13.79)
Other	-0.06 (-13.08)	-0.03 (-3.48)
Unskilled Blue Collars (Men)	0 (*)	0 (*)
age	0.50 (33.33)	0.55 (8.01)
age*age	-0.02 (-36.64)	-0.02 (-7.60)
anc	1.41 (45.04)	1.13 (8.30)
anc*anc	-0.02 (-25.45)	-0.01 (-2.83)
Unskilled Blue Collar (Women)	-0.12 (-34.72)	-0.15 (-11.92)
age	0.14 (6.88)	0.73 (7.80)
age*age	-0.01 (-8.32)	-0.03 (-6.29)
sen	1.42 (31.85)	1.32 (7.76)
sen*sen	-0.03 (-18.72)	-0.03 (-5.21)
Skilled Blue Collar (Men)	0.15 (56.32)	0.13 (14.51)
age	0.77 (43.11)	0.57 (12.46)
age*age	-0.03 (-41.81)	-0.02 (-11.80)
sen	0.98 (33.26)	1.08 (14.73)
sen*sen	-0.01 (-15.45)	-0.02 (-8.29)
Skilled blue Collar (Women)	0.03 (3.14)	-0.03 (-1.74)
age	0.50 (7.44)	0.61 (4.98)
age*age	-0.02 (-7.25)	-0.01 (-2.40)
sen	0.81 (6.48)	0.84 (4.20)
sen*sen	-0.01 (-2.50)	-0.02 (-3.00)
Foremen (Men)	0.41 (105.27)	0.41 (29.20)
age	1.55 (40.17)	0.90 (5.89)
age*age	-0.05 (-31.30)	-0.02 (-3.32)
sen	0.48 (10.52)	0.55 (3.19)
sen*sen	-0.00 (-1.71)	-0.01 (-1.37)
Foremen (Women)	0.26 (33.26)	0.27 (10.44)
age	0.88 (10.28)	0.82 (2.62)
age*age	-0.02 (-5.54)	-0.03 (-2.00)
sen	0.88 (8.10)	0.56 (1.36)
sen*sen	-0.01 (-4.96)	-0.00 (-0.42)

TABLE 4 (continued)

	1986	1992
Clerks (Men)	0.14 (29.67)	0.28 (19.12)
age	1.14 (26.75)	1.61 (11.179)
age*age	-0.05 (-26.53)	-0.04 (-7.16)
sen	1.10 (16.25)	0.04 (0.20)
sen*sen	-0.01 (-6.05)	0.00 (1.13)
Clerks (Women)	0.04 (11.05)	0.08 (7.19)
age	0.90 (34.12)	0.90 (9.81)
age*age	-0.03 (-27.43)	-0.02 (-4.77)
sen	1.14 (23.74)	0.99 (6.38)
sen*sen	-0.01 (-7.89)	-0.01 (-2.94)
Technician (Men)	0.38 (109.24)	0.40 (35.277)
age	1.93 (56.24)	1.37 (12.30)
age*age	-0.05 (-32.02)	-0.03 (-6.53)
sen	0.29 (6.43)	0.24 (1.72)
sen*sen	-0.00 (-3.78)	-0.00 (0.79)
Technician (Women)	0.25 (38.40)	0.31 (13.75)
age	0.74 (11.11)	1.44 (5.99)
age*age	-0.01 (-4.92)	-0.04 (-3.109)
sen	0.91 (9.54)	0.31 (0.82)
sen*sen	-0.01 (-3.80)	-0.01 (-0.66)
Eng, Prof and Man (Men)	0.80 (250.04)	0.81 (76.39)
age	2.84 (79.03)	2.50 (23.45)
age*age	-0.05 (-43.15)	-0.04 (-3.11)
sen	0.51 (13.70)	0.03 (0.26)
sen*sen	-0.01 (-11.71)	-0.01 (-2.37)
Eng, Prof, Man (Women)	0.66 (97.40)	0.69 (39.45)
age	1.58 (18.82)	2.33 (9.91)
age*age	-0.03 (-10.16)	-0.05 (-5.20)
sen	0.73 (6.70)	-0.55 (-1.84)
sen*sen	-0.02 (-5.19)	0.01 (0.67)
R Square	0.83	0.80
Standard Error	0.17	0.20
Standard Error of u	0.18	0.23

Notes: - Student of coefficients between parentheses.
- Age and seniority coefficients have been multiplied by 100.
- Stars correspond to identification constraints.
Data: Insee, ESS (Manufacturing Industries and Construction) 1986, 1992.

individual and firm effects. So, inequalities between individuals with the same measured characteristics working in the same firm are stronger in 1992 than 1986.

