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Spillover effect of the Minimum Wage in France: An Unconditional Quantile Regression Approach[‡]

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Abstract

This study evaluates the impact of the minimum wage on the earnings distribution, using an unconditional quantile regression method proposed by Firpo, Fortin, and Lemieux (2009). Our identification strategy relies on a unique setting that created an exogenous change in the minimum wage revaluation rule that occurred in France in the early 2000s. The gradual application of a working time reduction law resulted in the coexistence of several minimum wage levels. These levels were forced to converge to one single level between 2003 and 2005 resulting in exogenous variations of these different levels. For this specific period, we find that an increase in the minimum wage leads to significant but decreasing effects on the earnings distribution up to the seventh decile for men and up to the fifth for women.

Keywords: *Minimum wage, earnings distribution, unconditional quantile regressions.*

JEL: J31, J38, C21.

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1 Introduction

This study aims at providing new empirical stylized facts regarding the spillover impact of the minimum wage on the wage distribution during the last decade in France. To model the effect on the wage distribution, we rely on the "unconditional quantile regression" method proposed by Firpo, Fortin, and Lemieux (2009). It allows us to directly estimate the impact of a marginal change in the minimum wage level throughout the overall wage distribution, without changing the distribution of other observable characteristics. Our identification strategy uses a change in the French labor market regulation that occurred in the period 2003-2005. During this period, several levels of the minimum wage coexisted and evolved at different exogenous rates. Our results suggest significant effects of the changes in the minimum wage level that occurred at this period, up to the seventh decile for male employees of the private sector.

This paper relates to a large literature on the impact of the minimum wage on the labor market. While the original aim is to guarantee low skilled workers a "decent" standard of living, it is also considered as a redistributive tool (for a theoretical discussion see for instance Lee and Saez, 2012).[1] During the last decades, some emphasis has been placed on its potential impact on the reduction of earnings inequality. Di Nardo, Fortin, and Lemieux (1996) and Lee (1999) for instance conclude that part of the rise in wage inequality observed in the US during the 1980s is due to a decline, in real terms, of the minimum wage level. As emphasized for instance by Lee (1999), the minimum wage can have an impact on the overall earnings distribution through two main channels. First, an increase in the minimum wage level should have spillover effects on higher earnings. Beyond the *mechanical* effect at the very bottom of the earnings distribution, the wages of workers who earned more than the new minimum wage can also rise as if they benefited from a *global* increase at the bottom of the wage distribution (Katz and Krueger, 1992; Teulings, 2000, 2003). Indeed, a rise in the minimum wage lowers the relative price of high-skilled workers compared to the one of low-skilled workers, which can in turn affect their relative demand and therefore affect the wages of workers located higher in the wage distribution. Moreover, firms can be willing to maintain an upward compensation scheme as a way to stimulate the efforts of the employees. In line with the tournament model proposed by Lazear and Rosen (1981) and Rosen (1986), Chen and Shum (2010) find that a large part of intra-firm wage differentials could be interpreted as an incentive tool. This intuition is also supported by recent evidence on experimental data. According to Falk, Fehr,

[1] The French law states that the minimum wage should *"ensure that employees with the lowest wages, have a guaranteed purchasing power, and participate in the Nation's economic development".*

and Zehnder (2006), the minimum wage could modify the agents' perception of the "fair" level of remuneration (consistent with Akerlof and Yellen, 1988). It would therefore have a substantial impact on the employees' reservation wage. This could explain the presence of significant wage increases, even beyond the level imposed by the minimum wage. Second, the minimum wage can have indirect effects on the observed earnings distribution of *employed* individuals through an impact on employment which could go in different directions, either driving out low productivity workers from the labor market, or attracting previously unemployed individuals for whom the minimum wage did not meet their reservation wage. An increase in the minimum wage can thus change the composition and the size of the labor force. Beyond the reduction of earnings inequality, these spillover effects can be of great importance in terms of public policy in particular regarding the effect of the minimum wage on the overall labor cost and consequently on firm competitiveness. Lastly, the absence of spillovers was used in recent papers as an identification tool for public policy evaluation. Therefore, knowing when and where this assumption is reliable is of primary interest.

Despite the above mentioned theoretical channels that can lead to spillover effects, their very existence is still a controversial empirical issue. While there seems to be some agreement regarding the presence of spillover effects in the US, their magnitude is still debated. Lee (1999) finds substantial effects and explains that the decrease, in real terms, of the minimum wage, is the main determinant of the growth in inequality in the lower tail of the distribution. However Autor, Manning, and Smith (2010) seem to find much less evidence with a similar but more refined approach. Among the studies that try to quantify the distributional effects, some conclude that they are rather limited: up to the 5th to 10th percentiles in Card and Krueger (1995), and up to 1.2 to 1.3 times the minimum wage in Neumark and Wascher (2008). Neumark, Schweitzer, and Wascher (2004) on the contrary find significant effects up to a rather high level (2 to 3 times the minimum wage). Evidence for the UK is more mixed. On the one hand Dickens and Manning (2004) believe that *spreading* effects were probably modest in the United Kingdom in 1999 when a national minimum wage was reintroduced, Stewart (2012b) reports evidence of no spillover effects and Stewart (2012a) of at most very limited ones. Dickens, Machin, and Manning (1999) on the other hand report effects of the minimum wage up to the 40th percentile and Butcher, Dickens, and Manning (2012) up to the 25th.

Among the few studies concerning other countries than the US and the UK, Teulings (2000, 2003) finds that increases in the minimum wage significantly spread to higher wages in the Netherlands.

These very different results may come from differences between countries, time periods or populations of interest, but they also remind us that, from an empirical point of view, the identification of these spillover effects is a complicated issue. The estimation of the distributional impact of an increase in the minimum wage faces indeed a double challenge.

The first empirical issue is to deal with the identification of the impact of the minimum wage. When the minimum wage is the same for all employees (as it is the case in France), it is generally impossible to distinguish what pertains to its specific increase or to any wage trend or other cyclical effect. In previous literature, identification is usually achieved by spatial or sectoral variations: for the US, Lee (1999) uses state variation in the minimum wage levels. Similarly, Dickens, Machin, and Manning (1999) use the coexistence of several sectoral minimum wages in the United Kingdom (until 1993). The endogeneity issue is extremely severe in France, as the legal increase in the minimum wage is yearly adjusted according to the past trend in mean wages (it is indexed to the *blue-collar worker's basic hourly pay*).[2]

For our estimation, we use a unique setting corresponding to a specific period of time during which the usual revaluation rule was frozen. This natural experiment is an unintended consequence of the gradual implementation of the French Law on workweek reduction between 1998 and 2003. During this period, monthly Guaranteed Wages (GMR) were designed to maintain the monthly wage of the lowest paid employees despite the lower number of hours worked at the time of the switch to the 35-hour week. Each year, during five years, a new GMR was created for the firms that signed an agreement that year. After five years, it had indeed resulted in the coexistence of six levels of the minimum wage. The new (right-wing) government elected in 2002 decided to put an end to this situation and designed a convergence mechanism that resulted in the application of different discretionary increases to the different GMR between 2003 and 2005. This allows us to identify the distributional impact of an increase in the minimum wage. The pace of increase in the level of the minimum wage applying in one or other firm can be considered as exogenous. The identification will rely on the fact that the increase was quite large in firms that did not sign a reduction agreement before 2002, while it was more modest in other firms.

The second empirical issue for the estimation of the distributional impact of the minimum wage relates to modeling the overall distribution of earnings, which requires to use specific tools. A first stream of literature uses parametric specifications (see for instance Teulings, 2000; Meyer and Wise, 1983), but these speci-

[2] For this reason, little evidence exists on this subject in France. The only (in French) contributions are from Koubi and Lhommeau (2007) and Goarant and Muller (2011). They both conclude that increases in the minimum wage have a significant impact up to wages as high as twice the minimum wage.

fications could be sensitive to the functional form assumed for the distribution of wages (see for instance Dickens, Machin, and Manning, 1998, on a related topic). Over the last decade, many empirical tools have been proposed for a more detailed analysis of the entire wage distribution. Quantile regressions is one of them. It classically deals with the quantiles of the distribution of the variable of interest Y conditional on observable characteristics X (see Koenker and Hallock, 2001). In the same way as linear regressions approximate the conditional expectation of the variable of interest as a linear function of observables, it models the conditional quantile of the variable of interest as a linear function of observables. As both methods deal with conditional quantities, they do not inform directly on the impact of a change in the distribution of observables on the expectation or quantile of our variable of interest on the whole population (meaning their *unconditional* counterpart). Firpo, Fortin, and Lemieux (2009) propose a more direct method to deal with unconditional quantities. It relies on the influence functions, and requires only a local inversion of the distribution. We use this method in this paper. This allows us to disentangle as much as possible the spillover effects from changes in the composition of the labor force resulting for instance from the exclusion of low-productivity workers. We focus on the impact of the minimum wage on the various deciles of the distribution of annual earnings of workers in the private sector, using administrative business data (the DADS) that provide exhaustive records on yearly earnings of French workers in the private sector. We perform separate analyzes for men and women to account for different wage settings according to gender.

Section 2 presents the revaluation mechanisms of the minimum wage with a specific focus on the *convergence* period of the different levels of the minimum wage. The data and identification strategy are detailed in section 3, then section 4 presents the statistical method used for analyzing the distribution of earnings. Section 5 provides the results.

2 French labor market institutional setting: Minimum wage and workweek reduction

The French minimum wage ("SMIC") was introduced in 1970. Its hourly value is set by the French government for all French employees of the private sector.[3] It amounts to 9 € per hour in 2011, which makes it one of the highest minimum wages among developed countries.[4] The minimum wage level is an important reference in

[3]Rare exemptions concern for instance the catering sector because of the existence of fringe benefits as meals.

