Evaluation of the effects of payroll tax subsidies for low wage workers

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Abstract

In this article, we study the impact of payroll tax subsidies for low-wage workers on various outcomes of the firms, including employment, in particular that of young and less skilled workers. We concentrate on the effects of the 1995 and 1996 tax cuts policies, which permit large decreases in employer-paid contributions compared with the 1993 original policy. To distinguish between firms that are more or less concerned by these reductions, we compute for each firm in 1994 the changes in total labor costs, which are solely due to the changes in the tax reductions between 1994 and 1997. This variable lies between 0 and 10%, depending on the proportion of low wage workers in the firm in 1994.

This paper includes an important methodology part. It refers to the statistical framework of Rubin (1974, 1977, 1983) and Heckman, Ichimura and Todd (1997, 1998, 1999), appropriate for the evaluation of a unique treatment, like participation in a training program. For identifying the effects of the firm's *ex ante* labor cost reduction, we adapt the formalism to the case where the economic policies involve an infinite number of possible treatments. We also propose an estimation method based on the implementation of nonparametric series estimators. The empirical analysis makes use of matched employer-employee information originating from two main sources of data, the "Déclarations Annuelles de Données Sociales" (DADS) and the "Bénéfices réels normaux" (BRN). We find that, between 1994 and 1997, tax reductions are associated with very strong employment and wages effects in the economy. We also find that they permit the creation or the save of roughly 400.000 works.

Keywords: Total labor costs, Tax subsidies, matched employer-employee data, selection bias, econometric evaluation methods, semiparametric estimation, series estimators, continuous treatment

Classification JEL : C14, C20, H22, J23, J31

1. Introduction

To counteract the disappearance of unskilled jobs, payroll tax subsidies for low-wage workers were enacted in France in 1993. By gradually reducing the employer-paid contributions, they have permitted a substantial decrease in the labor costs of the lower paid. Tax subsidies increased considerably between 1995 and 1996. Indeed, in France, for a worker paid at the minimum wage, the subsidy increased from 5.4% of the wage in 1993 to 18.2% at the end of 1996. Budgetary spending devoted to tax exemptions also rose sharply over the nineties. The cost of tax cuts was almost 38 billion francs in 1996, as against 4 billion in 1993. It is still rising, although the impact of the measures already implemented is not really known.

In this article, we study, on the basis of observed evolutions, the impact of the 1995 and 1996 payroll tax subsidies on different variables such as value added, employment, capital, productivities of these factors and the proportions of different categories of employees broken down by age, skill level and sex. This study has been carried out at firm level, using information from very comprehensive data sources covering hundreds of thousands of firms and millions of employees. The identification of the effects of these measures relies on the fact that firms are affected differently by tax reductions because they do not all have the same wage distribution.The firms that should *a priori* benefit most are those with the highest proportion of lower-paid workers. To distinguish firms that are more between them, we computed for each firm in 1994 the changes in total labor costs, namely the *ex ante* reduction in average labor costs, which are solely due to the changes in tax reductions between 1994 and 1997. Our analysis then consisted of comparing the evolutions over the period 1994-1997 of employment, proportions of unskilled, young workers, etc. of different ex ante reduction firms. As firms have characteristics that influence both the ex ante labor cost reduction and the evolution of the variables of interest, it is necessary to isolate the specific effect of tax reductions. Our approach therefore consists of separating in the observed evolution of employment the part relating to the characteristics of the firm from the part relating to the tax reductions.

The statistical framework that we use refers to the one initially proposed by Rubin (1974, 1977, 1983) and recently developped by Heckman and his coauthors in a series of papers on evaluation methods (see Heckman, Lalonde and Smith (1999) for a survey, see also Heckman, Ichimura and Todd (1997, 1998) and Heckman, Ichimura, Smith et Todd (1999)). These methods are particularly well suited to the problems of selectivity and heterogeneity of the effect of the measures being examined. However, they are adapted to the case of a unique program generally called treatment, like participation in a training program, which affects only one part of the population, the other part being exempt. Hence, they are not directly usable for the evaluation of tax reductions, which concern all firms in the economy but differentially depending on their proportion of low-wage workers. To identify the effects of tax reductions, we adapt the formalism to the case of an infinite number of possible treatments. Our extension defines various parameters of interest and the conditions for their identification. We also propose an estimation method based on the implementation of non parametric series estimators (Andrews (1991)).

Our main conclusions are as follows. We find that the payroll tax subsidies have permitted a considerable number of job creations in France over the period 1994-1997. We show that these resulted from substitution effects between different categories and also, although on a smaller scale, between capital and labor. We also find that the job creations are explained by scale effects linked to falls in prices, which themselves are induced by the reduction in production costs.Finally, our results show that 470,000 jobs have been created or safeguarded in the economy between 1994 and 1997 as a result of the tax reductions for low-wage workers . Our evaluations are therefore very similar to those derived from general equilibrium models[1]. For example, Laffargue (2000) arrives at between 116,000 and 440,000 long-term job creations depending on the value of the elasticity of substitution between unskilled labor and the other factors. Audric, Givord and Prost (2000) obtain very similar results, ranging from 120,000 to 410,000 job creations. Using individual company data, Laroque and Salanié (1999) put the effect of these measures at 490,000 jobs. Also using individual data but a different methodology, Kramarz and Philippon (1999) highlight a similarly high degree of sensitivity of employment for the lower paid to their cost but without proposing macroeconomic quantification.

In the second part of this paper, we begin by setting out the statutory framework for the employer's payroll tax reductions for the low-wage workers . We also describe in detail the calculation of the firm's *ex ante* labor cost reduction and show its distribution in the sample of firms used. In Part 3, we use a factor supply-and-demand model identifying several categories

[1] Our evaluations are nevertheless not directly comparable with those derived from general equilibrium models. Unlike the *ex ante* evaluations made using general equilibrium models, our *ex post* evaluations do not take into account feedback loop effects or the financing of the measures.

of employee to describe the relationship between the variables of interest (employment, remuneration, value added, etc.) and the *ex ante* labor cost reduction. In Part 4, we describe the data sources used to construct the samples as well as the definition of the variables of interest and the control variables chosen. In Part 5, we show the initial results, assuming for simplicity that the relationship between the variables of interest and the *ex ante* cost reduction is linear. We then comment on the effects of payroll tax subsidies on the various firm level variables and, on the basis of these estimates, give a preliminary evaluation of the number of jobs created thanks to the measures for reducing taxes . In Part 6, we remove the assumption of homogeneity of the marginal effect of the *ex ante* cost reductions. For this, we base ourselves on the work of Rubin and Heckman, which we expand to analyze the effects of payroll tax subsidies. An important stage in these methods consists of describing the economic policy variable, the *ex ante* labor cost reduction, through the use of control variables. In Part 7,we present the results of this descriptive stage. In Part 8, we present the results obtained using a semi-parametric approach.

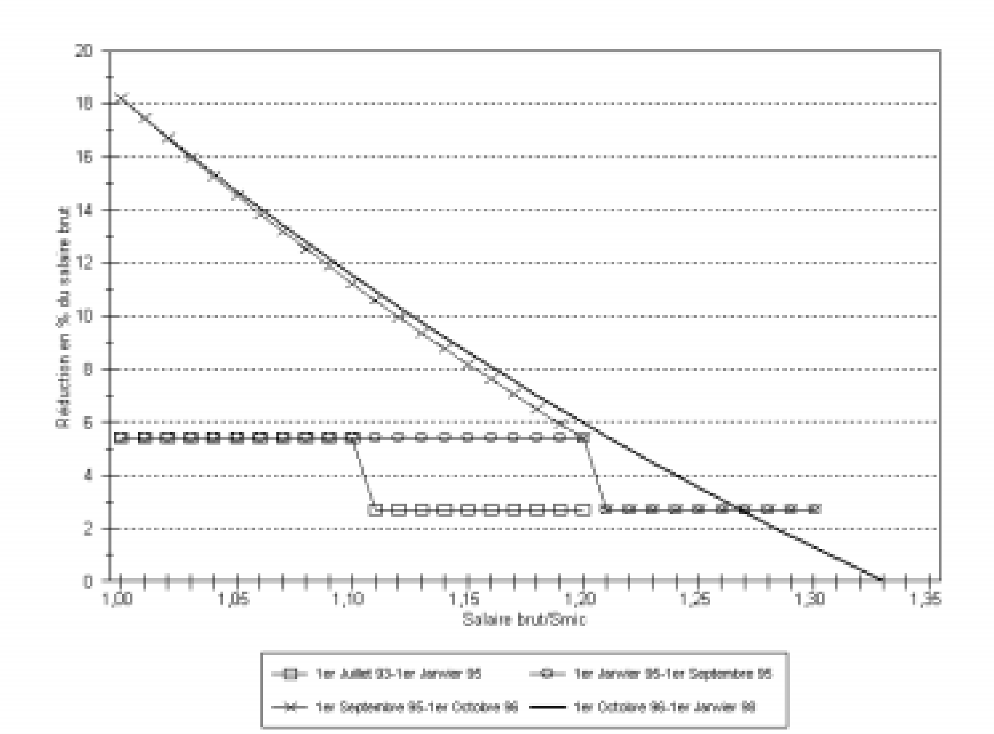
2. Payroll tax subsidies for low wage workers in France

2.1 The measures taken since 1993 to reduce payroll taxes for the lower paid

Starting in 1993, France implemented various measures to reduce payroll tax rates for low wage workers. The measures introduced in 1993, which came into force on 1 July, consisted of an exemption of 5.4 points in employers' payroll taxes for monthly wages below 1.1 times the minimum wage and a halving for those in the range between 1.1 and 1.2 times the minimum wage. This program was extended in 1995 through the raising, as of 1 January, of these two thresholds to 1.2 and 1.3 times the minimum wage respectively and the addition, as of 1 September, of a new degressive reduction between 1 and 1.2 times the minimum wage. These two programs were merged on 1 October 1996 into a single degressive reduction up to 1.33 times the minimum wage. The reduction is at its maximum at the level of the minimum wage with a reduction of 18.2 points in payroll taxes . Initially planned to last until 1 January 1998, this single degressive reduction has now been made indefinite.

Figure 1 shows the various measures introduced in France between July 1993 and January 1998. It indicates the reduction in employer's payroll taxes for a worker paid between 1 and 1.33 times the minimum wage over the period 1993-1997. It shows that payroll tax subsidies were modest between July 1993 and September 1995, but became substantial from then on.

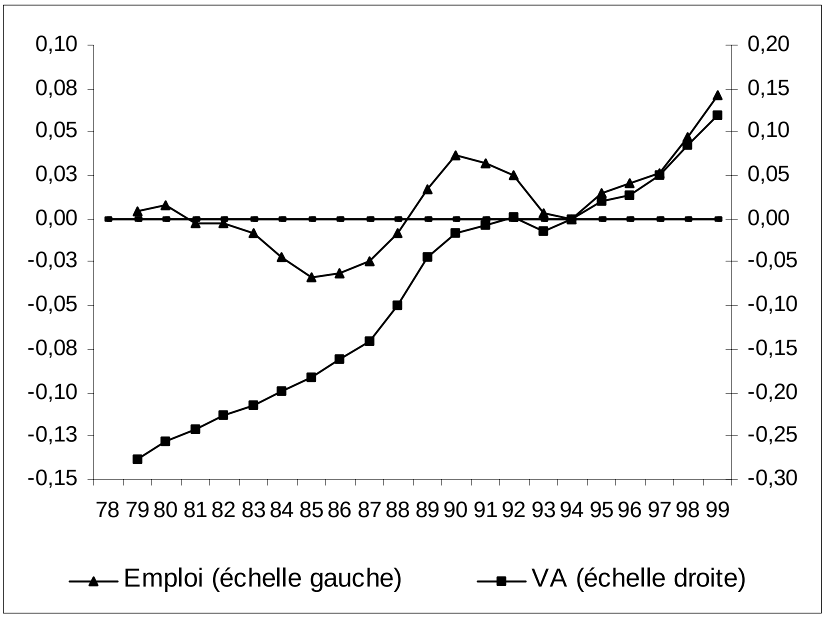
Figure 1 – Rules for the reduction in payroll taxes form 1993 to 1998



2.2 The general macroeconomic tendency at the time of the introduction of the policy.

The payroll tax subsidies for low wage workers took place in a macroeconomic context that had worsened considerably (figure 2), following a long period of stagnation in activity since the beginning of the 1990s that lasted until 1994. During this period, employment had fallen sharply. The year 1994 marked a break in this tendency with a substantial upturn in employment.