5 The Analysis of Employer Heterogeneity

In previous regressions, we introduced a fixed-effect for each establishment. These fixed-effects represent a base level of wage ; hence, they are an important feature of establishment-specific wage policies. In this section, we focus on these fixed-effects and try to explain their values from establishment-level variables [2].

In order to explain such fixed-effects, we perform regressions of these fixed-effects on establishment-level variables. These variables have different meaning. First, we put the size of the establishment [3], which has a well-known and much discussed influence on wage level (see BROWN-MEDOFF [1989], DAVIS-HALTIWANGER [1996] among many others). We also add two institutional dummies, the first one denoting firm-level bargaining and the second one branch-level collective bargaining [4].

Then, in order to take into account the composition of the workforce within each establishment, we add the percentage of workers in four categories (clerks ; foremen ; technicians ; engineers, professionals and managers), the fifth one (blue collar workers) is the reference group. Since qualification was already taken into account in the regressions by crossing all independent variables with the different categories, our qualification variables only aim at capturing spillover effects such as : 'working in an establishment with a large proportion of engineers entails a higher wage'. We also control for the proportion of employees working in shifts (two, three or more), since workers in such jobs receive specific bonuses. Finally, regressions include industry dummies [5]. Results are shown in Table 5.

The size has a positive impact, as expected. Dummies for firm-level or collective agreements are all but one significantly positive. They show the positive impact of any agreement on the level of wage. Moreover, they seem to be stronger in 1992. This may be due to the great number of small establishments in 1992 [6]: indeed, an additional regression shows that among small establishments, the impact of agreements on wage is greater than among large establishments. Furthermore, this result is consistent with our

2. One could ask why we perform a two step estimation on establishment-level variables whereas we could have introduced them in the first regression. In fact, the two procedures were experienced and results are, at least qualitatively, quite similar.

3. More exactly the logarithm of the size of the establishment which proves to be more adequate than the size itself.

4. As described in ABOWD-KRAMARZ [1993], branch-level collective bargaining (*conventions collectives de branche*) "are negotiated at the national, regional or local level between a confederation of unions and a confederation of employers within a broadly defined industry group. The convention would normally cover virtually all employment at the establishments represented by the employer confederation. The Minister of Labor may, and often does, extend a *convention collective de branche* to cover employees in the same industry even if their employers were not members of the confederation [...] In addition to the *conventions de branche* a large and increasing number of firms negotiate with their employees, whether or not they are represented by one or more unions, individual agreements covering only the employees at their establishments."

5. Here, industries are larger than the scope of collective agreements so that they include both covered and uncovered establishments. Thus, industry dummies are not redundant with the collective bargaining dummy.

6. The sizes of the two establishment samples are very different. However, their distributions in terms of industry or region are close to each other. They essentially differ on the size dimension, with numerous small establishments in 1992. Nevertheless, we ran regressions imposing the same distribution in the two years and observed that the results were not substantially modified.

TABLE 5

Regression of Establishment Fixed-Effects on Establishment-Level Variables in 1986 and 1992

	1986	1992
number of observations	10,718	3,801
intercept	11.0337 (1511)	11.101 (648)
log(size)	0.0123 (7.09)	0.0124 (3.07)
firm-level bargaining	0.0146 (3.48)	0.0299 (2.62)
collective bargaining	0.0076 (1.53)	0.0896 (7.25)
clerk	0.1217 (8.42)	0.0854 (4.09)
foreman	0.1911 (9.84)	0.2658 (9.69)
technician	0.2680 (15.85)	0.1849 (6.70)
eng, prof, man	0.3194 (20.21)	0.2934 (14.04)
two shifts	0.0278 (2.71)	-0.0094 (-0.41)
three shifts	0.0523 (3.43)	0.0739 (1.98)
$\geq$ four shifts	0.0759 (3.72)	0.2419 (4.37)
R square	0.15	0.18

(Student's T in parentheses ; coefficients for sector dummies are not shown.)
Data: INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992.

previous finding of a reinforcement of employer-specific wage policies (see also Table 6 below). Besides, the composition of the workforce plays an important role in the level of fixed-effects : all coefficients are significant. Dummies for work in shifts are also positive except one. They seem to denote a widening of the wage scale during the period.