[4]See for instance `http://stats.oecd.org/Index.aspx?DataSetCode=RMW`.

the French public debate. It is commonly used as a benchmark for living standard or earnings. Its update, that occurs every year according to a strict rule, is highly publicized and discussed. By law, the minimum wage increase cannot be smaller than the inflation rate observed the current year. It even exceeds it, as the annual increase in the purchasing power of the minimum wage corresponds at least to half the annual increase in the *purchasing power of blue-collar worker's basic hourly pay (SHBO hereafter).*[5] Besides, the French government can add to this strict rule an additional increase ("coup de pouce" or boost). The nominal rate of the minimum hourly wage can be written:

$$SMIC_t = SMIC_{t-1} \times \left(p_t/p_{t-1} + \frac{1}{2}\delta_{SHBO_t} + cdp_t \right)$$

where p_t is the price index in year t, δ_{SHBO_t} corresponds to the growth rate in the purchasing power of blue-collar workers' basic hourly pay and cdp_t represents the discretionary increase beyond the automatic revaluation rule (*boost*). For instance, the SMIC received a *boost* of 0.45 % on July 1, 1998, of 0.29 % on July 1, 2001 and of 0.30 % on July 1, 2006.

The gradual implementation of the new regulations on workweek reduction changed this situation. The 35-hour workweek was enacted in France by the so-called "Aubry Laws" (named after Martine Aubry, Minister of Labor), from a previous 39-hour workweek. All firms had to decrease the normal workweek time to 35 hours before January 1, 2000 (January 1, 2002 for the smallest ones) and to pay hours over 35 on an overtime basis. Incentives were provided to firms that negotiated an agreement before this binding limit, and the field implementation of the workweek reduction was thus gradual. Maintaining the hourly wage flat would have created a sharp drop in the monthly remuneration of workers. In order to avoid the loss of income for lowest-wage employees, the law imposed a new regulation for minimal wages. In firms that adopted the 35-hour workweek, a "monthly guaranteed wage" (hereafter GMR) was created. This GMR guaranteed that the monthly minimal remuneration would not be affected by the workweek reduction. In practice, it thus corresponded to a new legal hourly minimum wage for the firms that had signed a workweek reduction agreement. This was made possible by a generous cut in payroll taxes, in order to avoid a detrimental impact on employment.

However, if they guaranteed a maintained monthly remuneration at the time of the switch to the 35-hour week, the GMR did not then follow exactly the same

[5] A discussion of the consequences of this mechanism can be found in Cette and Wasmer (2010).

updating rules as the legal hourly minimum wage (which still applied to all firms that had not signed any workweek reduction agreement yet). While the former was updated according to changes in the blue-collar worker's basic *monthly* pay (SMBO), the latter followed the changes in the blue-collar worker's basic *hourly* pay (SHBO). This slight difference had unintended consequences. The SMBO evolved slower than the SHBO over the period because most workweek reduction agreements ensured the maintenance of a monthly salary despite the decrease in the number of hours, which mechanically translated into an increase in the hourly wage higher than that of the monthly wage. From one year to another the GMR thus benefited from lower updates than the hourly minimum wage which still applied to workers in the firms which had not signed yet any agreement. For firms negotiating workweek reductions later, the new minimum monthly wage, which would ensure no salary loss, was thus higher than the updated GMR of the previous year (see Figure 1 for the evolution of real GMR by date of the workweek reduction agreement and Table 1 for the precise creation calendar of the different GMR).

Because of the successive workweek reduction agreements, six levels of GMR coexisted in 2003. Beside the five GMR, the hourly minimum wage (which by abuse of language will be denoted by GMR 0) applied in firms which did not sign any workweek reduction agreement. At this date, the newly elected government put into place an adjustment mechanism in order to retrieve a unique level of minimum wage. From 2003 to 2005, the traditional revaluation rule of the minimum wage was frozen. While the highest hourly minimum wage rate (that applied to firms that had signed a workweek reduction agreement between July and December 2002) simply evolved as the inflation, the other hourly minimum wage levels received differential *boosts* so as to converge in 2005 to a unique hourly rate. The more they initially diverged the higher the *boosts* during this period (see Figure 1). Again, the impact on the labor cost was softened by substantial payroll tax exemptions.

To the best of our knowledge, the impact on the distribution of earnings of this convergence in the minimum wage levels has not been studied so far. Yet spillover effects due to an increase in the minimum wage are all the more credible in the French context. Besides, the exogenous increases that occurred during the convergence period provide a unique identification setting of these spillover effects.

3 Empirical Strategy

3.1 The data

We use the DADS panel (1/25th sample) over the period 2003-2005. This administrative business database starts in 1976 and provides yearly exhaustive data on workers in the private sector.[6] We have very accurate information on gross yearly earnings and number of hours worked during the year. We thus compute hourly wages which is our main variable of interest (expressed in euros of 2007). The data also provide detailed information on employees (age, seniority, gender, type of position...) and on firms (number of employees, industry, date of workweek reduction agreement). We restrict the sample to individuals aged 18 to 65.

For the minimum wage variable we use for year t the level of the minimum wage prevailing since July of year $t-1$, as this date corresponds to the annual update of the legal minimum wage rate over this period. This corresponds to the usual schedule of wage negotiations, that usually take place at the end of the year. As observed by Avouyi-Dovi, Fougère, and Gautier (2010), in France collective wage bargaining agreements usually take place at the end of the year and apply in the first months of the following year.[7] The empirical consequences of this choice are discussed in the Results section.

In the end, we have a panel of over 192,600 firms (among which about 98,700 were present all along between 2003 and 2005) and around 514,800 employees. As earnings distributions and wage negotiations could differ for men and women, we perform separate analyzes for both gender.

3.2 Identification Strategy

To identify the impact of the minimum wage on earnings, we rely on the specific convergence period of the different levels of the minimum wage. This peculiar situation creates a unique setting where we observe, during a short period of time, different legal minimum wage levels. Besides, these minimum wages exogenously increased at different paces for a three-year period. Our setting is close to a difference-in-differences strategy. The identification relies on the fact that we observe a steady increase in the minimum wage level in some firms (around 3.5% per year), while this increase was more modest in other firms.

[6] We limit the sample to wage earners and exclude self-employed persons.

[7] We can see an example of this in Koubi and Lhommeau (2007) and more recently in Goarant and Muller (2011): the analysis of quarterly effects shows a large peak of wage growth in the first quarter.

As shown in Table 2, the main part of our sample (around 70%) is constituted by employees who work in firms that did not sign a workweek reduction agreement (FGMR0) or firms that signed one quite early, between mid 1999 and mid 2000 (FGMR2). Very few firms signed workweek reduction agreements after mid 2002 or before mid 1999, as evidenced by the low number of employees belonging to FGMR 1 and FGMR 5 in our sample.

The signing date of a workweek reduction agreement is obviously not exogenous: in fact, the law was more restrictive for larger firms, which explains why they are overrepresented among those who contracted an agreement soon (see Table 3 in Appendix). Consequently, the manufacturing industry is also overrepresented in these groups. By contrast, those that had not signed an agreement in June 2002 (FGMR 0 and 5) are more often than average small and belonging to Trade or Services. We expect that these differences in the composition of the labor force of the GMR groups result in different observed wage distributions.

Besides, the negotiation date of workweek reduction agreements is probably related to the firm's wage policy and we cannot simply relate the minimum wage level with these observed distributions because of endogeneity issues.

However, the convergence period, which serves our identification purposes, was imposed a few years after for most of the firms and without having been anticipated. Therefore it seems plausible to consider this as a source of exogenous variation in the levels of the minimum wage.

As a first insight of the convergence period, we estimate a classical difference in differences regression of the log of the hourly wage on the log of the hourly minimum wage level. To capture systematic differences of wage policies in the different firms we use fixed effects for the different groups of GMR, and control for composition effects, that may change over time, by adding variables for characteristics of the employees and the firms (polynomial functions of age, socio-economic position, seniority, size, industry) and these variables interacted with year dummies. The specification can be written as:

$$log(y_{it}) = \alpha \underline{w}_{gt} + e_g + e_t + x'_{it}\gamma^t + \eta_{it} \quad (1)$$

where $\underline{w}_{gt}$ stands for the log of the hourly minimum wage level that prevails in FGMR group g to which wage earner i belongs, e_g stands for fixed effect for these groups, e_t for yearly trend and x_{it} for observed characteristics of the firms or the employees, whose effect γ^t can change over time. The implicit assumption is that in the absence of the increase in the minimum wage level, the wage would have evolved in the same way in all firms, whatever their GMR group. The interacted terms weaken this assumption, as we can assume for instance that the returns to education evolved differently. The dependent variable y_{it} is the hourly wage. As we observe yearly earnings in our data, this variable may change if the number of hours evolved because of the convergence period. To check that it is not the case, we also apply the same methodology using as a dependent variable the log of the number of hours worked.

Table 4 displays the results for the coefficient corresponding to the log of the hourly minimum wage level. We obtain that a 1 % increase in the minimum wage level results in a 0.2 % increase in the hourly wage. We do not observe any change in the number of hours worked because of the increase in the minimum wage level, which is reassuring for our identification strategy.