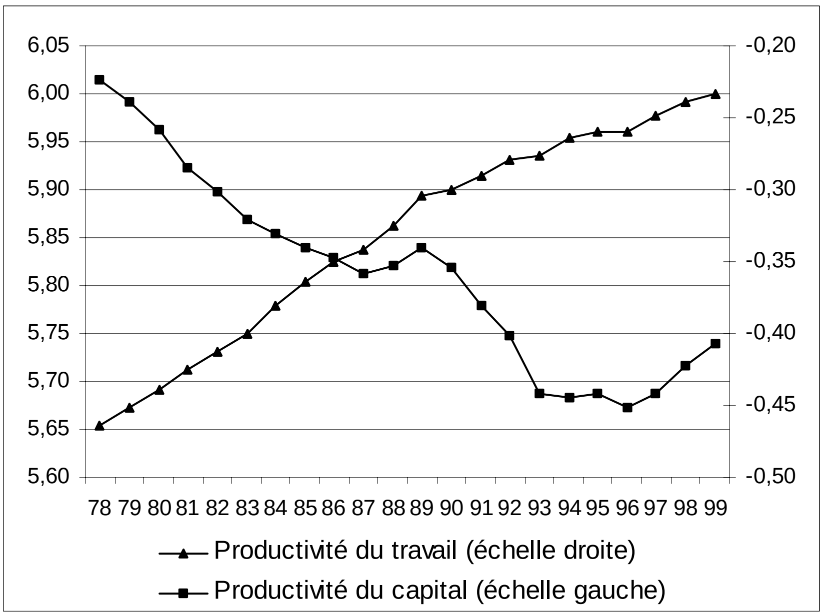
Figure 2 - Employment and Value added between 1978 and 1996, private sector



Source : National Account (logarithm normalized to zero in 1994)

Similarly, the decline in the productivity of capital stopped starting 1994, together with a rise in the job-content of growth (figure 3). The evolution in the productivity of capital is particularly interesting. The downward trend that had been virtually uninterrupted since the beginning of the 1980s was suddenly halted in 1993 and even turned upward starting in 1997.

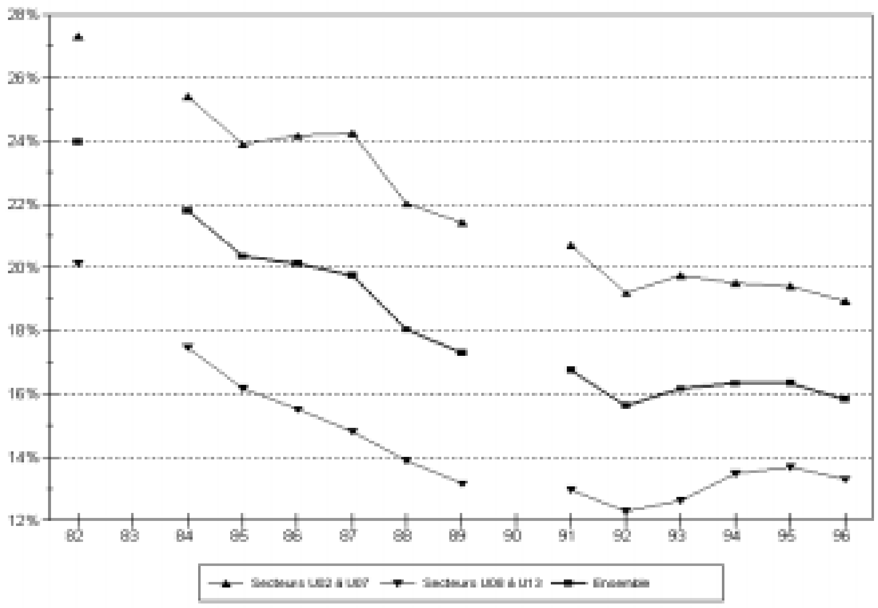
Figure 3 - Labor and capital productivity between 1978 and 1996, private sector



Source : National Account (logarithm)

Another feature was a halt to the decline in the proportion of unskilled workers in total employment at the beginning of the 1990s (figure 4).

Figure 4 - Share of unskilled workers between 1982 and 1996



Source : Audric, Givord et Prost (2000) from the DADS : full time employees of the non agricultural private sector. Not available for 1983 and 1990.

2.3 A subsidy paid to firms

While aggregate evolutions clearly show an upturn in employment and especially in unskilled employment concomitant with the introduction of payroll tax subsidies, they do not, on the other hand, make it possible to attribute them to this cause. The aim of this study is to evaluate, using individual company data, the effects of these measures on different characteristics of firms, including employment and especially employment of the unskilled and the younger workers.

For this purpose, we calculated, at individual firm level, the average *ex ante* labor cost reduction in 1994 associated with the measures taken in 1995 and 1996. In formal terms, using the gross wage $w_{j,i,94}$ of employee j in firm i in 1994, taken from the DADS (Annual Declarations of Social Data), we calculated two labor costs according to the payroll tax rules for 1994 and 1997. That is:

$$c_{j,i,94}(L_{94}) = (1 + T_{94}(w_{j,i,94}))w_{j,i,94}$$

$$c_{j,i,94}(L_{97}) = (1 + T_{97}(w_{j,i,94}))w_{j,i,94}$$

where $T_{94}(w_{j,i,94})$ (and $T_{97}(w_{j,i,94})$) is the payroll tax associated to the gross wage level $w_{j,i,94}$ and the 1994 tax rule (respectively the 1997 tax rule). The *ex ante* reduction in the average labor cost at the firm level is then equal to:

$$t_i = \frac{\sum_{j \in i} (c_{j,i,94}(L_{97}) - c_{j,i,94}(L_{94}))}{\sum_{j \in i} c_{j,i,94}(L_{94})}$$

This reduction ranges from 0% for firms having no worker paid below 1.33 times the minimum wage in 1994 to 9.5% for firms all of whose workers are paid the minimum wage. It increases with the proportion of lower-paid workers in the firm.

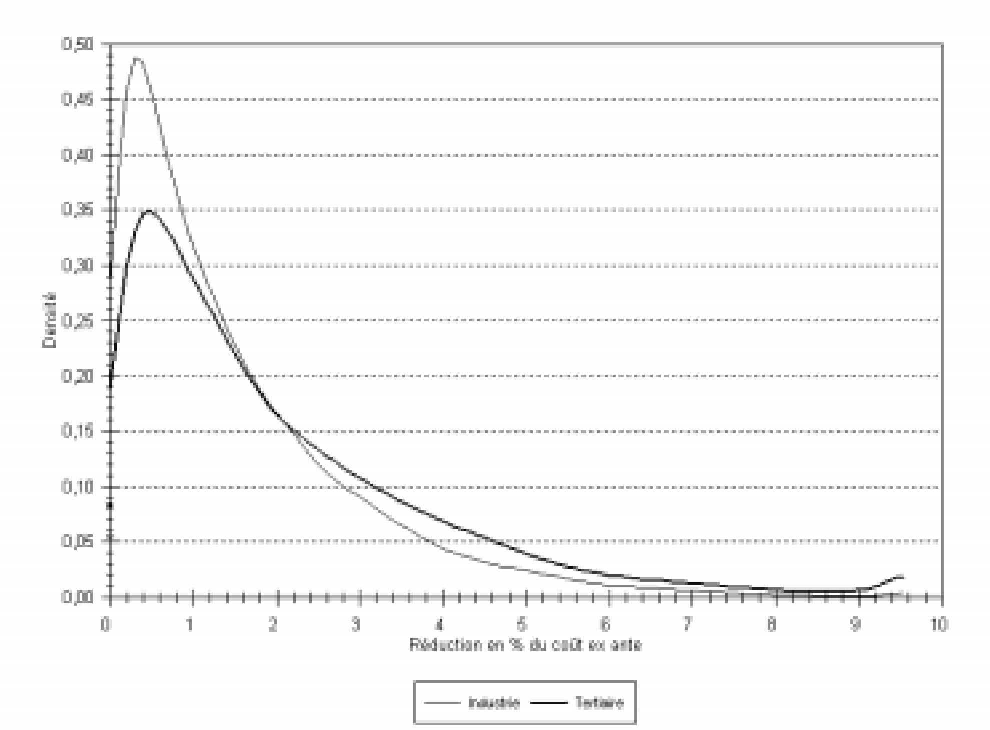
Figure 4 and table 1 show different features of the distribution of the *ex ante* reduction in average labor cost for manufacturing firms (energy and agricultural sectors excluded) and non manufacturing firms (except financial sectors). They show that this distribution is heavily concentrated around low values for the *ex ante* reduction but that the reduction is nevertheless substantial (between 1 and 6%) for around half the firms.

Table 1 - Ex ante reduction of labor cost

	0%	0-1%	1-6%	6-9.5%	9.5%
		Ensemble			
Percentage of firms	7.2	39.3	50.2	2.6	0.6
Percentage of employees	1.0	65.6	32.7	0.6	0.0
		Manufacturing			
Percentage of firms	5.6	45.9	46.9	1.5	0.1
Percentage of employees	0.8	73.3	25.6	0.3	0.0
		Non Manufacturing			
Percentage of firms	8.3	35.1	52.4	3.3	0.9
Percentage of employees	1.3	56.4	41.3	0.9	0.1

Note : Obtained from the study sample involving 87.720 firms, of which 34.371 (39%) are in manufacturing sectors and 53.349 (61%) in non manufacturing. Employment in these firms is 3.772.941, of which 2.053.777 (54%) are in manufacturing sectors and 1.719.164 (46%) in non manufacturing.

Figure 5 - Distribution of the ex ante reduction of labor cost in manufacturing and non manufacturing sectors.



Note : Kernel estimates of the density for positive values of the ex ante reduction.

3. Is it possible to identify the effect of payroll tax subsidies using only the ex-ante average labor cost reduction?

The approach we adopted to identify the effect of payroll tax subsidies consists of regressing the evolution between 1994 and 1997 of a certain number of variables of interest (employment, average labor cost, manpower structure, etc.) on the *ex ante* cost reduction and a set of control variables. In this section, we examine the conditions for the validity of such an approach.

For this purpose, we base ourselves on a labor supply-and-demand model with several types of workers which we set out in detail in annex AThe wages prevailing in a firm are assumed to be those that equalize supply and demand for each type of employee within the firm. The demand for labor is determined by firms operating under monopolistic competition in the product markets, while the supply is assumed to increase with the gross wage. The supply and demand show characteristics that are specific to the firms. We then show that on the hypothesis:

Hypothesis (H-1): the categories of employees affected by payroll tax subsidies are complementary

the evolution in the variables of interest from 1994 to 1997, such as growth in gross wages or in the employment of different categories of workers, can be written as a function of the *ex ante* cost reduction t_i and of factors Φ_i et ν_i:

$$\Delta y_i = \Delta y(\Phi_i, \nu_i, t_i)$$

where Φ_i involves structural parameters (elasticities of substitution, price-elasticity of demand and elasticities of labor supplies), the shares of the various production factors in total costs and the user cost of capital and ν_i the demand and productivity shocks. We also show that the *ex ante* labor cost reduction depends on a set of factors which we split into two groups Φ_i and ω_i:

$$t_i = t(\Phi_i, \omega_i)$$

where Φ_i denotes the above-mentioned factors and ω_i the unobserved characteristics affecting both the supply and demand for labor. Unlike ν_i, ω_i contains specific firm effects.

This model makes it possible to derive the three following results:

(i) the *ex post* variables of interest can be defined as functions of the *ex ante* average labor cost reduction once it is assumed that the categories of workers affected by payroll tax subsidies are complementary;

(ii) the variables of interest depend on numerous factors that affect also the *ex ante* labor cost reduction

(iii) there exist factors having an impact on the *ex ante* reduction but not on the *ex post* variables.

These are individual firm effects which influence the distribution of wages in 1994 and hence the *ex ante* labor cost reduction but have no direct effect on the *ex post* variables since the latter are in evolution.

These results show that it is possible to evaluate the impact of payroll tax subsidies on the evolutions of the variables of interest by comparing firms in terms of the size of their *ex ante* labor cost reduction. Given the existence of numerous factors that are common to the *ex ante* reduction and to the variables of interest, it is necessary to make comparisons between firms showing similar characteristics (i.e. all else being equal). In practice, we approach these characteristics through a set of observable variables and regress the variables of interest on these control variables and on the *ex ante* labor cost reduction. We then apply the following hypothesis:

Hypothesis (H-2): $\Phi_i = \Phi(X_i)$ ***and*** $V(t_i | X_i) > 0$

The first part of the hypothesis ensures that the control variables X_i make it possible to eliminate the selectivity bias. The second states that the X_i variables do not totally explain the *ex ante* labor cost reduction. After processing, there remains a source of random variation that is specific to the *ex ante* cost reduction: “comparable” firms can therefore benefit from different *ex ante* reductions. This hypothesis is essential for the identification of the impact of the *ex ante* reduction on the variables of interest. If the X_i variables correctly measure the common factors Φ_i, this source of variation must stem from specific firm effects (see above). The available empirical studies using matched enterprise and employee data show that this source is by no means negligible. In fact, they demonstrate the existence of wage differences that can be substantial between firms for jobs and individuals that are ostensibly identical. Abowd, Kramarz and Margolis (1999) estimate that individual skills explain only 80% of observed inter-firm wage differences.

We used four types of control variables: firms’ past characteristics, competition variables, financial variables and variables regarding skills structure. Most of these were measured in 1994 and in some cases in terms of an average evolution between a date prior to 1994 (generally the first year of presence in the sample) and 1994. The past characteristics reflect unobserved demand and productivity shocks. These are value added inclusive of labor costs, labor productivity, growth in total factor productivity and capital labor ratio . We introduce competition variables to approximate the unknown parameter of the price-elasticity of demand. These are measured at individual level using the markup and at sectoral level by the import and export ratios as well as by the entry and exit rates. In the case of the financial variables, these are intended to take into account shocks relating to the cost of capital resulting from the sharp rise in interest rates and the major changes in taxation over the period 1994-1997 (increase in the rate of flat-rate deduction in full discharge from 19.4% to 25%, in the corporation tax rate from 33.3% to 41.7% and in the rate of taxation of capital gains from 19.4% to 26%). They consist of the user cost of capital, debt as a share of total financing and the *ex ante* variation in the cost of capital, measuring the variation induced uniquely by changes in taxation over the period. Lastly, variables describing the skill structure denote the shares of different factors in the firm. These are measured at enterprise level by the shares in total hours worked for 18 categories of employees, created by intersections of the criteria of sex, skilled level and age.