The two previous regressions consider establishment wage policies at two different dates for the whole sample. But each establishment may have experienced specific changes of its own policy. We focus on the establishments that are present at both dates in our survey. The sample comprises 618 establishments [7] for which fixed-effects and other variables are available at each date. Then, we explain the changes in fixed-effects using the evolution of establishment-level variables.

Results are presented in Table 6.

First, an increase of the size induces an increase of the fixed-effects. Besides, the signature of a firm-level agreement also entails an increase of the fixed-effect. Moreover, a variation of the composition of the workforce – foremen, technicians and more obviously engineers, managers, professionals – affects the wage policy as well as paying a premium for work in shifts. In fact, most variables which explain the level of the fixed-effects also account for their changes.

7. Here again, the decrease of the sample size accounts for the difference between the numbers of establishments available for the two regressions in Table 5. This also explains the relatively small number of them which are sampled in both surveys and which we consider here.

TABLE 6

Regression of the Variation of the Fixed-Effect on the Variations of Establishment-Level Variables between 1986 and 1992

	Estimates (Student' s T)
number of observations	618
intercept	0.0546 (1.27)
log(size)	0.0197 (4.28)
firm-level bargaining	0.0219 (1.84)
collective bargaining	0.0043 (0.19)
clerk	0.0239 (0.58)
foreman	0.1169 (2.19)
technician	0.1520 (3.29)
eng, prof, man	0.2346 (4.72)
two shifts	0.0599 (2.18)
three shifts	0.1338 (3.65)
$\geq$ four shifts	0.2269 (4.84)
R square	0.24

(Dummies for industry sectors are not shown.)
Data: INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992.

6 An Establishment-Level Analysis of Wage Policies

In order to go a little further in the description of employer-specific wage policies, we estimate wage regressions within establishments, provided that a sufficient number of employees are sampled. Moreover, when establishments are present in both surveys, we analyze the evolution of the wage policy between the two dates. Hence, we select establishments that are present in both surveys. We estimate, for each of them and at each date, the following wage equation :

$$\log(w_{i,j,y}) = C_{j,y} + a_{j,y}exp_{i,j,y} + a'_{j,y}exp^2_{i,j,y} + b_{j,y}sen_{i,j,y} + b'_{j,y}sen^2_{i,j,y}$$

$$+\alpha_{j,y}BC_{i,j,y} + \beta_{j,y}CL_{i,j,y} + \gamma_{j,y}FOR_{i,j,y} + \delta_{j,y}TEC_{i,j,y} + u_{i,j,y}$$

Subscripts denote the individual (i), the establishment (j) and the year (y).

In this equation, *sen* denotes the seniority within the current firm, *exp* stands for the experience on the labor market before entering the current firm. The age of entry in the labor market is arbitrarily set to 20, so *sen* and *exp* are linked by the following equality : $exp = age - sen - 20$.

TABLE 7

Means and Standard Deviations of Establishment-Level Estimates and Mean Values of Experience and Seniority in 1986 and 1992

	Mean 1986 (standard deviation)	Mean 1992 (standard deviation)
number of observations	132	132
intercept	11.982 (0.4289)	12.174 (0.3555)
experience	0.0050 (0.0136)	0.0052 (0.0198)
experience squared	-0.0002 (0.0007)	-0.0001 (0.0013)
seniority	0.0196 (0.0387)	0.0198 (0.0245)
seniority squared	-0.0008 (0.0058)	-0.0004 (0.0011)
mean experience	5.98 (3.11)	4.76 (2.36)
mean seniority	12.38 (4.45)	15.77 (4.00)

Data: INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992.

TABLE 8

Correlation of Establishment-Level Estimates and Mean Values of Experience and Seniority in 1986

	intercept	experience coefficient	experience squared coefficient	seniority coefficient	seniority squared coefficient	mean experience	mean seniority
intercept	1 (0)						
experience coefficient	-0.0674 (0.4424)	1 (0)					
experience squared coefficient	-0.0149 (0.8654)	-0.8189 (0.0001)	1 (0)				
seniority coefficient	-0.2164 (0.0127)	0.1970 (0.0236)	-0.0718 (0.4130)	1 (0)			
seniority squared coefficient	0.0818 (0.3513)	0.0330 (0.7072)	-0.0687 (0.4341)	-0.8155 (0.0001)	1 (0)		
mean experience	0.1872 (0.0316)	0.1286 (0.1418)	-0.0240 (0.7849)	0.2160 (0.0129)	-0.2167 (0.0126)	1 (0)	
mean seniority	0.1147 (0.1902)	-0.2585 (0.0028)	0.1849 (0.0338)	-0.2559 (0.0031)	0.2389 (0.0058)	-0.5936 (0.0001)	1 (0)

Data : INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992. (in parentheses, probability of the estimate to be greater than the above value under the hypothesis that the correlation equals zero).