4 Modeling the distributional impact of an increase in the minimum wage

The change in the average hourly wage gives little information on the distributional impact of an increase in the minimum wage. Indeed, it is very unlikely that an increase in the minimum wage results in a translation shift of the earnings distribution. One expects on the contrary that it alters mostly the bottom of the distribution and that this effect differs depending on the position in the wage hierarchy. The empirical question is to determine how the shape of the distribution of earnings changes because of the increase in the minimum wage, and up to what level we observe an impact on the distribution. As illustrated in Figure 2, the cumulative distribution of log hourly real wages has clearly changed from 2003 to 2005. Moreover, a closer analysis suggests that these temporal shifts in the earnings distribution markedly changed depending on the increase in the minimum wage level in the GMR group. We observe that the sharp increases observed in the minimum wage level in FGMR0 (that yearly increased by 3.8 points of percentage over this period) results for this group in a shift of the distribution of wages that overcomes the sole very bottom of the distribution (see Figure 3). By contrast, the shift is much more reduced in FGMR2 where the increase is more modest (as it yearly increased by 1.1 points of percentage) and hardly noticeable in FGMR4. This empirical evidence suggests that the change in the shape of the cumulative distributions may be, at least partly, due to the increase in the minimum wage that was observed during this period. A more precise analysis requires to use tools that allows us to analyze this change in the distribution of earnings while also taking into account the potential effects due to changes in the composition of these groups.

In recent years, new methods to evaluate counterfactual distributions have emerged (a detailed presentation can be found in Fortin, Lemieux, and Firpo, 2011). The method, which is used here, is the so called *unconditional* quantile regression proposed by Firpo, Fortin, and Lemieux (2009). We rely on their estimator that allows for a direct measure of how a marginal change in the level of one variable (in our case, the minimum wage) will affect the distribution of wages in the population, keeping the distribution of other characteristics equal. More specifically, it provides a measure of the impact of a small location shift in the distribution of covariates X on some distributional statistic of a variable W, maintaining the conditional distribution of W given X unaffected. They call this notion "unconditional partial effect."

As shown by Firpo, Fortin, and Lemieux (2009), the vector α of partial derivatives representing the change in the distributional statistic q_τ of W with respect to a small location shift in the distribution of the covariates X is such that:

$$\alpha(q_\tau) = \int \frac{dE(RIF(W, q_\tau)|X = x)}{dx} dF_X(x) \tag{2}$$

Where RIF stands for the recentered influence function, which is notably simple in the case of a τth order quantile, as:

$$RIF(w_i, q_\tau) = \underbrace{q_\tau + (1-\tau)\frac{1}{f_W(q_\tau)}}_{c_{2,\tau}} + \underbrace{\frac{1}{f_W(q_\tau)}}_{c_{1,\tau}} 1(w_i > q_\tau). \tag{3}$$

We note $c_{1,\tau} = F_w^{'-1}(\tau) = 1/f_W(q_\tau)$ and $c_{2,\tau} = q_\tau + (1-\tau)c_{1,\tau}$. This expression corresponds to the sum of a constant $c_{2,\tau}$ and the probability $P(W > q_\tau|X = x)$ deflated by the density $f_W(q_\tau)$ of W evaluated at q_τ (corresponding to $c_{1,\tau}$). Both q_τ and $F_w^{'-1}(\tau)$ are constant and independent of X and can be easily estimated.[8] Table 5 provides the values of the deciles dec_j, $j \in [1, 9]$ and of the inverse of the density in each of these deciles (c_{1,dec_j}) in our sample.

As shown by Firpo, Fortin, and Lemieux (2009), a consistent estimator can be obtained once the dependance of $P(Y > q_\tau|X = x)$ in x is specified. We indeed have for each decile

$$E(RIF(W_i; dec_j)|X = x) = c_{1,dec_j} P(w_i > dec_j|X = x) + c_{2,dec_j} \tag{4}$$

We rely for our main specification on a RIF-OLS, meaning a linear specification for this conditional probability $P(W > q_\tau|X = x) = x'\beta$. As already noted by

[8] The density $f_W(q_\tau)$ can be estimated directly by kernel methods but it is computationally intensive. Following Koenker (2005), we rely on the approximation $F_w^{'-1}(\tau) = \frac{F^{-1}(\tau+h)-F^{-1}(\tau-h)}{2h}$. The optimal window verifies (under certain conditions): $h_n = n^{-1/5}\left(\frac{4.5\varphi^4(\Phi^{-1}(t))}{(2\Phi^{-1}(t)^2+1)^2}\right)^{1/5}$ where φ and Φ^{-1} respectively represent the probability density function (pdf) and the inverse of the cumulative distribution function (cdf) of the normal distribution and n is the sample size.

Firpo, Fortin, and Lemieux (2009), the results are quite robust to the specification choice but the linear setting has the advantage of providing a simpler framework which makes interpretations more straightforward (we check that our main conclusions are unchanged when using a Logit specification for this conditional probability). As made clear by Equation 2, when using a RIF-OLS specification the impact on the marginal quantile $\nu(F_W)$ of a marginal shift in X_k corresponds simply to

$$\hat{\alpha}_\tau^k = \frac{\hat{\beta}_\tau^k}{f(q_\tau)}$$

The intuition for this is illustrated in Figure 4. The issue is to determine how the quantile of the distribution of a variable of interest y changes in response to a marginal shift in the distribution of a covariate X: namely, that this variable becomes $X + \delta x$, letting the distribution of Y conditional on X unchanged. Assuming that this relation is locally linear, the small change in the distribution of X induces that $F(q_\tau)$ changes τ into $\tau - \beta\delta x$, and the τth order quantile in the new distribution is $q_\tau + \delta q$. As the slope of the distribution at this point is $f(q_\tau)$, the impact of the change in the quantile is thus $\delta q = \beta\delta x / f(q_\tau)$.

In practice, in our main specification we estimate for each decile the corresponding RIF as defined by (3) and regress the corresponding values on the complete set of covariates as in (1), namely the log minimum wage level, GMR group dummies, year dummies, observed characteristics for firms and employees and their interactions with year dummies. The estimation is thus

$$RIF(w_i; dec_j) = \alpha_\tau \underline{w}_{gt} + e_g + e_t + x_{it}'\gamma_\tau^t + \eta_{it} \tag{5}$$

Pointwise confidence intervals are obtained by bootstrap. A grouped data analysis is also provided as a robustness check in the Results Section which leads to more conservative confidence intervals.

5 Results

5.1 Main specification

Figure 5 reports, for each decile q_τ, the estimated effect of a marginal increase in the minimum wage. The estimates correspond to the $\hat{\alpha}_\tau^k = \hat{\beta}_\tau^k / f(q_\tau)$ presented above. According to our results, the increase in the minimum wage that occurred during the convergence period had a positive impact on the log-hourly wage distribution. The measured effect is decreasing with the wage level. Its magnitude is around 0.2 for the median hourly wage, and it is still significant up to the seventh decile for men. For the sake of comparison, note that the first decile (respectively seventh decile) corresponds roughly to 1.1 times (resp. 2.3 times) the average minimum wage at this time. These results are in line with the ones of Neumark, Schweitzer, and Wascher (2004) and Dickens, Machin, and Manning (1999). They are quite higher than most of the other results for the US and the UK. Detailed results for other covariates are presented in Table 6 (for the sake of simplicity, we present only the estimates for men).

Note that although the first decile is very close to the minimum wage, the effect of a one percent increase in the minimum wage level in July of year $t-1$ does not have a one percent effect on the first decile of the distribution of the log hourly wages over year t (see Figure 5) as could be expected from a mechanical increase. The measured effect is rather of about 0.5 %. This is in fact a natural result given that the available data is collected on a yearly basis and that during the studied period the minimum wage was revaluated each year on 1st July. Consequently, for workers whose earnings correspond to the minimum wage level, the level of minimum wage we use as a covariate for the estimation was the actual legal reference only for the first half of the year in question. Noting $\underline{w}_t$ the minimum wage prevailing on 1st July of year t, and $\tilde{w}_t$ the average hourly wage during year t, $\tilde{w}_t = (\underline{w}_{t-1} + \underline{w}_t)/2$ and a 1 % exogenous increase in $\underline{w}_{t-1}$ leads to a 0.5 % increase in $\tilde{w}_t$. For workers who are not mechanically affected by changes in the minimum wage, potential revaluations mostly occur at the beginning of the next year.

Estimations are also performed separately for the two-year periods 2003-2004 and 2004-2005 (see Figure 6). Although the confidence intervals are wider due to the smaller sample sizes we obtain patterns that are close to our main specification. This temporal decomposition emphasizes however that the spillover effects seem to increase over time. During the first year, they were mostly located at the bottom of the distribution, and they spread over the whole distribution the year after.

5.2 Robustness checks

In this section we provide two robustness checks: a placebo test and a replication of the previous analysis using a grouped data approach.

First, to check that our results are not due to specific wage dynamics in the different groups of firms, we provide a "placebo" test. More specifically, we use the two-year period available in our data *after* the convergence period (2006 and 2007), and simulate, for each GMR, fake increases in minimum wage levels of the same magnitude as the ones observed during the convergence period. Using the same estimates as before, we observe a negligible impact of the *not-happened* increase in minimum wage level (see Figure 7). These results are not due to the fact that using two years instead of three mechanically reduces the sample size. Indeed, as discussed above, when performing the restricted analysis to the two-year periods 2003-2004 and 2004-2005, there is still enough statistical power to observe significant impacts of the minimum wage quite high in the distribution.