4. The data

The data used are derived principally from the matching of two sources: the BRN (Real Normal Profits) and the DADS (Annual Declarations of Social Data). The BRN declarations are completed annually by firms with a turnover of more than 3.5 million francs (1992 threshold) liable for income tax in respect of BIC (Industrial and Commercial Profits). The BIC correspond to the profits declared by firms whose commercial, industrial or craftwork activity is carried out for lucrative purposes (60% of the firms, 94% of the turnover). The DADS declarations are completed annually by any firm employing workers. They cover all employers and their employees with the exception of paid agricultural workers and civil servants. Statistical processing of the DADS became exhaustive starting in 1993. At present, the employees covered by the DADS represent almost 80% of dependent employment.

The variables of interest examined in this work are shown as evolutions between 1994 and 1997 (the years immediately preceding and following the intensification of the payroll tax subsidies). They are calculated as the logarithmic differences of: real value added inclusive of labor costs , number of workers, average labor cost, capital labor ratio , labor productivity, productivity of capital, markup and unit production cost. The markup is defined as the ratio between value added and total costs (capital cost included). The growth rate in the unit cost of production is measured by the average of growth rates of labor and capital costs weighted by the cost shares of factors relative to total costs. Since this variable approximates the growth in prices, the growth in real value added is then measured by the difference between the growth in value added and the growth in the unit cost of production. All these variables are calculated on the basis of the BRN. We also examined the variation in the share in total hours worked of different categories of employees such as the unskilled, younger workers and unskilled younger workers. For this purpose we used the information provided by the DADS.

The control variables are introduced at their 1994 level and, for some of them, in terms of their average evolution over a past period. For past characteristics of firms reflecting demand shocks, we take the logarithm of value added inclusive of labor costs and its average difference. For past characteristics reflecting technical progress shocks, we take the logarithm of labor productivity as well as the average growth in total factor productivity, the logarithm and the average difference of the logarithm of the capital labor ratio. These various variables are calculated from the BRN. The competition variables are measured at individual level by the markup (absolute level and evolution). At the sectoral level, we use the import and export ratios (two-digit level), which are themselves derived from the national accounts, together with the entry and exit rates (three digit level) compiled from the BRN (see Dunne, Roberts and Samuelson (1988)). The financial variables, derived from the BRN, consist of the user cost of capital, the share of debt in total financing as well as the *ex ante* variation in the cost of capital, measuring the variation induced uniquely by changes in taxation over the period. Last, the variables measuring the workforce heterogeneity at the firm level are measured by the shares in hours worked of 18 categories of employees, built up from the DADS through the intersection of sex, three skill levels defined on the basis of their occupation and three age groups. They also comprise at individual firm level the share of wages in costs and at the sectoral level (two-digit) the average cost and share of unskilled workers, obtained by aggregation of the DADS information at this level. Annex B sets out the definitions of the above-mentioned variables of interest and the control variables. The constitution of the sample of firms is set out in annex C. It consists of 87,720 firms, of which 34,371 (39%) are in manufacturing and 53,349 (61%) in non manufacturing. These firms employ a total of 3,772,941 people, of whom 2,053,777 (54%) work in manufacturing and 1,719,164 (46%) in non manufacturing.

5. Parametric estimation of the impact of payroll tax reductions on various characteristics of firms.

We now go on to examine the results of regressing the different variables of interest on the *ex ante* labor cost reduction and the different control variables:

(1) $\Delta y_i = a\,t_i + x_i b + u_i$,

where Δy_i represents the evolution in the variable of interest between 1994 1997, t_i the average *ex ante* labor cost reduction, x_i the control variables and u_i the error term. In this initial specification, the impact of a marginal increase in the *ex ante* labor cost reduction is assumed to be constant (linear relationship) and identical from one firm to another. The results are reported in table 2.

Note that the parameters in equation (1) have no structural interpretation. The elasticities in relation to the *ex ante* cost reduction in fact combine various parameters for the supply and demand of factors that cannot be dissociated (see annex A). By examining the elasticities to the *ex ante* labor cost reduction of certain variables of interest, it is nevertheless possible to obtain an idea of the importance of the mechanisms at work, in the form of substitution and volume effects.

5.1 The observed effects of ex ante labor cost reductions

The results show a positive relationship between employment and the *ex ante* labor cost reduction. It turns out that an increase of the *ex ante* reduction of 1 percentage point leads to a rise in employment of 1.6% in manufacturing and 1.8% in non manufacturing. These substantial evolutions in employment are related to two types of mechanism. The first effect corresponds to substitutions between factors of production: the unskilled-labor content of production increases. The second corresponds to a profitability effect: the reduction in production costs enables firms to lower their prices and this in turn produces an increase in demand and hence an increase in all factors of production.

Our results show that the structure of the workforce is heavily skewed in favour of the least well-paid workers, especially the unskilled and the unskilled younger workers. The *ex ante* labor cost reductions lead to an increase in the proportion of unskilled and unskilled younger workers. These effects are substantial in non manufacturing but smaller in manufacturing, where the relationships that emerge in the case of the proportion of younger workers and the proportion of unskilled younger workers are non-significant. The *ex ante* cost reductions also leads to a substantial fall in the *ex post* cost of labor. An increase in the *ex ante* labor cost reduction of 1 percentage point leads to a fall in the average *ex post* labor cost of 2.3% in each of the two sectors. Our results indicate that the substitutions also operate between labor and capital, with the capital labor ratio falling sharply under the impact of tax reductions in each sector. Similarly, labor productivity falls and capital productivity rises. However, the relationship between labor productivity and the *ex ante* labor cost reduction is neither substantial nor statistically significant in the industrial sector.

Table 2 : Effect of the ex ante reduction in labor cost on some firm variables between 1994 and 1997.

Variables	Elasticities		Growth rate	
	Manufacturing	Non Manufacturing	Manufacturing	Non Manufacturing
Employment[a]	1.60	1.79	1.28	2.35
	(0.14)	(0.10)	(0.12)	(0.19)
Average labor cost[a]	-2.30	-2.25	-1.84	-2.96
	(0.10)	(0.09)	(0.09)	(0.20)
Share of unskilled workers	0.38	0.49	0.30	0.65
	(0.09)	(0.07)	(0.07)	(0.10)
Share of young workers	0.04	0.30	0.03	0.40
	(0.30)	(0.05)	(0.05)	(0.07)
Share of young unskilled workers	0.03	0.17	0.02	0.22
	(0.04)	(0.04)	(0.03)	(0.05)
Capital labor ratio[a]	-1.40	-1.58	-1.12	-2.07
	(0.17)	(0.13)	(0.14)	(0.20)
Productivity of capital[a,b]	1.28	1.27	1.03	1.67
	(0.17)	(0.13)	(0.14)	(0.19)
Labor productivity[a,b]	-0.12	-0.31	-0.10	-0.40
	(0.08)	(0.07)	(0.07)	(0.09)
Mark-up[a]	0.13	-0.05	0.10	-0.07
	(0.08)	(0.06)	(0.06)	(0.08)
Unit cost[a]	-1.87	-1.82	-1.50	-2.39
	(0.09)	(0.08)	(0.08)	(0.17)
Value added[a,b]	1.48	1.49	1.19	1.95
	(0.15)	(0.11)	(0.13)	(0.18)

Note : These results are obtained by the OLS regression of the variable of interest on the ex ante reduction in labor cost and a set of control variables in 1994and for some of them in evolution over the past period. They are performed on 32.459 observations in manufacturing and 48.930 in non manufacturing. Firms with a zero ex ante reduction in labor costs were discarded. The [a] superscript means that the variable is expressed in logarithm and the [b] superscript that it has been deflated by a divisia index of the unit cost of production weighting labor and capital costs by their share in total costs.

Job creations also stem from the expansion of markets resulting from price falls, which are themselves made possible by the reduction in production costs. Although there is no information on prices at firm level, it is possible, on the assumption of constant returns to scale, to approximate the markup by the ratio between value added and total costs (as already seen above). The estimations carried out indicate that the markup has not risen because of the impact of the *ex ante* cost reductions. The cost reductions associated with payroll tax subsidies have therefore been passed on into prices. Moreover, the fall in production costs was substantial, implying that the price falls were as well[2]. Consequently activity has risen sharply, notably in manufacturing.

5.2 A macroeconomic evaluation of the tax reductions of 1995 and 1996

These preliminary results make it possible to give a first macroeconomic evaluation of the number of jobs created or saved thanks to the payroll tax subsidy policy. This evaluation is obtained simply by applying the growth rate in employment that can be attributed to payroll tax subsidies to the number of jobs in the economy. For each firm, the growth attributable to payroll tax subsidies is obtained by comparing the *ex post* situation of firms Δy_i with the one that would have prevailed in the absence of tax reductions $\Delta y_i(0) = x_i b + u_i$. It is then defined by:

$$(2) \qquad E[\varpi_i(\Delta y_i - \Delta y_i(0))] = a\, E[\varpi_i t_i]$$

where ϖ_i is a normalized weighting variable: $\varpi_i = N_i / E(N_i)$ and N_i denotes the number of employees in the firm in 1994. The growth rates attributable to the tax reductions are shown in the last two columns of table 2 for manufacturing and non manufacturing.

In manufacturing and non manufacturing, the growth rates in employment attributable to the policy are estimated to be 1.3% and 2.35%, respectively. Given that dependent employment totals 5 million people in manufacturing and 10 million in non manufacturing, the number of jobs created or saved between 1994 and 1997 can be put at 300,000 for the economy as a whole, of which 65,000 in manufacturing and 235,000 in non manufacturing. By applying shares of unskilled employment in the total of 22% in manufacturing and 30% in non manufacturing, the elasticities of unskilled and skilled employment are respectively 3.5% and 1.1% in both manufacturing and non manufacturing[3]. This gives us roughly 165,000 unskilled

[2] Substantial evolutions in prices in the industrial sector explain why *ex ante* cost reductions have had little impact on labor productivity. This substantial fall in prices has appreciably limited the fall in the real *ex post* cost of labor in manufacturing. An increase in the *ex ante* cost reduction of 1% leads to a fall in the real cost of labor of only 0.5% in manufacturing, compared with a fall of 0.9% in non manufacturing.

[3] The elasticity of unskilled employment to the *ex ante* labor cost reduction is obtained by the following relationship:

jobs created or saved at the end of two years, 30,000 in manufacturing and 135,000 in non manufacturing. They therefore represent more than half the total job creations.

6. Semi-parametric estimation: a more general framework

The previous approach is based on the very restricted hypothesis of homogeneity as regards the marginal effect of the *ex ante* labor cost reduction throughout the population: $E(\Delta y_i | x_i, t_i) = a\, t_i + x_i b$. If not verified, this can lead to biased evaluations. In this part, we identify the effects of the *ex ante* reduction without having recourse to this hypothesis. We take an unknown and unspecified function of the *ex ante* labor cost reduction and the control variables: $E(\Delta y_i | x_i, t_i) = g(x_i, t_i)$.

This semi-parametric approach is based on the work by Rubin (1974,1977,1983) and Heckman, Ichimura and Todd (1997, 1999) and Heckman, Ichimura, Smith and Todd (1998) on the evaluation of discrete measures with the assumption of independence conditional on observables. The population is divided into two groups, with the "treated" group of individuals made up of individuals benefiting from the measure and the control group consisting of the other individuals. The "treatment" then takes values in $\{0,1\}$ depending on whether the individual does or does not benefit from the measure in question. The causal effect of the program on one variable of interest y_i is defined as the difference of the potential outcomes y_{1i} and y_{0i} corresponding to what would be the situation of individual if they receive or not the treatment One important aspect of this causal framework is the heterogeneity of the treatment effect among the population. As potential outcomes are not simultaneously observable, it is not possible to identify directly the individual causal effect. One important

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$$\Delta \log N_J = \Delta \log N + \Delta \log \frac{N_J}{N} = \Delta \log N + \frac{\Delta(N_J/N)}{(N_j/N)},\ \forall J$$

The growth rate of workers employed in category J, attributable to the reductions in taxes , is then given by $\left[\hat{a}_N + \frac{\hat{a}_{N_j}}{(N_j/N)}\right]\bar{t}$ where $\hat{a}_N$ and $\hat{a}_{N_j}$ are the estimations of the coefficients of the *ex ante* reduction in the relationships where the explained variables are, respectively, the evolution in the logarithm of employment and the evolution in the share of employment taken by type J. Considering two categories of workers, skilled and unskilled, the elasticities of unskilled and skilled labor to the *ex ante* labor cost reduction are respectively equal to $1.6 + (0.4/0.22)$ and $1.6 - (0.4/0.78)$ in manufacturing and: $1.8 + (0.5/0.30)$ and $1.8 - (0.5/0.70)$ in non manufacturing.

identifying assumption is independence of potential outcomes y_{1i} and y_{0i} with treatment conditional on a set of observables, or under a weaker form the independence of y_{0i} with treatment conditional on a set of observables. This identifying assumption has been extensively worked out by Heckman, Ichimura and Todd (1997, 1999) and Heckman, Ichimura, Smith and Todd (1998). In these papers, they develop a kernel matching estimator and evaluate its performance using the experimental data available from the JTPA program. Other identifying assumption allowing for selectivity on unobservables have been studied in part on theses papers and more recently by Heckman and Vytlacil (1999, 2000) In these papers they propose a general framework to address the identification issue with instrumental variables and introduce the concept of Local Instrumental Variables and Marginal treatment effect.