TABLE 9

***Correlation of Establishment-Level Estimates and Mean Values of Experience and Seniority in* 1992**

	intercept	experience coefficient	experience squared coefficient	seniority coefficient	seniority squared coefficient	mean experience	mean seniority
intercept	1 (0)						
experience coefficient	-0.2462 (0.0044)	1 (0)					
experience squared coefficient	0.1309 (0.1347)	-0.7433 (0.0001)	1 (0)				
seniority coefficient	-0.3042 (0.0004)	0.3439 (0.0001)	-0.2246 (0.0096)	1 (0)			
seniority squared coefficient	0.1816 (0.0372)	-0.1956 (0.0246)	0.1289 (0.1408)	-0.8849 (0.0001)	1 (0)		
mean experience	0.2055 (0.0181)	0.0180 (0.8375)	0.0233 (0.7911)	0.0693 (0.4300)	-0.1090 (0.2133)	1 (0)	
mean seniority	-0.0454 (0.6056)	0.0476 (0.5879)	-0.1103 (0.2079)	-0.1414 (0.1058)	0.2353 (0.0066)	-0.5890 (0.0001)	1 (0)

Data : INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992. (in parentheses, probability of the estimate to be greater than the above value under the hypothesis that the correlation equals zero).

We add dummies for qualification groups as control variables : BC (blue-collar workers), CL (clerks), FOR (foremen), TEC (technicians).

At each date, we estimate, for each establishment, the coefficients of the wage equation. We compute the correlation between these parameters at the two dates. We also calculate the mean values of *exp* and *sen* within each establishment, since we can expect some relation between the wage policy of an establishment and its workforce' characteristics.

The wage equation includes nine exogenous variables. Therefore, in order to identify the coefficients for one establishment, we need at least nine employees belonging to that establishment in our survey. Indeed, the more observations, the smaller the standard-errors of the estimates. But as we require more observations per establishment, fewer of them fulfill the condition. There is a trade-off between the number of employees (thus the estimation accuracy) and the number of establishments available in our survey for estimation. Several specifications were used, one of which is presented below. We estimate the wage equation in 132 establishments present in both surveys, with at least 25 employees in 1986 and 20 in 1992. The requirement is less stringent in 1992 since, in that survey, the sample size is quite smaller. On the whole, results were the same whatever the number of employees required. First, we provide simple statistics about the estimates.

TABLE 10

Correlation of Establishment-Level Estimates and Mean Values of Experience and Seniority between 1986 and 1992

		1986						
		intercept	experience coefficient	experience squared coefficient	seniority coefficient	seniority squared coefficient	mean experience	mean seniority
1992	intercept	0.0002 (0.9981)	-0.0018 (0.9833)	0.0215 (0.8068)	0.0728 (0.4067)	-0.0764 (0.3838)	0.1130 (0.1970)	0.0443 (0.6139)
	experience coefficient	-0.0465 (0.5964)	0.1464 (0.0940)	-0.1415 (0.1056)	0.0471 (0.5919)	0.0194 (0.8249)	0.1219 (0.1639)	-0.0596 (0.4972)
	experience squared coefficient	-0.0344 (0.6952)	-0.1246 (0.1545)	0.1453 (0.0965)	-0.0290 (0.7417)	-0.0179 (08387)	-0.0943 (0.2820)	0.0295 (0.7369)
	seniority coefficient	0.0969 (0.2690)	0.1146 (0.1909)	-0.1284 (0.1425)	-0.0182 (0.8362)	0.0498 (0.5704)	0.2209 (0.0109)	-0.2513 (0.0037)
	seniority squared coefficient	-0.0978 (0.2646)	-0.0289 (0.7420)	0.0527 (0.5483)	0.0416 (0.6357)	-0.0384 (0.6620)	-0.2831 (0.0010)	0.2650 (0.0021)
	mean experience	0.1774 (0.0418)	0.0480 (0.5851)	0.0227 (0.7959)	-0.0805 (0.3589)	0.0642 (0.4647)	0.2332 (0.0071)	-0.0288 (0.7428)
	mean seniority	0.0657 (0.4540)	0.1091 (0.2133)	-0.1588 (0.0689)	-0.0087 (0.9215)	0.0919 (0.2946)	-0.532 (0.5445)	0.0577 (0.5111)

Data : INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992. (in parentheses, probability of the estimate to be greater than the above value under the hypothesis that the correlation equals zero).