Second, note that the identification of the impact of the minimum wage is mainly due to variation at the group$\times$year level. As emphasized by Moulton (1990) or more recently by Bertrand, Duflo, and Mullainathan (2004) in a differences-in-differences context, this clustered structure of the data may lead to a biased inference. If the unobserved components of individual earnings are correlated within clusters, ignoring this correlation may seriously bias upward the estimated precision. Indeed, while we have a very large sample, the estimated precision under the iid assumption is very high. We thus perform, as a robustness check, estimations using aggregated data at the group$\times$year level following Wooldridge (2003). More specifically, for each decile, we estimate in each FGMR$\times$year cluster the proportion of workers whose earnings are higher than the decile, correcting for composition effects due to variations in individual covariates. We regress these 6$\times$3 proportions $\tilde{\pi}^j_{gt}$ on the minimum wage level, controlling for year and group effects (see Appendix for further details). This leads to a dramatic decrease in the sample size but the results appear surprisingly robust to these conservative tests. The obtained coefficients in this 18-observation sample are significant up to a high level. We provide a graphical illustration of these estimations in Figures 8 and 9. Each plot corresponds to a given decile. For each of them, the empirical counterparts $\tilde{\pi}^j_{gt} - \hat{e}_t - \hat{e}_g$ of $P(w_i > dec_j|g,t)$, purged from composition effects and estimated group and year effects, are plotted against their corresponding level of the minimum wage. The variables are transformed such that the slopes can be directly interpreted as the effect on each decile of a marginal increase in the minimum wage. As with the main specification, the impact of the minimum wage decreases with the deciles. For men, the slope coefficient is not significant for the 5th and 6th deciles but it becomes significant again for the 7th decile. The impacts on the 8th and 9th deciles are much smaller and not significant.

This procedure can be seen as an overidentified minimum distance estimation whose overidentifying restrictions can be easily tested (see Appendix). Interestingly we systematically fail to reject the null hypothesis which confirms the chosen approach.

6 Conclusion

We propose an empirical evaluation of the impact on the earnings distribution of French workers of the repeated and exogenous increases in the minimum wage that occurred between 2003 and 2005. Our identification strategy relies on this specific period when some firms were compelled to apply a steady increase in the minimum wage level, while the increase was much smaller in other ones. To describe the effect on the whole distribution of earnings, we use a model of unconditional quantile regression. We provide evidence that this increase has spread over a large part of the wage distribution. As we use a complete administrative dataset, and the "treatment" (*i.e.* the minimum wage increases) was unusually high, the statistical power is very strong. It is thus more likely to observe significant effects far to the right of the distribution. During the considered period, a 1% increase in the minimum wage level increases by around 0.5% the first decile of earnings. The impact is decreasing over the distribution, but is as high as 0.2% at the level of the median of the distribution of earnings. If the impact is not significant at a higher level of the distribution of earnings of female wage earners, it is still significant for the seventh decile of the one of male wage earners.

Compared to results obtained in other countries and with other methods, our estimates belong to the upper part in terms of magnitude. They are for instance in line with the ones of Neumark, Schweitzer, and Wascher (2004) and Dickens, Machin, and Manning (1999). A few specific features of the French labor market may explain our quite strong results. First, the "bite" of the minimum wage is particularly high in France. According to the OECD, the French minimum wage represents as much as 63% of the median wage (Kaitz index) and 50% of the average wage of full-time French workers, that is the largest one of all OECD countries. Besides, it is commonplace that collective agreements specifically refer to the minimum wage in France. Indeed, our results are consistent with theoretical models on incentives and recent experimental contributions that highlight the role of the minimum wage in wage bargaining (see Falk, Fehr, and Zehnder, 2006). Commitment at work also has complex links with the situation in the wage ladder (see for instance Card, Mas, Moretti, and Saez, 2010).

Finally, our results are obtained in a specific context, when the minimum wage increased at an unusually high pace in a few (rather small) firms. One can thus question their external validity. Be that as it may, they provide new empirical elements on wage setting. Therefore these elements have to be taken into account when analyzing the impact of the minimum wage both on employment and earnings inequality. These effects must also be taken into account in terms of global labor cost and firm competitiveness because they show that the impact of the minimum wage goes way beyond the lower paid individuals.

Appendix A Data description

We use the DADS panel (1/25th sample) over the period 2003-2005. This administrative business database starts in 1976 and provides yearly exhaustive data on workers in the private sector. The data contain very accurate information on gross yearly earnings and number of hours worked during the year but also on employees (age, seniority, gender, type of position...) and on the firm (number of employees, industry, date of workweek reduction agreement).

We restrict the analysis to the private sector and exclude a few formerly state-owned firms whose legal status changed during the period studied here. Namely we drop France Telecom, EDF and GDF. We also exclude self-employed persons and restrict the sample to individuals aged 18 to 65. In some cases, the gross annual remuneration of certain employees falls below the annual minimum wage. Several possibilities can explain that. First, the minimum wage regulations do not apply to all occupations (*e.g.* in the case of employees whose working time is difficult to measure, such as some traveling salesmen). Moreover, the strict definition of the minimum wage also includes some fringe benefits that are not always valued in the DADS. It can also be due to reporting problems, for instance unearned bonuses from one year to another, or problems in the number of days worked during the year. We thus choose to exclude workers whose hourly wage is lower than the minimum wage.

Appendix B RIF-OLS estimation on clustered data

We follow here the procedure described in Wooldridge (2003). From Equation (3) we define the average RIF at a GMR and year level as:

$$E(RIF(w_i, dec_j)|e_g, e_t) = c_{1,dec_j} P(w_i > dec_j|e_t, e_g) + c_{2,dec_j} \tag{6}$$

Without additional covariates and using the RIF-OLS specification, $P(w_i > dec_j|e_t, e_g)$ can be approximated by the corresponding proportion $\hat{\pi}^j_{ft}$ in each year and GMR group. We can simply regress this average on year and GMR group dummies as well as the minimum wage level that apply in this GMR group during this year. With weights equal to the number of individuals per group, this regression gives the same estimates as the one on the full population. However, this estimate can be biased by composition effects. In practice we thus use the following two step procedure: in the first step $1_{x_i>dec_j}$ is regressed on all the additional covariates x plus FGMR×year dummies.

$$1_{y_i>dec_j} = \tilde{\pi}_{gt} + x'\xi + \tilde{\varepsilon}^j_i$$

This way we obtain group fixed effects $\hat{\hat{\pi}}^j_{gt}$ which are purged of the variations in the additional covariates. In a second step, these $\hat{\hat{\pi}}^j_{gt}$ replace the $P(w_i > dec_j | e_t, e_g)$ from Equation (6) and are regressed on year dummies, FGMR dummies and $\underline{w}_{gt}$.

$$\hat{\hat{\pi}}_{gt} = e^j_t + e^j_g + \beta^j \underline{w}_{gt} + \upsilon^j_{gt} \tag{7}$$

Note that with weights equal to the number of individuals per group, this procedure gives the same estimates as the ones on the full population.

An efficient version of this estimator can be computed with weighted least squares using as weights $1/SE(\hat{\hat{\pi}}_{gt})^2$ where the asymptotic variances of $\hat{\hat{\pi}}_{gt}$ are made fully robust to heteroscedasticity.

This procedure corresponds to an overidentified minimum distance estimation of β^j where the number of overidentifying restrictions corresponds to the number of degrees of freedom in the above equation.

Testing these overidentifying restrictions can be easily done by computing the SSR which follow a χ^2_{df} under the null hypothesis that $P(w_i > dec_j)$ can be approximated as in Equation (7). Failing to reject this test would confirm the chosen approach and would tell that critiques such as the one from Donald and Lang (2007) do not apply here.

Tables

Table 1: Schedule for the creation of the different monthly guaranteed wages (GMR)

GMR	Date of the workweek reduction agreement
GMR1	Between June 15, 1998 and June 30, 1999
GMR2	Between July 1, 1999 and June 30, 2000
GMR3	Between July 1, 2000 and June 30, 2001
GMR4	Between July 1, 2001 and June 30, 2002
GMR5	After July 1, 2002

Table 2: Number of firms and employees for each GMR group in 2003

	FGMR0	FGMR1	FGMR2	FGMR3	FGMR4	FGMR5	Total
Firms							
Number	90,013	3,344	17,492	11,340	18,737	2,073	142,999
Share (%)	63.0	2.3	12.2	7.9	13.1	1.5	100
Employees							
Number	147,425	25,015	122,285	58,117	32,416	4,179	389,437
Share (%)	37.9	6.4	31.4	14.9	8.3	1.1	100

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.

Table 3: Characteristics of firms in each GMR group (2003 to 2005)

	FGMR0	FGMR1	FGMR2	FGMR3	FGMR4	FGMR5	Total
Number of employees (in %)							
less than 10	46.2	9.7	6.1	13.9	40.8	45.3	37.2
10 to 49	43.8	39.0	40.5	44.7	49.9	40.7	44.1
50 to 499	9.6	45.0	46.9	37.3	8.7	12.9	17.1
500 to 4,999	0.4	6.0	6.2	3.9	0.6	1.1	1.5
over 5,000	0.0	0.2	0.3	0.2	0.0	0.0	0.1
Gender							
Share of women	32.0	35.8	35.4	34.9	33.2	33.9	33.9
Industry							
Agriculture	2.0	1.3	1.1	0.5	1.3	1.9	1.7
Manufacturing industry	18.8	35.6	33.3	33.2	22.9	20.0	22.7
Construction	15.0	9.0	7.9	7.9	13.4	12.1	13.2
Trade	25.3	21.4	21.5	28.3	24.8	0.0	24.9
Services	38.9	32.8	36.2	30.0	37.6	39.3	37.6

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.