We propose an extension of the causal framework for the evaluation of payroll tax subsidies. Unlike the preceding programs studied by Heckman et al., the payroll tax subsidy program concern all firms but do so differently depending on the wage distribution inside the firm. The treatment, is in this case the *ex ante* labor cost reduction, and ranges between 0 and 9.5%, depending on the proportion of low-wage workers employed by the firm. To examine the effects of this policy, we propose a formalization of the problems of evaluating “continuous” measures. Unlike the discrete measures, these measures affect all individuals with an intensity that varies continuously within a given interval. Our extension consists of defining various parameters of interest and the conditions for their identification. For example, we identify the average effect of a marginal increase in the *ex ante* labor cost reduction received by each firm $E[\varpi_i \, \partial g(x_i, t_i)/\partial t]$, which is the analogue of parameter a in the previous approach. We also identify the growth rate in the variables of interest between 1994 and 1997 that are uniquely attributable to reductions in taxes: $E[\varpi_i(\Delta y_i - g(x_i, 0))]$, which is the analogue of $a\,E[\varpi_i t_i]$. For the application, we developed a semi-parametric estimation procedure based on series estimators.

6.1 Notations and definition of individual effects

In the statistical model being considered, firms i, $i = 1, \ldots, N$, can receive any treatment t falling in the interval $[\underline{t}, \bar{t}]$. In addition, the model introduces for each firm as many latent output variables $y_i(t)$ as there are possible treatments t, in other words an infinity. However, only one of these variables is observed, this being the one associated with the treatment that the firm has in fact received, i.e. $y_i(t_i)$[4]. It is from these latent variables that the various individual effects of the measure can be defined. By analogy with the Rubin causal model, one can compare for an individual the situations in which he benefits from treatments t_1 and t_0, in other words $c_i(t_0, t_1) = y_i(t_1) - y_i(t_0)$. One can also define the effect of a marginal increase in treatment t_0 i.e. $d_i(t_0) = \partial y_i(t_0)/\partial t$. One can also look at the effect of the measure for an individual: $e_i = y_i - y_i(0)$, or the effect of a marginal increase in the treatment received: $f_i = \partial y_i(t_i)/\partial t$. This formalization is very general since it makes no assumptions regarding the constancy of effects as between individuals. These effects are, however, unobservable, being defined on the basis of unobservable latent variables $y_i(t)$. On certain assumptions, however, it is possible to identify and estimate the expectations for these parameters.

We therefore defined the four following parameters of interest:

(3)	$E_1 = E(y_i(t_0))$	Average-Constant-Treatment-Output (ACTO)
(4)	$E_2 = E(\partial y_i(t_0)/\partial t)$	Average-Marginal-Effect-of-Constant-Treatment (AMECT)
(5)	$E_3 = E((y_i - y_i(0)))$	Average-Treatment-Effect (ATE)
(6)	$E_4 = E(\partial y_i(t_i)/\partial t)$	Average-Marginal-Treatment-Effect (AMTE)

The first parameter E_1 (ACTO) represents the average output that would have been observed if all the individuals had received a treatment of identical intensity t_0. The second parameter E_2 (AMECT) represents the average effect of a marginal increase in the treatment when it has a constant value equal to t_0 throughout the population. The third parameter E_3 (ATE)

[4] In the case of payroll tax subsidies, the latent variables can be, for example, all the growth rates in employment between 1994 and 1997 associated with the various possible values of the *ex ante* labor cost reduction. However, for each firm, one takes only the growth in employment corresponding to the reduction in taxes it has received *ex ante*.

represents the average effect of the treatment received $y_i(t_i)$ compared with the situation in which all individuals would have received a null treatment. Interpreting this parameter as the effect of the tax reductions therefore assumes that this latter situation $y_i(0)$ is identical to the one that would have prevailed in the absence of the program y_i^0, which amount to assume that there are no indirect effect (Heckman, Lalonde and Smith (1999)):

Hypothesis (H-3) : $y_i(0) = y_i^0$

The last parameter E_4 (AMTE) is the average effect of a marginal variation of the treatment t_i received by each individual.

6.2 Identification of the parameters of interest

We make the assumption that their exist a set of observable variables x_i conditional on which potential outcomes $y_i(t)$ and treatment t_i are independent:

Hypothesis (H-4) : $\exists x_i$ such that $y_i(t) \perp t_i | x_i, \forall t \in [\underline{t}, \bar{t}]$

Proposition (P-1):

Under hypothesis (H-3), parameters E_1, E_2, E_3 *et* E_4 *are identifiable.*

Proof. The general idea underlying proposition (P-1) is simple. We define $g(x_i, t_i) = E(y_i | x_i, t_i)$ which is identifiable from the data. For the first parameter we have: $E_1 = E_x[E(y_i(t_0)|x_i)]$. Given the independence property (H-4), we also have: $E(y_i(t_0)|x_i) = E(y_i(t_0)|x_i, t_i = t_0) = E(y_i|x_i, t_i = t_0)$. As this latter quantity is identifiable and equal to $g(x_i, t_0)$, the parameter E_1 is simply $E_1 = E[g(x_i, t_0)]$. Similarly the second parameter can be rewritten as $E_2 = E[\partial g(x_i, t_0)/\partial t]$, the third one as $E_3 = E[y_i - g(x_i, 0)]$ and the last one as $E_4 = E[\partial g(x_i, t_i)/\partial t]$. Thus they all are means of functions identifiable from the data.

Whereas in the previous approach, a particular form ($g(x,t) = a\,t + x\,b$) was given to the function g, in this new approach no functional form is specified. The counterpart of this generalization is the practical difficulty of estimating function $g(x,t)$. However, we show that it is sufficient to estimate a function of smaller dimension $g(s(x_i), t_i)$ where $s(x_i)$ is the score, properly defined. For this purpose, we generalize the property of Rosenbaum and Rubin (1983) in the case of discrete treatment to the case of continuous treatment.

Proposition (P-2):

For any indicator $s(x)$ *of observable variables such as*

$$l(t_i | x_i) = l(t_i | s(x_i))$$

where $l(t_i | x_i)$ *is the distribution of the treatment conditionally on* x_i*, the hypothesis (H-4) of independence conditionally on the observables implies that of independence conditionally on the score:*

$$y_i(t) \perp t_i | s(x_i), \forall t \in [\underline{t}, \bar{t}]$$

Proof: it is sufficient to show that $l(t_i | y_i(t), s(x_i)) = l(t_i | s(x_i))$ under the hypotheses $y_i(t) \perp t_i | x_i$ and $l(t_i | x_i) = f(t, s(x_i))$. We calculate the two quantities and show that they are equal. On the one hand we have:

$$l(t_i | y_i(t), s(x_i)) = \int l(t_i | y_i(t), x_i) l(x_i | s(x_i)) dx_i = \int l(t_i | x_i) l(x_i | s(x_i)) dx_i$$

$$= \int f(t_i, s(x_i)) l(x_i | s(x_i)) dx_i = f(t_i, s(x_i)) \int l(x_i | s(x_i)) dx_i = f(t_i, s(x_i))$$

since given the assumptions $l(t_i | y_i(t), x_i) = l(t_i | x_i) = f(t_i, s(x_i))$. Furthermore, we have:

$$l(t_i | s(x_i)) = \int l(t_i | x_i) l(x_i | s(x_i)) dx_i = \int f(t_i, s(x_i)) l(x_i | s(x_i)) dx_i$$

$$= f(t_i, s(x_i)) \int l(x_i | s(x_i)) dx_i = f(t_i, s(x_i))$$

since $l(t_i|x_i) = f(t_i, s(x_i))$.

In the case of a single treatment, the score is of dimension 1 and corresponds to the probability of treatment conditionally on the control variables. There is nothing to say that it would be of dimension 1 in the case of continuous treatment. We shall nevertheless make this assumption in the application to the payroll tax subsidies[5]. The function $g(s(x_i), t_i)$ is then bi-variate.

6.3 Semi-parametric estimators and asymptotic distribution

Our estimators are defined as sample means of suitable functional transformation of an non-parametric estimate $\hat{g}(s,t)$ of $g(s,t) = E(y_i|s_i = s, t_i = t)$. We considered the following estimators for the parameters of interest ACTO, AMECT, ATE and AMTE defined by equations (3) to (6):

(7) $\hat{E}_1 = \frac{1}{N}\sum_{i=1}^{N} \hat{g}(s_i, t_0)$

(8) $\hat{E}_2 = \frac{1}{N}\sum_{i=1}^{N} \frac{\partial}{\partial t}\hat{g}(s_i, t_0)$

(9) $\hat{E}_3 = \frac{1}{N}\sum_{i=1}^{N} (y_i - \hat{g}(s_i, 0))$

(10) $\hat{E}_4 = \frac{1}{N}\sum_{i=1}^{N} \frac{\partial}{\partial t}\hat{g}(s_i, t_i)$

We could use the popular kernel estimator of the function g :

$$\hat{g}(s,t) = \frac{1}{N}\sum_j \frac{K_h(s_j - s, t_j - t)}{\sum_i K_h(s_j - s, t_j - t)} y_j$$

However such a procedure would lead to intense computational burden as the function as to be estimated on all sample points to implement our parameters. We used non-parametric series estimators to estimate the bivariate function $E(y_i|s_i, t_i) = g(s_i, t_i)$. Function g is approximated by a polynomial function of treatment and score whose degree increases with the size of the sample. We therefore consider a polynomial base (P_k), used to define the set of regressors $P_i = (P_k(s_i)P_l(t_i))_{k+l \le d_n}$, with k and l the degrees for each of the polynomials and d_n the maximum degree of the polynomials used[6]. We then consider the coefficients $\hat{\theta}' = (\hat{\theta}_{k,l})_{k+l \le d_N}$ of the regression of variables y_i on this set of regressors in the sample: $\hat{\theta} = (P'P)^+ (P'Y)$ with $P = (P_i)$ and $Y = (y_i)$, $\forall i = 1, \ldots, N$.. The estimators for functions g and $\partial g/\partial t$ are then given by

$$\hat{g}(s,t) = \sum_{k+l \le d_n} P_k(s) P_l(t)\, \hat{\theta}_{k,l} \quad \text{and} \quad \frac{\partial}{\partial t}\hat{g}(s,t) = \sum_{k+l \le d_n} P_k(s) \frac{\partial P_l}{\partial t}(t)\, \hat{\theta}_{k,l}$$

for any value of treatment and the score. The characteristic of these estimators is that the degree d_n of the polynomials entering the approximation of functions g and $\partial g/\partial t$ increases with the size of the sample. This determines the number of regressors, which is equal to $(d_n + 1)(d_n + 2)/2$. In our application we chose a maximum degree of 6, which leads to the introduction of 28 regressors.

We referred to Andrews' work (1991) to study the asymptotic properties of the estimators of the parameters E_1 to E_4 (consistency and asymptotic normality). Parameters E_1, E_2, E_3 et E_4 are of the form $\Gamma(g)$ and the estimators $\hat{E}_1$, $\hat{E}_2$, $\hat{E}_3$ and $\hat{E}_4$ are written $\Gamma_n(\hat{g})$, where $\hat{g}$ is a series estimator of function g. Andrews (1991) gives conditions for the nature of the functional Γ, function g and the family of functions serving in his approximation $\hat{g}$ to describe the asymptotic behavior of the estimators. He examines several types of functionals, one of which comprises those that interest us here. This is $\Gamma_n(g) = \int D^\lambda g(t,s) d\eta(t,s)$, where η is a distribution of probabilities that can depend on the size of sample n and D^λ is any partial

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[5] We shall describe later the procedure adopted to estimate the score.

[6] We used Legendre polynomials, defined by the following recurrence formula:

$$(k+1)P_{k+1}(x) - (2k+1)x\,P_k(x) + k\,P_{k-1}(x) = 0 \text{ with } P_0(x) = 1 \text{ and } P_1(x) = x$$

where $P_k(x)$ is the polynomial of order k for variable x. These polynomials constitute an orthogonal base on $[-1,1]$.

derivative operator (cf. case 6 of example 2.7 on page 310 of Andrews (1991)). In our case, we have, depending on the parameters considered: $\lambda = 0$ or $\lambda = 1$ and $\eta = \delta_{t_0} \otimes \hat{f}_n(s)$ or again $\eta = \hat{f}_n(s,t)$ with $\hat{f}_n(s)$ empirical distribution of s and δ_{t_0} is a Dirac delta function in t_0. Moreover, Andrews deals precisely with the case of polynomial approximations of functions with compact support. The results then enunciated (theorems 1 and 2 and their application to example II) make it possible to establish that our estimators are convergent and asymptotically normal on certain assumptions of regularity of function g and when the degree of the polynomials increases with the size of sample at a rate below $n^{1/6}$. On the other hand, Andrews' results do not make it possible to identify the convergence rates of the estimators. In the application used in our study we use bootstrap to calculate the standard deviations by a random drawing of 500 samples and applying the same estimation procedure to each drawing.