At each date, we observe, as usual, concave profiles for experience or seniority impact on wages. Besides, we observe a strong increase of mean seniority and a decrease of mean experience. We will discuss later (in Table 11) the changes in these parameters.

We provide first the matrix of correlation at each date.

We observe, as expected, a negative correlation between the coefficients of experience and experience squared, as well as between seniority and seniority squared. Besides, the intercept and the coefficient of seniority are negatively correlated : this shows a trade-off between the base level and the slope of the wage curve as a function of seniority (see ABOWD, KRAMARZ, MARGOLIS [1994]). We do not observe the same effect with experience but remember that the variable we use here measures experience on the labor market before entering the firm. The mean seniority is negatively correlated with the coefficient of seniority : firms which keep their workers for a long time reward seniority at a lower rate. Means of experience and seniority are strongly negatively correlated. This explains the correlation between mean experience and the coefficient of seniority. Coefficients of experience and coefficients of seniority appear to be correlated but this result is not systematic.

In order to analyze the stability of the estimated parameters, we calculate their correlation between the two dates.

TABLE 11

Correlation of Variations of Establishment-Level Estimates and Mean Values of Experience and Seniority

	Δ(intercept	Δ(experience coefficient)	Δ(experience squared coefficient)	Δ(seniority coefficient)	Δ(seniority squared coefficient)	Δ(mean experience)	Δ(mean seniority)
Δintercept	1 (0)						
Δ(experience coefficient)	-0.1386 (0.1131)	1 (0)					
Δ(experience squared coefficient)	0.0914 (0.2975)	-0.7356 (0.0001)	1 (0)				
Δ(seniority coefficient)	-0.3212 (0.0002)	0.1905 (0.0287)	-0.0857 (0.3286)	1 (0)			
Δ(seniority squared coefficient)	0.1436 (0.1005)	-0.0253 (0.7731)	0.0005 (0.9957)	-0.7842 (0.0001)	1 (0)		
Δ(mean experience)	0.0615 (0.4839)	-0.0361 (0.6809)	0.0772 (0.3792)	0.1291 (0.1403)	-0.2019 (0.0203)	1 (0)	
Δ(mean seniority)	-0.0088 (0.9204)	-0.0966 (0.2706)	0.0313 (0.7216)	-0.1087 (0.2146)	0.1095 (0.2114)	-0.6414 (0.0001)	1 (0)

Data : INSEE, ESS (Manufacturing Industries and Construction) 1986, 1992. (in parentheses, probability of the estimate to be greater than the above value under the hypothesis that the correlation equals zero).

Table 10 shows that few coefficients are identical or close at the two dates. Once again, this could come from the small number of observations. Yet, coefficients of experience are positively correlated across years, as well as the mean values of experience, which denotes a relative stability of these parameters. In other cases, the estimated correlation is not significant or rather difficult to explain.

To describe a little further the evolution during the period, we focus on the changes in the parameters between 1986 and 1992 (Table 11).

From Table 7, we could see that the intercept increased during the period. Here, we observe a negative correlation between the variation of the intercept and the variation of the seniority coefficient : the average level of wage increased at the expense of the reward of that individual characteristic. This may be explained by the trade-off between the base level and the slope of the wage curve, and stresses the decline of individual variables. One might suggest that results presented in Table 7 are inconsistent with what is observed in Section 4 (*i.e.* a strong decrease of seniority coefficients). However, in Table 7, standard errors relative to the coefficient are so high that a significant decrease would be difficult to prove. But the present correlation between the variations of intercepts and seniority coefficients, in association with an increase of the intercept, would indicate that the importance of seniority declined. Besides, there is a strong negative correlation between the changes of the mean experience and of the mean seniority. The overall trend (cf. Table 7) consists in a dramatic increase of mean seniority and a decrease of mean experience, which here, results in an increase of the mean age, as already observed in section 3. This evolution is consistent with decisions of hiring few young workers and firing workers with relatively little firm-specific seniority.

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