Table 4: Difference in differences coefficients for different variables of interest

	DiD coefficient	
	Men	Women
Log of the hourly wage	0.27^{***} (0.03)	0.20^{***} (-0.02)
Number of annual hours worked	-44.2 (34.6)	57.9 (47.6)

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.
*Note: * corresponds to a 10% significance level, ** to a 5% level and *** to a 1% level.*

Table 5: Deciles q_τ of the hourly earnings (in level) and coefficients $c_{1,\tau}$ (log)

	Men		Women	
τ	q_τ	$c_{1,\tau}$	q_τ	$c_{1,\tau}$
0.1	9.68	1.12	9.02	0.83
0.2	10.71	0.94	9.78	0.79
0.3	11.73	0.91	10.57	0.80
0.4	12.87	0.97	11.51	0.89
0.5	14.25	1.10	12.64	1.00
0.6	16.05	1.32	14.06	1.12
0.7	18.64	1.72	15.88	1.36
0.8	22.92	2.54	18.57	1.91
0.9	30.55	3.51	23.90	3.33

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.
Note: For the sake of readability values of q_τ correspond to the deciles of the distribution of hourly earnings ($q_\tau = F_w^{-1}(\tau)$), while the coefficients $c_{1,\tau} = F'^{-1}_{log(w_i)}(\tau)$ correspond to the distribution of log-earnings.

Table 6: RIF-OLS estimation for deciles of log hourly wage (Men, 2003-2005)

	1st decile	2nd decile	3rd decile	4th decile
Minimum wage (log)	0.57 *** (0.033)	0.45 *** (0.034)	0.39 *** (0.036)	0.33 *** (0.039)
Year Dummy				
2002	*ref.*	*ref.*	*ref.*	*ref.*
2003	-0.053^{***} (0.017)	-0.0037 (0.015)	0.024 (0.015)	0.0019 (0.015)
2004	-0.014 (0.017)	0.019 (0.016)	0.051 *** (0.014)	0.057 *** (0.014)
GMR group				
GMR 0	*ref.*	*ref.*	*ref.*	*ref.*
GMR 1	0.066 *** (0.002)	0.069 *** (0.0021)	0.061 *** (0.0023)	0.047 *** (0.0024)
GMR 2	0.019 *** (0.0017)	0.027 *** (0.0018)	0.03 *** (0.002)	0.029 *** (0.0022)
GMR 3	0.013 *** (0.0022)	0.024 *** (0.0023)	0.026 *** (0.0025)	0.024 *** (0.0028)
GMR 4	-0.014^{***} (0.0025)	-0.017^{***} (0.0028)	-0.022^{***} (0.003)	-0.029^{***} (0.0032)
GMR 5	-0.0072 (0.0042)	-0.0017 (0.0043)	-0.0083 (0.0047)	-0.014^{***} (0.005)
Socio-economic position				
CEOs	0.0013 *** (0.00051)	0.001 (0.00056)	0.00088 (0.00057)	0.0016 (0.00062)
Executives	0.0012 (0.0005)	0.0031 *** (0.00054)	0.0032 *** (0.00058)	0.0042 *** (0.00062)
Technicians and associate professionals	$-7.6e-05$ $(4.2e{-}05)$	$-6.7e-05$ $(4.8e{-}05)$	$-5.1e-05$ $(5.2e{-}05)$	$-9.8e-05$ $(5.7e{-}05)$
Office clerks and service workers	$-6.7e-05$ $(4.1e{-}05)$	-0.00024^{***} $(4.6e{-}05)$	-0.00023^{***} $(5e{-}05)$	-0.00029^{***} $(5.6e{-}05)$
Skilled and unskilled workers	*ref.*	*ref.*	*ref.*	*ref.*
Seniority				
Seniority	0.0091 *** (0.00034)	0.0095 *** (0.00039)	0.0082 *** (0.0004)	0.006 *** (0.00043)
Seniority^2	-0.00049^{***} $(2.8e{-}05)$	-0.00038^{***} $(3.3e{-}05)$	-0.00022^{***} $(3.5e{-}05)$	$7.2e-06$ $(4e{-}05)$
Seniority^3	$7.7e-06^{***}$ $(5.9e{-}07)$	$4.8e-06^{***}$ $(7e{-}07)$	$1.3e-06$ $(7.7e{-}07)$	$-3.2e-06^{***}$ $(8.9e{-}07)$
Age				
Age	$-1.3e-11$ $(6.1e{-}12)$	$-1.6e-11$ $(6.8e{-}12)$	$-5.6e-12$ $(7.3e{-}12)$	$-1.1e-11$ $(8.7e{-}12)$
Age^2	$-2.2e-11^{***}$ $(6e{-}12)$	$-2.2e-11^{***}$ $(6.4e{-}12)$	$-1.7e-11$ $(7.4e{-}12)$	$-3e-11^{***}$ $(8.8e{-}12)$
Industry				
Agriculture	$1.5e-06$ $(9.2e{-}07)$	$1.4e-06$ $(1e{-}06)$	$1e-06$ $(1.2e{-}06)$	$2.1e-06$ $(1.3e{-}06)$
Industry	$1.2e-06$ $(8.7e{-}07)$	$4.7e-06^{***}$ $(9.7e{-}07)$	$4.3e-06^{***}$ $(1.1e{-}06)$	$5.7e-06^{***}$ $(1.2e{-}06)$
Construction	$8.8e-07^{***}$ $(2.9e{-}07)$	$6.7e-07$ $(3e{-}07)$	$5.5e-07$ $(3.2e{-}07)$	$1.1e-06^{***}$ $(3.7e{-}07)$
Trade	$4.8e-07$ $(2.9e{-}07)$	$5e-07$ $(3.1e{-}07)$	$-2.7e-09$ $(3.2e{-}07)$	$4.1e-07$ $(3.8e{-}07)$
Services	*ref.*	*ref.*	*ref.*	*ref.*
Size of the firm				
Size	$2.1e-06^{***}$ $(2.2e{-}07)$	$2.4e-06^{***}$ $(2.2e{-}07)$	$2.9e-06^{***}$ $(2.4e{-}07)$	$3.5e-06^{***}$ $(2.7e{-}07)$
Size^2	$-1.5e-11^{***}$ $(4.7e{-}12)$	$-1.4e-11^{***}$ $(5.1e{-}12)$	$-1.8e-11^{***}$ $(5.5e{-}12)$	$-2.2e-11^{***}$ $(6.6e{-}12)$
Size^3	$8.5e-18$ $(2.4e{-}17)$	$-1.3e-17$ $(2.6e{-}17)$	$2.7e-18$ $(2.9e{-}17)$	$1e-17$ $(3.6e{-}17)$

Source: DADS panel, 1/25th sample.
Field: Male employees from the private sector aged 18 to 65, excluding interns and apprentices. Covariates also include interactions of year dummies and observable characteristics.
Note: * *corresponds to a 10% significance level,* ** *to a 5% level and* *** *to a 1% level. The minimum wage variable was divided by* $c_{1,\tau}$ *so that the obtained coefficient directly corresponds to the effect of the minimum wage on the deciles.*

Table 7: RIF-OLS estimation on deciles of log hourly wage (Men, 2003-2005)

	5th decile	6th decile	7th decile	8th decile	9th decile
Minimum wage (log)	0.25*** (0.043)	0.19*** (0.048)	0.14*** (0.055)	0.13 (0.071)	−0.044 (0.088)
Year Dummy					
2002	*ref.*	*ref.*	*ref.*	*ref.*	
2003	−0.0049 (0.015)	−0.0071 (0.017)	−0.01 (0.02)	−0.055 (0.024)	−0.14*** (0.031)
2004	0.044*** (0.016)	0.064*** (0.017)	0.084*** (0.019)	0.048 (0.025)	0.027 (0.033)
GMR group					
GMR 0	*ref.*	*ref.*	*ref.*	*ref.*	
GMR 1	0.03*** (0.0026)	0.014*** (0.0028)	0.002 (0.0032)	0.006 (0.0041)	0.0032 (0.005)
GMR 2	0.024*** (0.0023)	0.018*** (0.0027)	0.0087*** (0.0032)	0.0091 (0.004)	0.017*** (0.0048)
GMR 3	0.019*** (0.003)	0.014*** (0.0034)	0.0071 (0.0039)	0.01 (0.005)	0.013 (0.0061)
GMR 4	−0.037*** (0.0033)	−0.037*** (0.0036)	−0.04*** (0.0043)	−0.043*** (0.0057)	−0.038*** (0.0068)
GMR 5	−0.021*** (0.0054)	−0.025*** (0.006)	−0.033*** (0.007)	−0.037*** (0.0084)	−0.032*** (0.011)
Socio-economic position					
CEO	0.0017 (0.00069)	0.002 (0.00081)	0.0016 (0.00095)	−0.0024 (0.0013)	−0.0041 (0.0017)
Executives	0.0042*** (0.00068)	0.004*** (0.00078)	0.0034*** (0.001)	0.00094 (0.0013)	−0.0018 (0.0016)
Technicians and associate professionals	−9.7e − 05 (6.5e−05)	−0.00012 (7.5e−05)	−0.00013 (9.1e−05)	0.00018 (0.00013)	0.00024 (0.00017)
Office clerks and service workers	−0.00028*** (6.4e−05)	−0.00024*** (7.4e−05)	−0.00024 (9.7e−05)	−8.1e − 05 (0.00012)	0.0001 (0.00016)
Skilled and unskilled workers	*ref.*	*ref.*	*ref.*	*ref.*	*ref.*
Seniority					
Seniority	0.0048*** (0.00049)	0.0042*** (0.00055)	0.0057*** (0.00067)	0.0088*** (0.00086)	0.011*** (0.0012)
Seniority^2	9.4e − 05 (4.6e−05)	7.1e − 05 (5.2e−05)	−0.00017*** (6.4e−05)	−0.0007*** (8.5e−05)	−0.001*** (0.00011)
Seniority^3	−4.5e − 06*** (1e−06)	−3.6e − 06*** (1.2e−06)	1.7e − 06 (1.5e−06)	1.4e − 05*** (2e−06)	2e − 05*** (2.8e−06)
Age					
Age	−1.9e − 11 (9.7e−12)	−1.9e − 11 (1.2e−11)	−2.3e − 11 (1.3e−11)	1.5e − 11 (1.7e−11)	−3.7e − 11 (2e−11)
Age^2	−2.9e − 11*** (1e−11)	−4.6e − 11*** (1.2e−11)	−5.8e − 11*** (1.3e−11)	1.3e − 11 (1.6e−11)	−4.2e − 11 (2.1e−11)
Industry					
Agriculture	1.9e − 06 (1.5e−06)	2.8e − 06 (1.7e−06)	3.8e − 06 (2.1e−06)	−2.6e − 06 (3.2e−06)	−3.3e − 06 (4.2e−06)
Industry	5.9e − 06*** (1.5e−06)	5e − 06*** (1.7e−06)	6e − 06*** (2.3e−06)	3.4e − 06 (3e−06)	−6.9e − 07 (4e−06)
Construction	1.2e − 06*** (4.2e−07)	1.8e − 06*** (5e−07)	1.7e − 06*** (5.7e−07)	−1.8e − 06 (7.4e−07)	1.1e − 06 (9.7e−07)
Trade	8.6e − 07 (4.1e−07)	8.8e − 07 (5e−07)	9.8e − 07 (5.9e−07)	−1.2e − 06 (7.7e−07)	1.1e − 06 (9e−07)
Services	*ref.*	*ref.*	*ref.*	*ref.*	*ref.*
Size of the firm					
Size	4.3e − 06*** (3.1e−07)	3.9e − 06*** (3.7e−07)	5.2e − 06*** (4.2e−07)	5.6e − 06*** (5.5e−07)	2.2e − 06*** (6.8e−07)
Size^2	−3.6e − 11*** (7.6e−12)	−2.4e − 11*** (9e−12)	−5.6e − 11*** (1e−11)	−7.5e − 11*** (1.3e−11)	−1.1e − 11 (1.6e−11)
Size^3	7e − 17 (4.2e−17)	−4.9e − 17 (4.9e−17)	6.6e − 17 (5.5e−17)	1.5e − 16 (6.8e−17)	−1.7e − 16 (8.4e−17)