We also defined the following weighted average effects:

$E_1^{\varpi} = E(\varpi_i y_i(t_0))$ Weighted-Average-Constant-Treatment-Output (WACTO)

$E_2^{\varpi} = E(\varpi_i \,\partial y_i(t_0)/\partial t)$ Weighted-Average-Marginal-Effect-of-Constant-Treatment (WAMECT)

$E_3^{\varpi} = E(\varpi_i (y_i - y_i(0)))$ Weighted-Average-Treatment-Effect (WATE)

$E_4^{\varpi} = E(\varpi_i \,\partial y_i(t_i)/\partial t)$ Weighted-Average-Marginal-Treatment-Effect (WAMTE)

where ϖ_i is the normaliszed weighting variable i.e. $\varpi_i = N_i / E(N_i)$ in which N_i denotes employment. The identification and estimation of these parameters requires additional assumptions. For identification, the weighting variable has to be included in the list of conditioning variables. The identifying hypothesis is then (H-5): $y_i(t) \perp t_i | x_i, \varpi_i$, $\forall t \in [\underline{t}, \bar{t}]$. For the estimation, it is now the quantity $E(y_i | s_i, t_i, \varpi_i)$ that has to be estimated and not $E(y_i | s_i, t_i)$. Although it is possible in theory to envisage non-parametric series estimators for these three variables, it is necessary in practice to make assumptions about the form taken by this function. We chose to introduce the weighting variable as an additional regressor, i.e. $E(y_i | s_i, t_i, \varpi_i) = \alpha\varpi_i + f(s_i, t_i)$.

6.4 The support

The work by Heckman et al. (1998) on kernel matching estimators in the case of a single treatment has highlighted the importance of the so-called support condition. This condition is also very important in the case of continuous treatment. The estimations of the parameters E_1 are to be compared with each other. They therefore are only of interest if it is possible to estimate quantities of the type $E(y_i(t_0) - y_i(t_1))$, in other words if one can simultaneously estimate both $E(y_i | s_i, t_i = t_0)$ and $E(y_i | s_i, t_i = t_1)$. For this, it is necessary that the score s_i of the enterprise being considered belongs to the intersection of the supports of the distribution used for firms with treatments t_0 et t_1. Since one wants to make comparisons over an interval of treatment variables ($t \in [\underline{t}, \bar{t}]$), it is necessary to examine the support of the conditional distribution of the score, knowing treatment $f(s_i | t)$, and to determine for the interval $[\underline{t}, \bar{t}]$ the common support of the corresponding distributions $S = \bigcap_{t \in [\underline{t}, \bar{t}]} \text{Supp}(f(s_i | t))$. The parameters in which we are interested in this case are then the local parameters $E(y_i(t) | s_i \in S)$ which can be compared two by two. In the case of parameter E_2, the support condition is automatically met since the conditional distribution of the score depends continuously on the treatment. However, if one wants to compare the effect of a marginal increase in the treatment for different values of the treatment, here too we have to consider the local parameters $E(\partial y_i(t_0)/\partial t | s_i \in S)$. The same is true for parameter E_3: in this case one considers $E(y_i - y_i(0) | s_i \in S)$, since for any firm it must be possible to find identical firms with small *ex ante* labor cost reductions. On the other hand, the support associated with parameter E_4 is the total support.

7. The estimation of the score

For the application of semi-parametric estimators it is necessary to determine the score. For this purpose, we assume that the distribution of the variable for the *ex ante* average labor cost reduction conditional on the control variables is a bivariate function of the *ex ante* cost reduction and an indicator of dimension 1 of the control variables, defined as a linear combination of these variables, i.e.

$$l(t|x) = f(t, x\beta)$$

Under these conditions, for any function h, there exists a function $\tilde{h}$ such that:

$$E(h(t)|x) = \tilde{h}(x\beta).$$

We consider the transformation $h(t) = \log(t/(0.10 - t))$ that defines the set of real values as support for the *ex ante* cost reduction and we assume that the corresponding $\tilde{h}$ function is the identity[7]. The score is then defined as: $s(x) = h^{-1}(x\hat{\beta})$ and its values therefore belong to the range 0-10%.

In the end we chose to use some 40 variables out of the 50 or so initially introduced. The eliminated variables are a few sectoral competition indicators as well as some financial variables. The representation of the treatment variable by conditioning variables was satisfactory, with $R^2 = 0.5$ in both manufacturing and non manufacturing. The variance of the treatment variable is then substantially reduced but an important source of variability still persists that will make it possible to compare firms with identical scores but different *ex ante* cost reductions. The results obtained are qualitatively the same for manufacturing and non manufacturing and presented in D1 of annex D.

8. The results of the semi-parametric estimation

We show first the estimations of parameters E_1 (ACTO) and E_2 (AMECT), by examining various points such as the heterogeneity of the effect of the treatment, the accuracy of the estimations and the importance of selectivity biases. We then go on to examine the effects of the payroll tax subsidies on the basis of parameters E_3 (ATE) and E_4 (AMTE). We also give a macroeconomic evaluation of the number of jobs created thanks to the measures. Finally, we discuss the support condition.

8.1 Heterogeneity of the effect, precision of the estimations and selectivity biases

Figures E1 in annex E show the estimations for parameters E_1 and E_2 relating to the growth rate of employment in non manufacturing[89]. Two important points should be noted. First, the effect of a marginal increase in treatment is not constant from one treatment value to another. Initially nil for small values of the treatment, it then rises sharply to reach a maximum value of 5% for an *ex ante* reduction of 2.5% and then decreases to stabilize at 2% for large values of the *ex ante* reduction. Second, the estimations are very imprecise. This lack of precision is particularly important at the edges of the range of variation for the treatment variable. The imprecision at the level of high values for the treatment is explained by the small number of firms available at these points to estimate functions $E(y|t,s)$. Figures F2 in annex F illustrate the problem of selectivity biases by comparing the previous estimations $E_1 = E(y(t))$ and $E_2 = E(\partial y(t)/\partial t)$ with naive estimations of the effect of treatment $E(y(t)|t)$ and the marginal effect $E(\partial y(t)/\partial t\,|t)$[10]. Despite the imprecision of estimators E_1 and E_2, we observed substantial differences from the naive estimators both as regards orders of magnitude and evolutions.

8.2 Substitution and profitability effects

Tables 3 and 4 show respectively the semi-parametric estimations of parameters E_3 (ATE and WATE) and E_4 (AMTE and WAMTE). The results (tables 3 and 4) are qualitatively similar to those of table 2. However, the orders of magnitude are often greater, although also less precise. These differences show that it is important to take into account the heterogeneity of the effect of a marginal increase in the *ex ante* labor cost reduction and therefore to apply semi-parametric estimation methods. The effect on employment is more substantial. In manufacturing, the growth rate in employment attributable to the policy now comes to 2.6% as against 1.3% using the previous parametric method. In non manufacturing, the figures are 3.4% and 2.3% respectively. This impact on employment is explained by greater substitution

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[7] It is possible to determine $x\beta$ through recourse to a semi-parametric procedure such as that of Ichimura (1989). This makes it possible to estimate models of the form $E(h(t)|x) = \tilde{h}(x\beta)$ leaving the function $\tilde{h}$ unspecified.

[8] In the application, we used 50 treatment variables spread uniformly within the interval $[0,10\%]$.

[9] The estimations relating to the various variables considered will be found in Desplatz (2000).

[10] The naive estimations are obtained by using uniquely the treatment polynomials (introduction of 7 regressors for a degree of 6 for the average unweighted effects). If the hypothesis of independence conditional on the observables holds, the difference between our estimations and the naive estimations corresponds to the selectivity bias.

between categories of employees and larger volume effects. The fall in the average cost of labor attributable to the policy is larger: in manufacturing, 3.1% as against 1.8% previously and, in non manufacturing, 4.4% as against 2.9%. Similarly, the proportion of unskilled workers increases more sharply[11]. The volume effects are also greater: the growth rate of output attributable to the policy comes to 1.9% as against 1.2% in manufacturing and 3.2% as against 2.0% in non manufacturing. This greater impact is a result of larger falls in prices and production costs of 2.4% as against 1.5% in manufacturing and 3.5% as against 2.4% in non manufacturing. On the other hand, substitution between capital and labor is less clear. While there is indeed a fall in capital labor ratio that is fairly similar to the one previously observed, the estimated effects on labor productivity and capital productivity are no longer significant. To be more precise, the effects on the productivity of capital are greater but much less precisely estimated. Similarly, whereas in the parametric approach there was an observed joint fall in the productivity of labor and the real cost of labor in non manufacturing, in the semi-parametric approach, there is no longer a significant fall in the labor productivity despite a similar fall in the real cost of labor. According to our results, the fall in the labor productivity observed at aggregate level over this period cannot be attributed to the payroll tax subsidy policy, partly because of the lack of accuracy of our estimates.

[11] For younger workers the positive effect obtained using the parametric approach disappears to become negative and significant. This negative sign is also observed for the proportion of unskilled younger workers, although this time it is not significant.

Table 3 : Semi parametric evaluation of a marginal increase of the exante reduction in labor cost

Variables	Manufacturing		Non Manufacturing	
	1	Employment	1	Employment
Employment[a]	2.87	3.30	3.27	5.16
	(0.36)	(0.82)	(0.26)	(0.68)
Average labor cost[a]	-3.73	-4.37	-4.12	-6.40
	(0.31)	(0.69)	(0.21)	(0.58)
Share of unskilled workers	0.85	0.71	0.52	0.73
	(0.20)	(0.49)	(0.13)	(0.34)
Share of young workers	-0.14	-0.54	-0.02	-0.31
	(0.10)	(0.28)	(0.07)	(0.25)
Share of young unskilled workers	-0.13	-0.64	0.19	-0.05
	(0.16)	(0.38)	(0.11)	(0.34)
Capital labor ratio[a]	-1.86	-1.58	-1.99	-2.21
	(0.49)	(1.08)	(0.34)	(0.97)
Productivity of capital[a,b]	1.28	0.89	1.57	1.69
	(0.54)	(1.20)	(0.35)	(0.95)
Labor productivity[a,b]	-0.58	-0.69	-0.42	-0.52
	(0.27)	(0.61)	(0.17)	(0.46)
Mark-up[a]	-01.7	-0.16	-0.05	-0.06
	(0.28)	(0.63)	(0.17)	(0.43)
Unit cost[a]	-2.92	-3.29	-3.41	-5.37
	(0.29)	(0.63)	(0.18)	(0.51)
Value added[a,b]	2.29	2.61	2.85	4.65
	(0.44)	(1.01)	(0.29)	(0.81)

Note : These figures are the semi parametric estimates of the parameter $E_4^{\varpi} = E(\varpi_i \, \partial y_i(t_i)/\partial t)$, *obtained with and without weighting firms by their employment. They are performed on 32.459 observations in manufacturing and 48.930 in non manufacturing. Firms with a zero ex ante reduction in labor costs were discarded. The [a] superscript means that the variable is expressed in logarithm and the [b] superscript that it has been deflated by a divisia index of the unit cost of production weighting labor and capital costs by their share in total costs.*

8.3 A macroeconomic evaluation

The results shown in table 4 make it possible to provide another macroeconomic evaluation by multiplying, as previously, the growth rate in employment attributable to reductions in charges by the number of employees in the economy. This leads to an estimate of 470,000 for the number of jobs created or saved in the economy, 340,000 in non manufacturing and 130,000 in manufacturing. However, this estimation is highly imprecise: taking into account the estimated standard deviations, the number of jobs created lies in the range from 260,000 to 690,000. At the same time, it is estimated that half the job creations were for unskilled workers. Of these 235,000 unskilled job creations, 65,000 were in manufacturing and 170,000 in non manufacturing.

Table 4 : Semi parametric evaluation of growth rates due to payroll tax reduction for low wage workers.

Variables	Manufacturing		Non Manufacturing	
	1	Employment	1	Employment
Employment[a]	4.63	2.62	2.49	3.44
	(1.14)	(0.58)	(1.03)	(0.78)
Average labor cost[a]	-5.51	-3.10	-3.67	-4.36
	(0.93)	(0.48)	(0.93)	(0.86)
Share of unskilled workers	1.30	0.69	0.87	0.69
	(0.64)	(0.30)	(0.63)	(0.34)
Share of young workers	-0.11	-0.23	-1.19	-0.50
	(0.32)	(0.14)	(0.45)	(0.22)
Share of young unskilled workers	0.37	-0.14	-1.13	-0.30
	(0.57)	(0.26)	(0.60)	(0.29)
Capital labor ratio[a]	-2.88	-1.41	-0.96	-1.61
	(1.23)	(0.66)	(1.36)	(0.84)
Productivity of capital[a,b]	1.49	0.73	0.86	1.36
	(1.23)	(0.69)	(1.33)	(0.79)
Labor productivity[a,b]	-1.39	-0.69	-0.10	-0.25
	(0.68)	(0.37)	(0.66)	(0.37)
Mark-up[a]	-0.73	-0.34	0.34	0.16
	(0.64)	(0.36)	(0.63)	(0.35)
Unit cost[a]	-4.41	-2.44	-2.56	-3.48
	(0.83)	(0.43)	(0.78)	(0.70)
Value added[a,b]	3.24	1.94	2.38	3.19
	(1.19)	(0.63)	(1.04)	(0.74)

Note *: These figures are the semi parametric estimates of the parameter* $E_3^{\varpi} = E(\varpi_i(y_i - y_i(0)))$, *obtained with and without weighting firms by their employment. They are performed on 32.459 observations in manufacturing and 48.930 in non manufacturing. Firms with a zero ex ante reduction in labor costs were discarded. The* [a] *superscript means that the variable is expressed in logarithm and the* [b] *superscript that it has been deflated by a divisia index of the unit cost of production weighting labor and capital costs by their share in total costs.*

8.4 The support condition

Finally, we examined the sensitivity of the estimations of parameter E_3 to the choice of the support. In order to estimate parameter ATE, one has to be able to find, for any given firm, similar firms that did not benefit from *ex ante* labor cost reductions. We first define the common support and see how the results are modified when taking only this support. In order not to present too many results, we show only the estimations of parameter ATE in non manufacturing.

To determine the common support, we took 28 treatment classes of identical length equal to 0.5%. In each of these classes, we calculated different quantiles (the median, first and third quartiles, the first and ninth deciles, the quantiles of order 0.25% and 99.75%), as well as the minimum and maximum values of the score. These are shown in figure E4 of annex E. This figure clearly shows the very high concentration of the score around low values for low values of treatment and the gradual widening-out as the treatment values increase.

We also defined two alternative supports. The resulting rates of elimination are shown by treatment class in table E1 in annex E. The first support is the one for which no restriction was made: the treatment and score variables then each take a value in the interval $[0,10\%]$, and no firm is eliminated. The second support still uses all the values of the score at the lower end of the distribution (the lower bound is therefore zero) but only values for the score that are below the minimum value for the percentile of order 99.75% among the different treatment classes. This leads to eliminating firms with scores above 5%. It turns out that the percentage of firms eliminated is small (4.9%). Indeed, the rate of elimination is highest in the classes with high treatment values which contain only very few firms.

Table E2 in annex E shows the results obtained for these two supports. It will be seen that the estimations are not appreciably different from one support to the other. The modifications observed are minor and largely smaller than the standard deviations of the parameters. This result stems from the fact that the distributions for the score and the *ex ante* reduction are heavily skewed towards zero (see figure F3).

9 Conclusion

In this work, based on a huge volume of data for firms and employees, we have tried to examine the effect of payroll tax subsidies on various characteristics of the firms. The study is based on the calculation of the firm's *ex ante* labor cost reduction attributable to the extension of the original policy in 1995 and 1996. It has been built up from the wage distribution inside firms in 1994. The underlying principle was to compare the results for firms benefiting from different *ex ante* reductions. For this purpose, we developed a statistical model, based on the work of Rubin and Heckman on the evaluation of economic policies. We also developed a semi-parametric estimation procedure using series estimators.

We found that payroll tax subsidies introduced between 1994 and 1997 led to the creation or the save of 470,000 jobs in the economy, half of which involved unskilled workers (representing 20% of the stock of jobs). These job creations mainly reflected substitution effects between categories of employee: the *ex-post* elasticities of the average cost of labor and the proportion of unskilled workers to the *ex ante* labor cost reduction were large. They also reflected substantial volume effects. Indeed, we obtained high elasticities to the *ex ante* reduction for production, the unit cost of production and prices.

These results are based on an identifying hypothesis of independence between treatment (the ex ante labor cost reduction) and outcomes (variables of interest associated with the different values of the treatment) conditional on observables (control variables). We used a factor supply-and-demand model to discuss this condition and the choice of control variables. Our results are sensitive to the introduction of control variables in the regressions, as shown by the differences seen between our estimations and “naive” estimations in which no control variable is introduced. They illustrate the importance of selectivity bias when directly comparing firms according to the size of the *ex ante* reductions in charges. Moreover, our results show that the impact of a marginal increase in tax reductions is very heterogeneous within the population and differs from one treatment value to another. They show the importance of tackling the question of evaluating reductions in taxes in a suitable causal framework. Introducing control variables into a direct regression of output variables on the *ex ante* cost reduction would lead to an erroneous evaluation of its effect.

The main interest of this statistical approach lies in the fact that it is not based on the specification and estimation of structural models. The effects measured are a combination of different structural parameters (elasticities of substitution, demand elasticities and factor-supply elasticities), within which it is not possible to distinguish the different components. As a result, our evaluation does not require the estimation of elasticities of substitution between various categories of worker, nor the elasticity of the demand for labor to its cost.However, its main disadvantage is to be valid only for the measure in force over the period 1994-1997. The evaluations cannot be used to study alternative policy such as, for example, the extension of the tax cuts to a broader population or, on the contrary, the intensification of payroll tax subsidies for the population already involved, or, yet again, modifications in the minimum wage. This study shows, nevertheless, that the attempts to change the distribution of earnings in the economy, especially for the lower paid, have substantial effects on employment of all categories of employees as well as on activity.

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Annex A : Economic model of labor demand and supply with heterogeneous workers.

We develop a simple theoretical model to: i)- identify the factors that affect the firm's ex ante labor cost reduction, ii)- assess the conditions under which the growth rate of the variables of interest can be defined as functions of the ex ante reduction and iii)- examine whether there exists a specific source of heterogeneity in the ex ante reduction. To keep notations as simple as possible, we omit time and individual indices.

Notations, labor demand and supply

We consider a firm with a technology of production that incorporates L various types of workers N_j and capital K. We assume constant elasticities of substitution between the various factors and constant returns to scale.

We suppose that the demand addressed to the firm has a constant price elasticity ε: $Q = \delta P^{-\varepsilon}$ where P is the price and δ a scale parameter.

$N = (N_1, \ldots, N_L)'$, $c_N = (c_{N_1}, \ldots, c_{N_L})'$ et $w = (w_1, \ldots, w_L)'$ denote respectively the vectors of employment, costs and gross wages of the L types of workers and $\pi_N = \mathrm{Diag}(\pi_{N_1}, \ldots, \pi_{N_L})$ the diagonal matrix of their cost shares relative to total cost (capital cost included) ($\sum_{j=1}^{L} \pi_{N_j} + \pi_K = 1$). Costs and gross wages are linked by the relationship: $c_{N_j} = (1 + T(w_j))w_j$, where $T(\ )$ is the payroll tax rate function. It is that function which has been changed by the introduction of payroll tax reductions for low wage workers.

The firm level demand of workers can be written as:

$$\log N = \Sigma_{N,N} \pi_N \log c_N + \Sigma_{N,K} \pi_K \log c_K + u$$

where $\Sigma_{N,N} = \lfloor \sigma_{N_j,N_l} - \varepsilon \rfloor_{j,l=1,\ldots,L}$ and $\Sigma_{N,K,} = \lfloor \sigma_{N_j,K} - \varepsilon \rfloor_{j=1,\ldots,L}$ represent the matrices of Allen substitution elasticities reduced by the price elasticity of demand and u the vector of perturbations that include technology and demand shocks.

The supply of workers addressed to each firm is defined as:

$$\log N = R \log w + v$$

where R is the diagonal matrix of the elasticities of supply to gross wages : $P = \mathrm{Diag}\left(\rho_{N_1}, \ldots, \rho_{N_L}\right)'$ and $v = \left(v_1, \ldots, v_L\right)'$ the vector of the firm specific component of wages.

Distribution of wages and the ex ante reduction of labor cost

Assuming first that wages at the firm level balance demand and supply for each type of employee within the firm and second that the cost of capital is held constant, we have :

$$\log w = \left[R - \Sigma_{N,N}\pi_N\right]^{-1}\left[\Sigma_{N,N}\pi_N \log(1 + T(w)) + \Sigma_{N,K}\pi_K \log c_K + u - v\right]$$

where $\log(1 + T(w)) = \left(\log(1 + T(w_1)), \ldots, \log(1 + T(w_L))\right)'$.

The distribution of gross wages inside firms is a function of structural parameters (elasticities of substitution, price elasticity of demand and wages elasticities of supply), cost-shares of inputs , payroll tax rates and unobserved components affecting both the demand and supply labor equations ,[12].

Using the wage equations above, we can derive in each firm the expression of wages by categories of workers at date 1994 with tax rules of 1994 and 1997. This permits to deduce the firm's ex ante labor cost reduction t, which is simply equal to the sum of the L workers-level ex ante reductions. Indeed, we have $t = t_{N_1} + \ldots + t_N$, with t_{N_j} the ex ante labor cost reduction for workers of type j defined as $t_{N_j} = \tilde{\pi}_j \left[\log(1 + T_{97}(w_{j,94})) - \log(1 + T_{94}(w_{j,94}))\right]$, $\tilde{\pi}_j$ being the share in labor costs (capital cost excluded). The firm's ex ante labor cost reduction can thus be expressed as:

$$t = t\left(\Sigma, R, \pi_{N,94}, \pi_{K,94}, c_{K,94}, u_{94}, v_{94}\right)$$

Ex post variation in gross wages and employment

The growth rate of the interest variables (mainly employment and wages of the various types of workers) over the period 1994-1997 are obtained by differentiating the wage and labor

[12] The share of each factor in total costs enters this equation. These variables could be expressed as a function of the true underlying heterogeneity sources of the model (the structural parameters and the firm specific component of heterogeneity in productivity and demand and wages). However this is not important for our analysis. The crucial point in this equation is that given the shares, the distribution of wages inside the firm is a function of the firm specific components u and v.

equations above. Assuming the labor cost- shares of the different categories of workers ($\tilde{\pi}_{N_j}$) are held constant and denoting $\tilde{\pi}_N = \mathrm{Diag}(\tilde{\pi}_{N_1},\ldots,\tilde{\pi}_{N_L})$ the vector of these shares and $t_N = (t_{N_1},\ldots,t_{N_L})'$ the vector of the ex ante labor cost reductions associated to the various types of workers, the growth rate of wages can be written as:

$$[R - \Sigma_{N,N}\pi_N]\Delta \log w = [(1-\pi_K)\Sigma_{N,N}\tilde{\pi}_N \Delta \log(1+T(w)) + \Sigma_{N,K}\pi_K \Delta \log c_K + \Delta u - \Delta v]$$

The term $\Delta \log(1+T(w))$ can be split into two terms : the variation only due to the change in the tax rule (holding gross wages constant) and the variation due to the adjustment of gross wages between 1994 and 1997 (the tax rule being that of 1997):

$$\begin{aligned}\tilde{\pi}_N \Delta \log(1+T(w)) &= \tilde{\pi}_N [\log(1+T_{97}(w_{97})) - \log(1+T_{94}(w_{94}))] \\ &= \tilde{\pi}_N [\log(1+T_{97}(w_{94})) - \log(1+T_{94}(w_{94}))] \\ &\quad + \tilde{\pi}_N [\log(1+T_{97}(w_{97})) - \log(1+T_{97}(w_{94}))]\end{aligned}$$

The first term $t_N = \tilde{\pi}_N [\log(1+T_{97}(w_{94})) - \log(1+T_{94}(w_{94}))]$ is simply the vector of the ex ante reductions corresponding to the different types of workers. Assuming then that the second term is as a function of the growth rate of gross wages i.e. $\tilde{\pi}_N [\log(1+T_{97}(w_{97})) - \log(1+T_{97}(w_{94}))] = \tilde{\pi}_N A\Delta \log(w)$, we have:

$$\Delta \log w = [R - \Sigma_{N,N}\pi_N (I - A)]^{-1}[(1-\pi_K)\Sigma_{N,N} t_N + \Sigma_{N,K}\pi_K \Delta \log c_K + \Delta u - \Delta v]$$

To define the changes in wages as functions of the ex ante labor cost reduction, we have to assume some restrictions on the elasticies of substitution between the $\bar{l}$ categories of workers concerned by payroll tax reductions :

$$\sigma_{N_j,N_k} = \sigma_{N_j,N_l} = \sigma_{N_j} \quad \forall k, \forall l \le \bar{l} \text{ et } \forall j = 1,\ldots,L$$ [13]

This is the case if workers concerned by tax reductions are perfect complement (sufficient condition)[14]. Under this assumption, the growth rates of gross wages and employment can be expressed in each firm as :

[13]These restrictions imply $\sigma_{N_l,N_k} = \sigma \quad \forall \quad l,k \le \bar{l}$ et $\sigma_{N_l,N_k} = \sigma_{N_l} \quad \forall \quad l > \bar{l}$ and $k \le \bar{l}$ where $l \in \{1,\ldots,\bar{l}\}$ is the set of worker types whose wages is below the upper threshold 1.33 times the minimum wages of the change in the tax rule (which is firm specific)

$$\Delta \log w = \left[R - \Sigma_{N,N,}\pi_N (I - A)\right]^{-1}\left[\varsigma_N (1 - \pi_K)t + \Sigma_{N,K,}\pi_K \Delta \log c_K + \Delta u - \Delta v\right]$$

$$\Delta \log N = R\Delta \log w + \Delta v$$

where ς_N is the $(L \times 1)$ vector equal to $\varsigma_N = \left[\sigma_{N_j} - \varepsilon\right]_{j=1,\ldots,L}$. The interest variables (employment, gross wages, capital, value added,…) can therefore be written as :

$$\Delta y = \Delta y(\Sigma, R, A, \pi_N, \pi_K, c_K, \Delta u, \Delta v, t)$$

Comparing factors affecting both the ex ante labor cost reduction and the interest variables , we can note that unobserved firm level components (technology and demand for the demand side and firm level wage component on the supply side) are specific to the ex ante reduction. The variables of interest are indeed taken in evolution and thus are not dependant of these specific effects. Common factors include structural parameters, shares of inputs in total cost, capital cost and demand and technology shocks. Finally, equations can be rewritten as:

$$\Delta y = \Delta y(\Phi, \nu, t) \text{ and } t = t(\Phi, \omega)$$

where Φ represent common factors, ν and ω demand and productivity shocks affecting respectively the variables of interest and the ex ante labor cost reduction. Note that firm specific components enter ω but not ν [15].