Source: DADS panel, 1/25th sample.
Field: Male employees from the private sector aged 18 to 65, excluding interns and apprentices. Covariates also include interactions of year dummies and observable characteristics.
*Note: * corresponds to a 10% significance level, ** to a 5% level and *** to a 1% level. The coefficients correspond to the effect of the covariates on the deciles.*

Figures

Figure 1: Level and evolution of hourly GMRs (2002-2005)

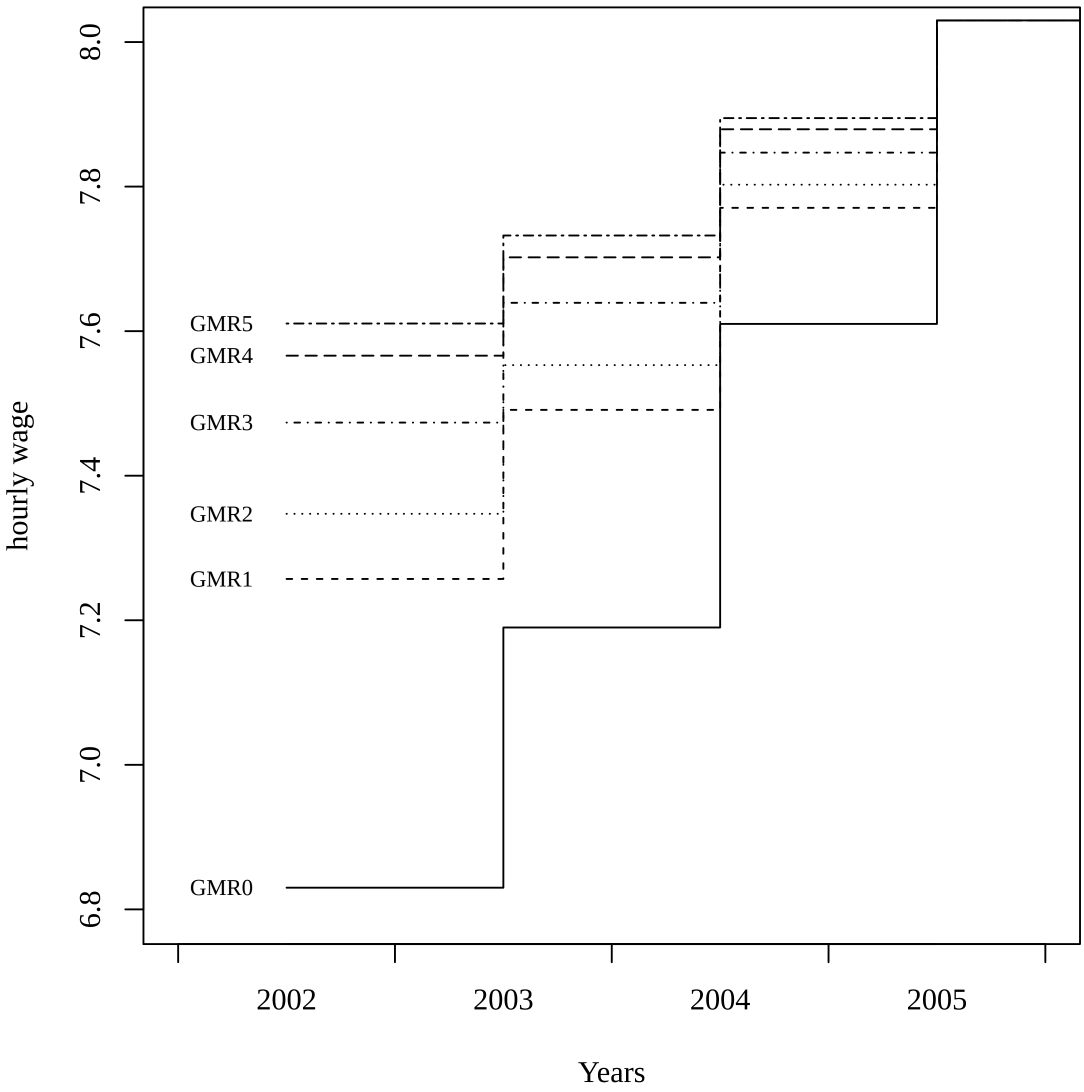
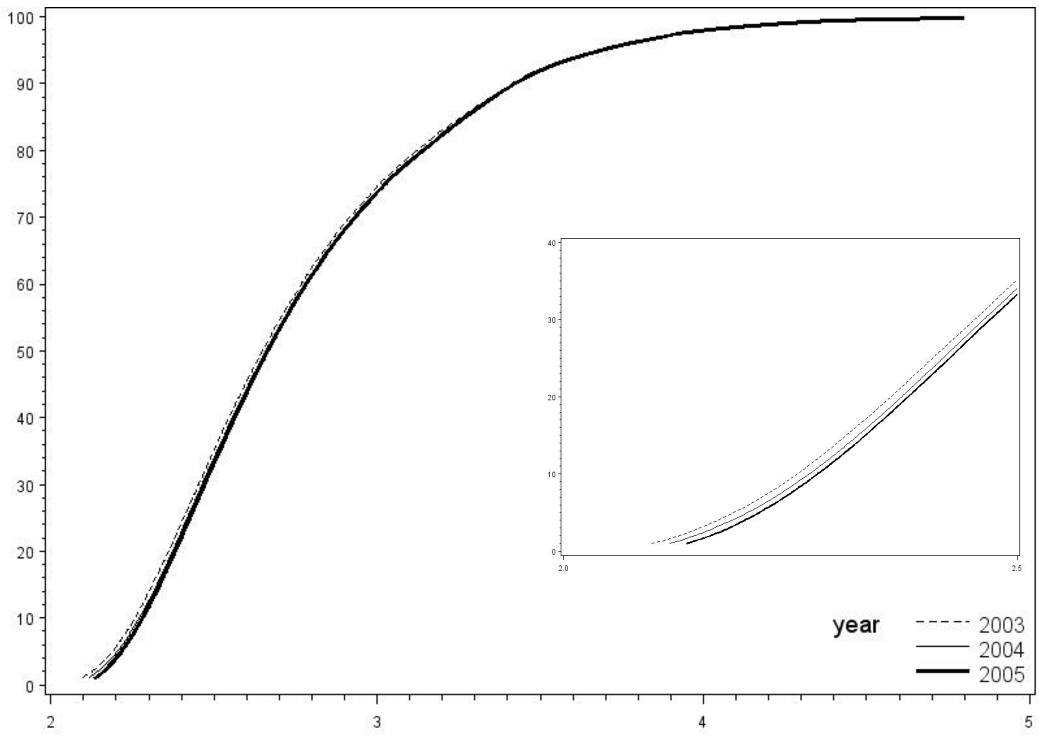
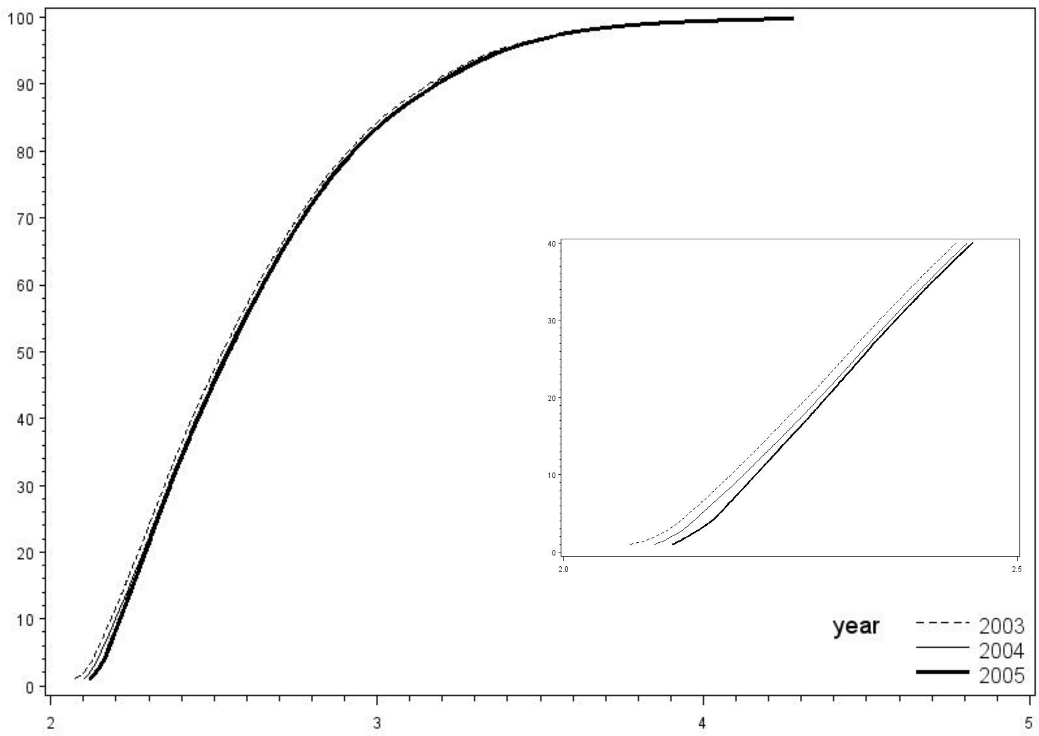