[14] Given the definition of Allen Uzawa substitution elasticities $\sigma_{i,j} = CC_{i,j} / C_i C_j$ where C is the cost function, it is straightforward to see that the condition is satisfied as long as the cost function can be written as $C\left(c_{N_1}, \ldots C_{N_i}, C_{N_{i+1}}, \ldots C_{N_L}\right) = C\left(\lambda_{N_1} c_{N_1} + \cdots + \lambda_{N_i} C_{N_i}, C_{N_{i+1}}, \ldots C_{N_L}\right)$

[15] If supply and demand shocks are modeled as: $u_{it} = u_i^{(1)} + u_i^{(2)} t + u_{i,t}^{(3)}$ et $v_{it} = v_i^{(1)} + v_i^{(2)} t + v_{i,t}^{(3)}$, then $u_i^{(1)}$ and $v_i^{(1)}$ does not enter Δu_{it} and Δv_{it} but u_{it} and v_{it}.

Annex B : Definition of the variables of interest and the control variables

We present all of the variable definitions, the variables of interest and the control variables, which are measured from two main sources of data, the BRN (Real normal profits) and the DADS (Annual Declaration of Social Data).

Variables from the BRN

The value added inclusive of labor costs (*valeur ajoutée au coût des facteurs*) is defined as the difference between unfinished production (*production immobilisée*) and materials, added to incentives for production (*subventions d'exploitation*) minus value added tax and other accrued taxes or credits for production (*impôts et taxes*). The result was divided by the national accounts industry value added price index (*prix de la valeur ajoutée*) at the two-digit level of the French industrial classification (*NAP 40 in the "Nomenclature des activités et produits"*), to yield real value added inclusive of labor costs (*valeur ajoutée réelle au coût des facteurs*).

The measure of employment is the annual average number of firm workers.

The average labor cost is equal to total labor compensation costs (salaries and payroll taxes) divided by employment.

The real capital stock measure is computed as the gross book value of fixed assets (*immobilisations corporelles brutes)*, construction, technical installations and other fixed assets, approximately adjusted for inflation on the basis of an estimated average age of fixed assets derived from the net to gross book value ratio.

The capital labor ratio is then defined as real capital stock divided by employment, labor productivity as real value added divided by employment, and operating income per unit of capital as the difference between value added inclusive of labor costs and total labor compensation, divided by capital stock.

The measurement of the user cost of capital is firm-specific, as in Auerbach (1983). It is defined as the sum of terms reflecting economic depreciation of assets and inflation and the mean of the costs of debt (*endettement*) and equity (*fonds propres d'entreprise*), weighted by their respective share in the firm's financial structure. The debt cost is proxied by the observed interest rate, the owners' equity cost by using capital income taxes and the rate of economic depreciation as the net to gross book value ratio (i.e. amortized portion of the capital).

We also calculate for each firm in 1994 the changes in the capital cost, which are solely due to changes in taxes over the period 1994-1997. It is obtained by comparing two capital costs, which are computed in 1994 from the firm accounts using the 1994 and 1997 tax rates.

Total cost in the enterprise is equal to the sum of labor and capital costs, as defined above. Using this information, we calculate the wage share in total costs and the ratio between value added and total costs to approximate the mark-up. We also compute total factor productivity growth as the difference between the growth rate of real value added and the cost share-weighted average of the growth rates of labor and real capital. Finally, we measure the growth of average cost as the cost share-weighted average of the growth rates of labor and capital prices. We also used this former variable to compute some of our interest variables like the growth rate of real value added and labor productivity, in so far as growth in average cost is a proxy of the growth rate in the value added price.

Variables from the DADS

The DADS provide information on the total net nominal earnings during the year for each employee, the individual's sex, date of birth, occupation, number of days and hours during the calendar year the individual worked in the establishment, status of the employee: full-time, part-time, intermittent workers and. We selected only full-time and part-time employees.

Using the information on the individual's occupational category (1-digit in the occupational classification), we created three skill levels: high skilled workers (included business heads, senior executives and intermediate occupations), skilled workers (skilled blue- and white-collar workers) and unskilled workers (unskilled blue- and white-collar workers). We also grouped age responses into three categories, separating the youngest employees (less than 25 years old) from the middle aged (between 25 and 49 years old) and the oldest one (50 and more). We thus obtained 18 categories of employees, by multiplying the three skill levels with the three age-level categories and sex.

We converted net nominal earnings to total compensation cost for each employee, by calculating employee and employer payroll taxes, after accounting for tax reductions for low-wage workers.

We then aggregated the variables concerning total compensation costs and hours worked by individual's category at the firm level, to yield cost- and hour- shares of the eighteen categories of employees per firm.

Finally, we used the DADS to compute the firm's ex ante labor cost reduction.

Annex C : Elimination of outliers and data set creation

Our data set was constructed from two main sources of data collected by Insee, the “Déclarations Annuelles de Données Sociales” (DADS) and the “Bénéfices Réels Normaux” (BRN). The DADS is a large-scale administrative database of employee information, which covers nearly all individuals employed in French enterprises (80% of the employed persons in the economy). The DADS file contains a variable concerning the gross earnings of each employee, in addition to many other variables such as individual’s sex, status, occupation and so on. The BRN is a firm-level database, that covers nearly 600,000 French firms with more than 3,5 millions of francs (1992 threshold) and subjected to commercial and industrial profit tax regime (Bénéfices Industriels et Commerciaux, BIC :60% of the firms and 94% of the sales). The BRN data set provides detailed information on firms current accounts and balance sheets, like value added, employment, investment, gross and net stocks of fixed assets, labor costs, operating income, etc.

The data set construction was realized in several steps. We built up a balanced sample of firms from the BRN source, selecting those which belonged to Manufacturing industries and Services (except the energy, agricultural and financial sectors) and were present in all five years 1993-1997 (we also kept information from the date of firms first appearance in the data source BRN to be able to take lags of the variables as controls). We then eliminated from our BRN sample the firms for which one our main variables had erroneous values (to keep a balanced sample, firms were eliminated all years of data). We deleted firms with non positive values for value added, number of employees, capital and wage bill. We also deleted the firms for which the annual growth rates of the value added, capital and number of employees are less than the 1^{st} percentiles of these distributions or more than the 99^{th} percentiles in the sector (NAP 40, approximately 2- digit). In the same manner, we eliminated firms which were extreme outliers in the distributions of the logarithms and the first differences of logarithms of labor productivity, capital labor ratio, labor and capital costs. These restrictions yielded to loose 50% of the firms in Industry and 60% in Services, essentially because of the “clean-up” on the capital cost. We then merged the data from BRN with DADS, selecting only firms which were in both data sources for the years 1993 to 1997. The merging eliminated 9% of the remaining firms in both sectors. We then realized a last clean-up on our variables of interest and control variables, which caused an additional elimination of 14% of firms in industry and services. Finally, about 30% of the 295,118 firms, which were available in the BRN initial balanced sample, were kept in the analysis data set. This gave us a usable sample of 87,720 firms, in which 34,371 firms (39%) belonged to Industry and 53,349 (61%) to Services. Those firms employed 3,772,941 individuals, 2,053,777 (54%) in Industry and 1,719,164 (46%) in Services.

Despite an important elimination of observations, our analysis sample provides an overall growth rate of employment which is close to that of the national accounts (see table below). Both changes indicate a fall of employment in Industry and a rise in Services. Nevertheless, our sample gives a weight too much important to Industry and not enough to Services, as compared to national accounts. Moreover, it underestimates the decrease of employment in Industry (-1,2% over the period 1994-1997 against –2,6% in national accounts) and the increase in Services (+2,5% against +5,7%). For the building sector, the growth rates of employment from both sources are however quite similar (-5,3% against –5,6%). Such differences are frequent, when comparing data from firms and national accounts.

Annex D : Results of the estimation of the score

Table D1 : Logistic regression of the treatment variable on controls (manufacturing, 32769 observations and non manufacturing, 49614 observations).

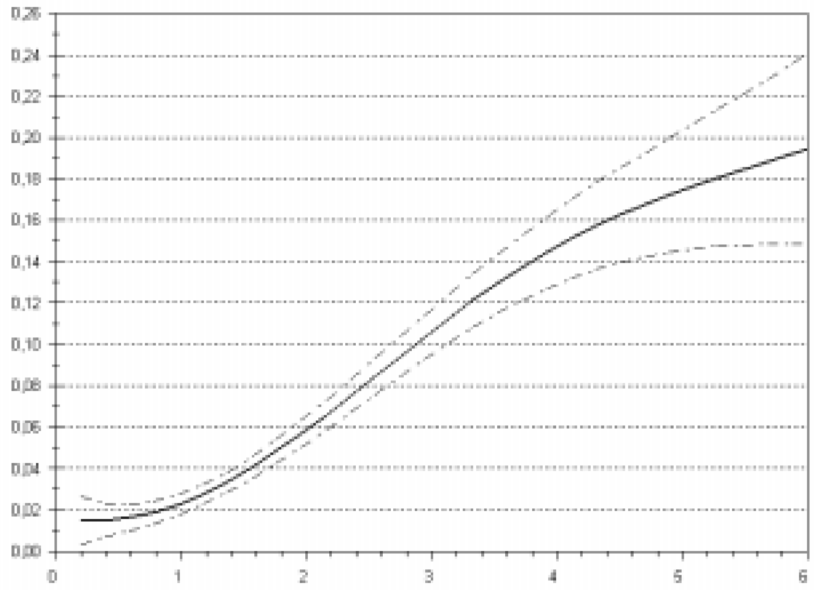
Variables	Manufacturing	Non Manufacturing
Firm level variables (general)		
Value added (log)	-0.161 (0.005)	-0.235 (0.005)
Labor productivity (log)	-2.457 (0.031)	-1.662 (0.024)
Capital labor ratio (log)	-0.066 (0.020)	-0.322 (0.017)
Operating income capital ratio	-0.153 (0.023)	-0.138 (0.012)
Value added (growth rate)	-0.462 (0.072)	-0.356 (0.056)
Total factor productivity (growth rate)	2.423 (0.112)	1.308 (0.072)
Capital labor ratio (growth rate)	0.161 (0.060)	0.405 (0.044)
Operating income capital ratio (growth rate)	-	0.072 (0.020)
Competition variables		
Firm mark-up	2.461 (0.054)	1.811 (0.027)
Firm mark-up (difference)	-1.330 (0.127)	-0.877 (0.069)
Entry rate (three digit level)	-0.264 (0.018)	1.132 (0.051)
Exit rate (three digit level)	0.093 (0.136)	-0.991 (0.094)
Import rate (two digit level)	-0.833 (0.052)	2.981 (0.134)
Export rate (two digit level)	0.213 (0.055)	-4.852 (0.175)
Work force heterogeneity		
Share of labor cost	-3.859 (0.139)	-5.192 (0.108)
Share of labor cost (difference)	1.172 (0.519)	2.703 (0.264)
Average labor cost of unskilled worker (log two digit level)	0.609 (0.078)	-0.921 (0.040)
unskilled worker (growth rate two digit level)	-0.037 (0.051)	-0.207 (0.056)
Share of young unskilled men	0.756 (0.097)	0.735 (0.076)
Share of prime age unskilled men	-0.397 (0.054)	-0.170 (0.038)
Share of old unskilled men	-0.325 (0.135)	-0.380 (0.103)
Share of young skilled men	0.174 (0.076)	0.230 (0.061)
Share of prime age skilled men	-0.906 (0.041)	-0.680 (0.028)
Share of old skilled men	-0.926 (0.071)	-0.807 (0.069)
Share of young highly skilled men	-0.120 (0.231)	-0.003 (0.141)
Share of prime age highly skilled men	-1.612 (0.053)	-1.316 (0.033)
Share of old highly skilled men	-1.737 (0.080)	-1.363 (0.053)
Share of young unskilled women	1.124 (0.158)	1.137 (0.053)
Share of prime age unskilled women	0.876 (0.066)	0.745 (0.030)
Share of old unskilled women	0.677 (0.152)	0.673 (0.062)
Share of young skilled women	1.426 (0.188)	1.248 (0.114)
Share of prime age skilled women	0.305 (0.068)	0.549 (0.042)
Share of old skilled women	-0.055 (0.143)	0.234 (0.092)
Share of young highly skilled women	2.608 (0.411)	0.630 (0.111)
Share of prime age highly skilled women	-0.942 (0.104)	-0.483 (0.034)
Share of old highly skilled women	-0.924 (0.183)	-0.980 (0.071)
Financial variables		
Cost of capital (log)	0.450 (0.027)	-0.048 (0.028)
Cost of capital (growth rate)	-0.138 (0.081)	-
Debt ratio	0.219 (0.023)	0.109 (0.018)
Ex ante variation in the cost of capital	2.291 (0.088)	1.522 (0.214)
R^2	**0.508**	**0.504**

Note : Also include sector dummy variables at the one digit level. Level variables are taken in 1994, variables in difference or growth rate are taken over the longest period available when at the firm level and over the period 1990-1994 for variables at the sectoral level.

Annex E : Semi parametric estimation results for parameters E_1 and E_2

Figure E1 : Growth rate of employment in non manufacturing, 48.930 observations

a)- Estimation of parameter $E_1 = E(y_i(t_0))$



b)- Estimation of parameter $E_2 = E(\partial y_i(t_0)/\partial t)$

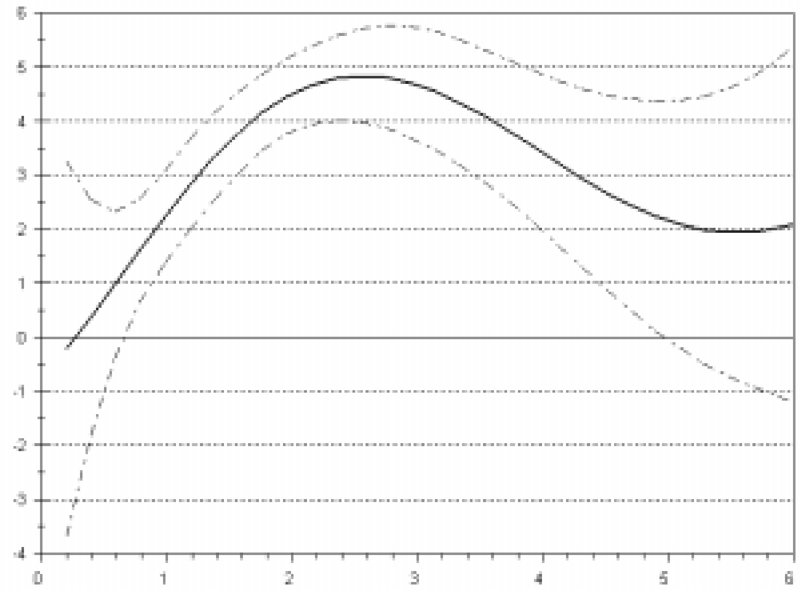
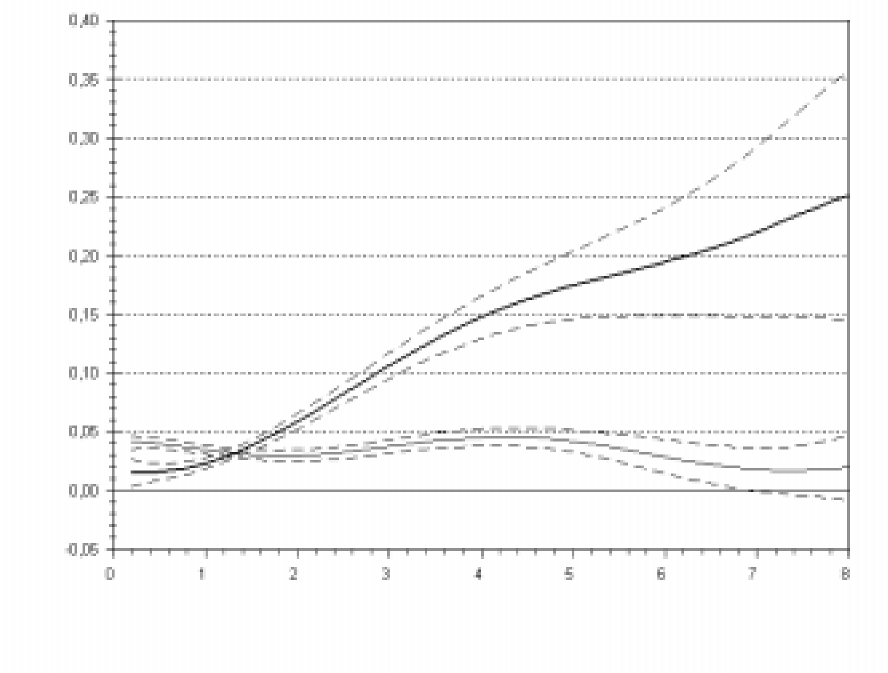


Figure E2 : Growth rate of employment in non manufacturing, 48.930 observations

a)- Comparison of $E_1 = E(y_i(t))$ with $E(y_i(t)|t)$



b)- Comparison of $E_2 = E(\partial y_i(t)/\partial t)$ with $E(\partial y_i(t)/\partial t|t)$

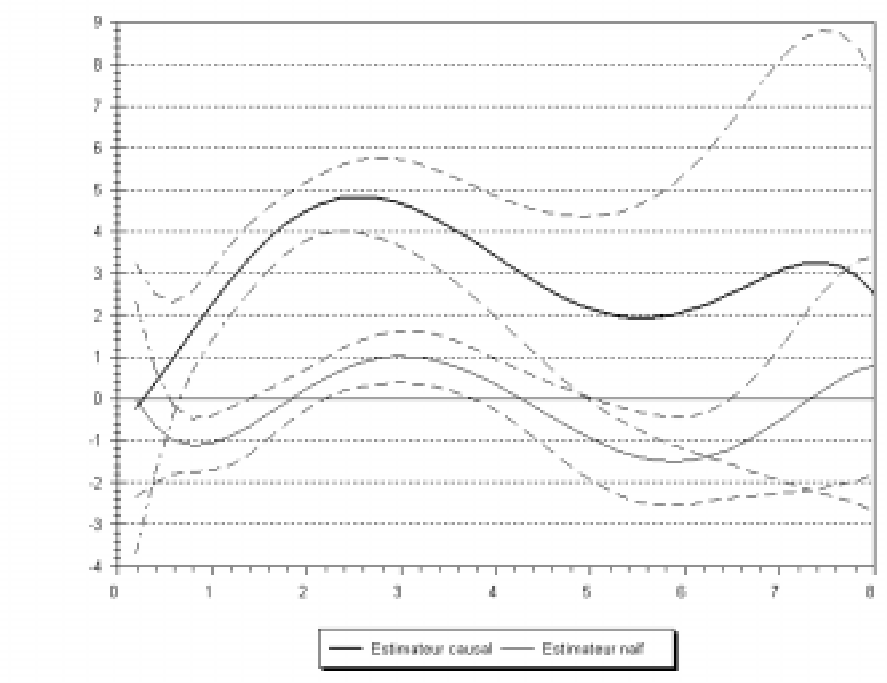
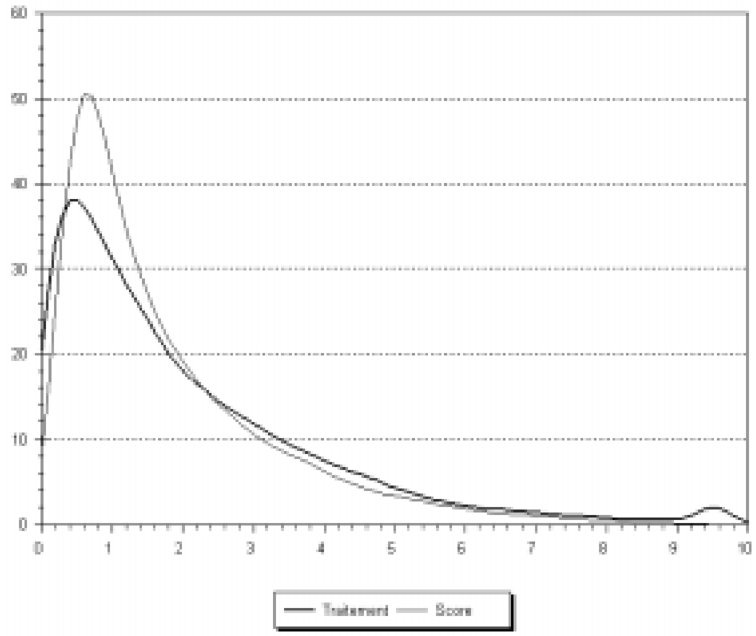
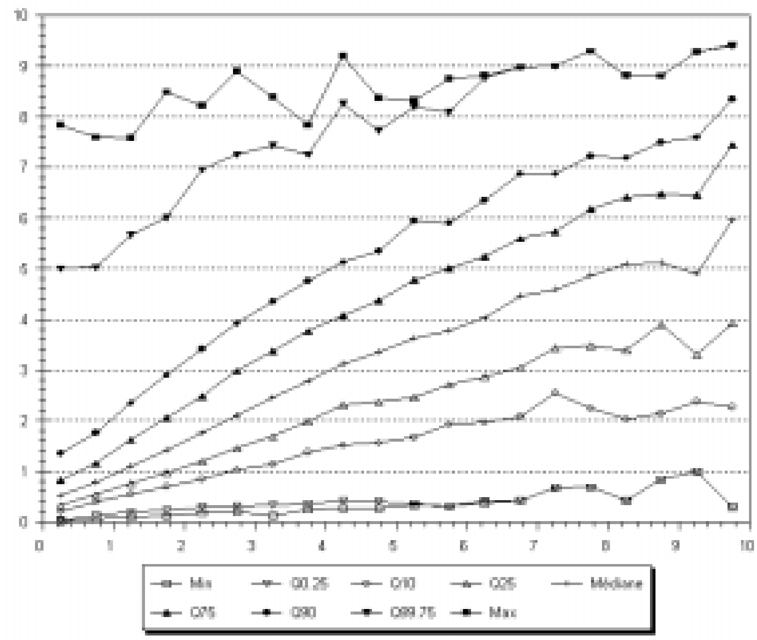


Figure E3 : Distributions of treatment and score - non manufacturing



Note : Kernel density estimation (gaussian kernel)

Figure E4 : Percentiles of the score distribution by class of treatment – non manufacturing



Note : 20 classes of treatment, width 0.5% : [0,0.5%[, [0.5%,1%[, [1%,1.5%[, [1.5%,2%[, …

Table E1 : Number of firms by class of treatment and number of firms discarded in the restricted support (non manufacturing)

	Whole support		Number and % of discarded firms in the restricted support	
0.0%-0.5%	20.4	10001	0.26	26
0.5%-1.0%	17.8	8725	0.25	22
1.0%-1.5%	13.7	6721	0.51	34
1.5%-2.0%	10.2	5011	1.14	57
2.0%-2.5%	8.2	3986	2.11	84
2.5%-3.0%	6.5	3188	3.92	125
3.0%-3.5%	5.2	2561	5.35	137
3.5%-4.0%	4.3	2092	7.55	158
4.0%-4.5%	3.3	1599	11.13	178
4.5%-5.0%	2.6	1272	14.47	184
5.0%-5.5%	1.8	870	20.92	182
5.5%-6.0%	1.3	631	25.67	162
6.0%-6.5%	0.9	462	29.44	136
6.5%-7.0%	0.8	384	37.24	143
7.0%-7.5%	0.6	291	41.58	121
7.5%-8.0%	0.4	216	48.61	105
8.0%-8.5%	0.3	149	52.35	78
8.5%-9.0%	0.3	154	52.60	81
9.0%-9.5%	0.2	118	50.00	59
9.5%-10.0%	1.0	499	61.92	309
Total	100.0	48390	4.87	2381

Table E2 : Comparison of estimations of parameter E_3 on the restricted and whole support (non manufacturing)

Variables	**Whole support**		**Restricted support**	
	1	Employment	1	Employment
Employment[a]	2.47 (0.91)	3.52 (0.78)	2.49 (1.00)	3.44 (0.78)
Average labor cost[a]	-3.14 (0.78)	-4.36 (0.86)	-3.70 (0.92)	-4.36 (0.86)
Share of unskilled workers	0.83 (0.57)	0.70 (0.33)	0.87 (0.63)	0.69 (0.34)
Share of young workers	-0.77 (0.37)	-0.43 (0.20)	-1.19 (0.45)	-0.50 (0.22)
Share of young unskilled workers	-0.52 (0.49)	-0.20 (0.27)	-1.13 (0.60)	-0.30 (0.29)
Capital labor ratio[a]	-0.81 (1.23)	-1.71 (0.82)	-0.96 (1.36)	-1.61 (0.84)
Productivity of capital[a,b]	1.77 (1.34)	1.32 (0.82)	2.13 (1.56)	1.39 (0.85)
Labor productivity[a,b]	0.96 (1.87)	-0.39 (1.10)	1.16 (2.11)	-0.22 (1.13)
Mark-up[a]	0.45 (0.57)	0.15 (0.34)	0.34 (0.63)	0.16 (0.35)
Unit cost[a]	-3.15 (1.65)	-3.39 (1.08)	-3.82 (1.92)	-3.50 (1.11)
Value added[a,b]	3.43 (1.47)	3.13 (0.96)	3.65 (1.70)	3.22 0.98)