Figure 2: Distributions of log hourly real earnings of male and female employees in the private sector, from 2003 to 2005

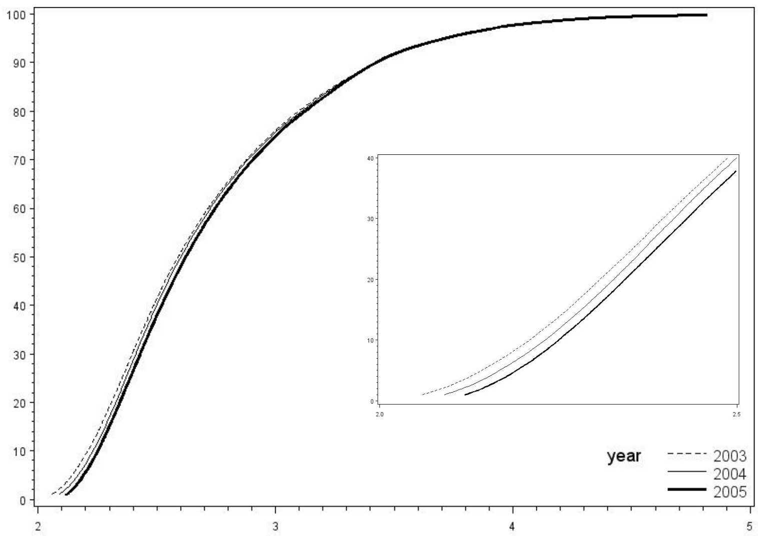


(men)

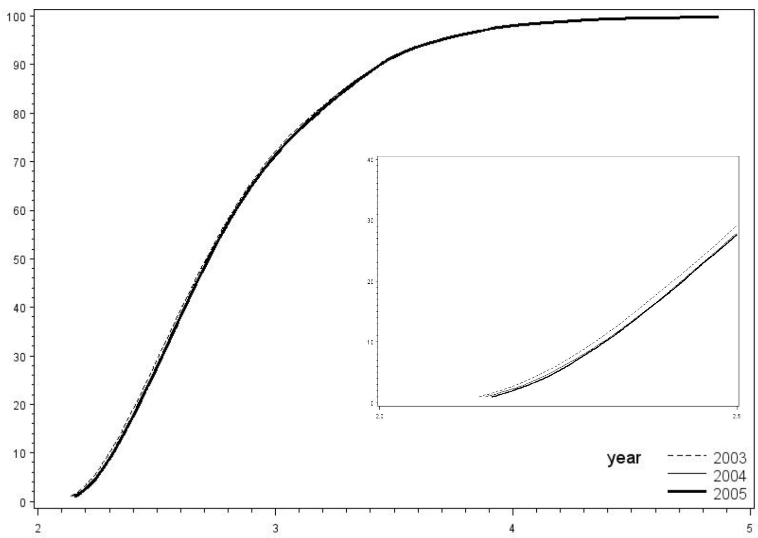


(women)

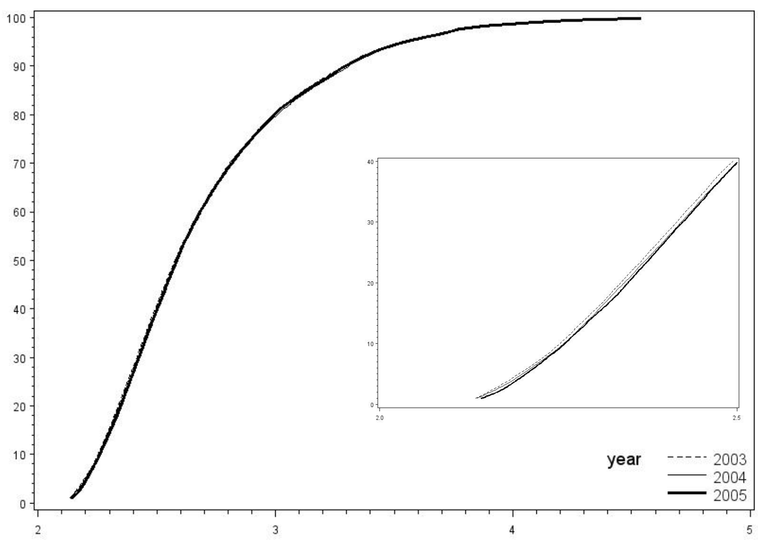
Figure 3: Distributions of log hourly real earnings of male and female employees in the private sector in GMR0, GMR2 and GMR4 from 2003 to 2005



(GMR0)

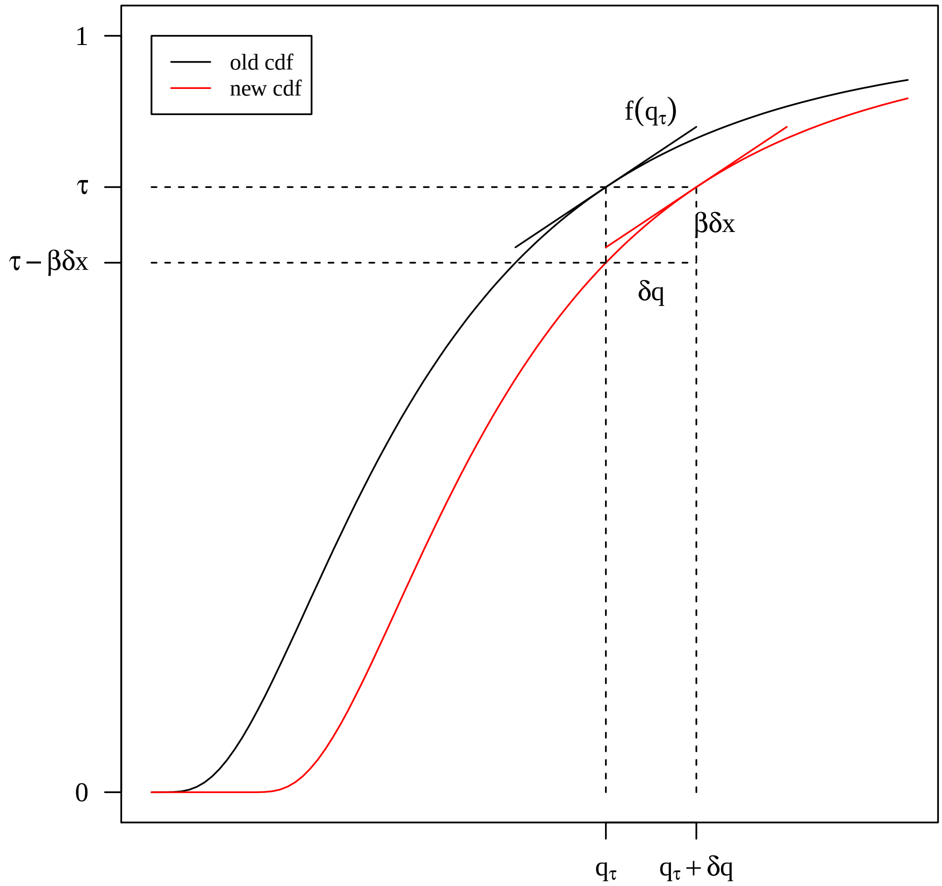


(GMR2)



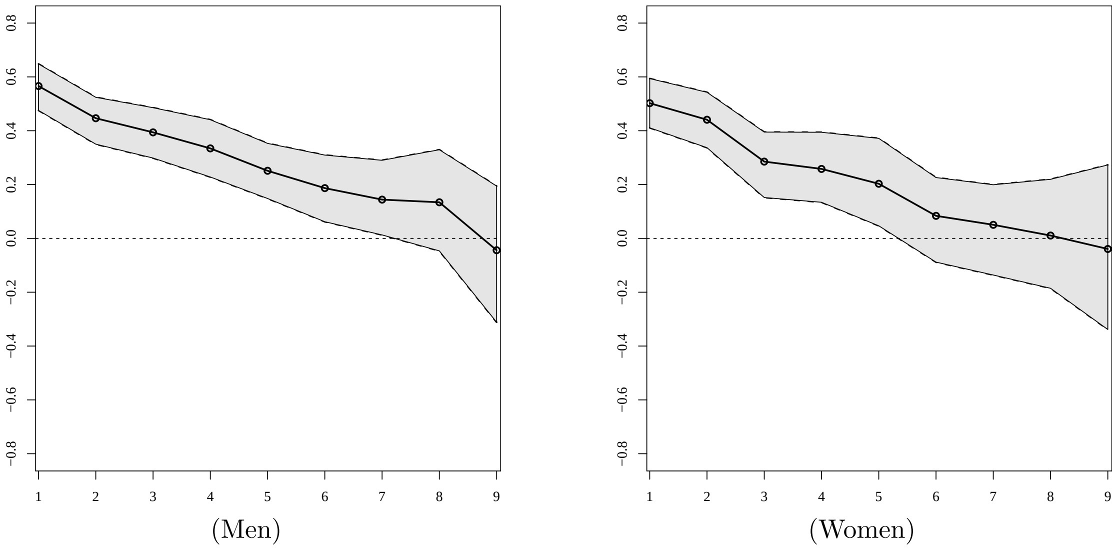
(GMR4)

Figure 4: Illustration of the UQR method for quantiles



Note: a small change in a covariate X induces a change of $P(y_i > q_\tau)$ to $P(y_i > q_\tau) + \beta\delta x$ which in turn changes τ into $\tau - \beta\delta x$ and induces a change in the τth order quantile from q_τ to $q_\tau + \delta q$, such that $\delta q = \beta\delta x / f(q_\tau)$.

Figure 5: Impact of the minimum wage on the different deciles of log hourly wage 2003-2005

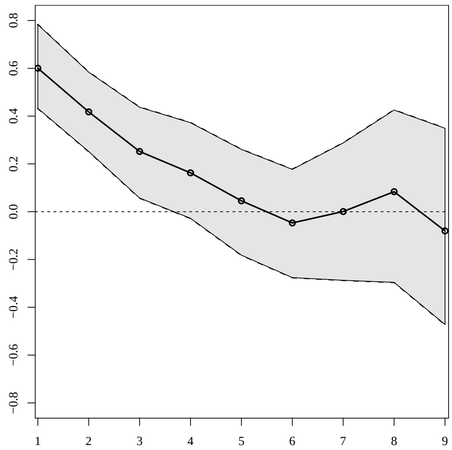


Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.
Note: The variable of interest is the log hourly wage. The graphs report for each decile (x-axis) the impact of a small change in the minimum wage. The y-coordinate corresponds to $\beta_\tau/f(q_\tau)$ *where* β_τ *is the estimated coefficient,* q_τ *is the* τ*th order decile and* f *is the density of the wage distribution. Additional controls include age, year dummies, socio-economic position, industry, GMR group, Size of the firm, seniority. The shaded area corresponds to a 99% confidence interval estimated by bootstrap (500 replications).*

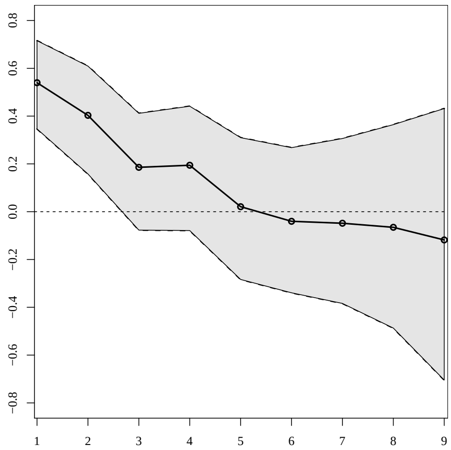
Figure 6: Impact of the minimum wage on the different deciles of log hourly wages

Estimations on two-year periods

Years 2003 and 2004 only

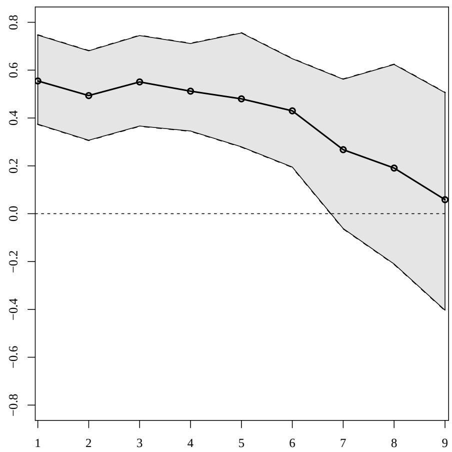


(Men)

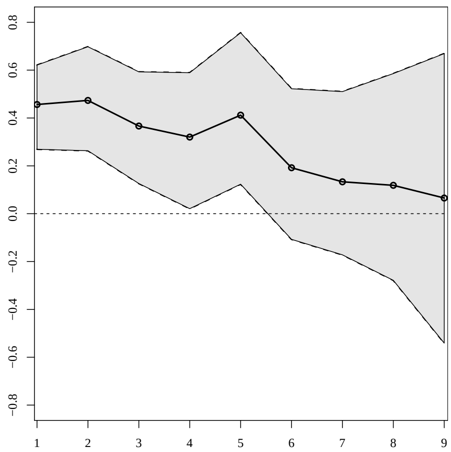


(Women)

Years 2004 and 2005 only



(Men)



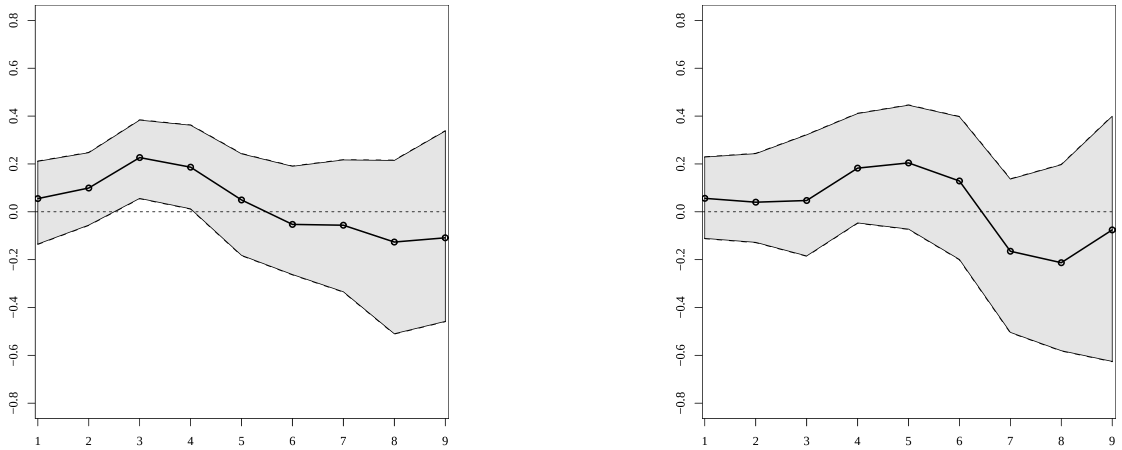
(Women)

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.
Note: The estimation is the same as in the previous figure, but estimated on two-year periods. The shaded area corresponds to a 99% confidence interval estimated by bootstrap (500 replications).

Figure 7: Impact of the minimum wage on the different deciles of log hourly wages

Placebo test

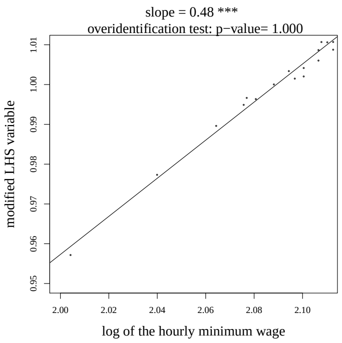
Years 2006 and 2007 only



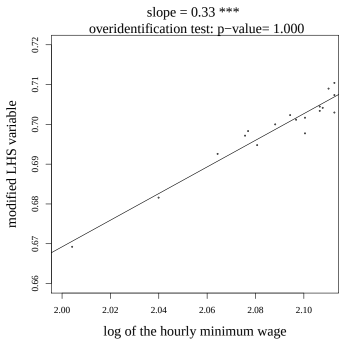
(Men) (Women)

Source: DADS panel, 1/25th sample.
Field: employees from firms of the private sector aged 18 to 65, excluding interns and apprentices.
Note: The estimation is the same as in the previous figures, but estimated on the two-year period 2006-2007 which falls right after the reform and is therefore used to perform a placebo test. The shaded area corresponds to a 99% confidence interval estimated by bootstrap (500 replications).

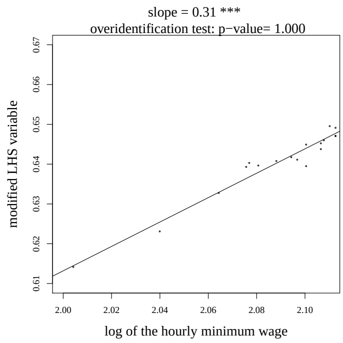
Figure 8: Estimation on clustered data - Men



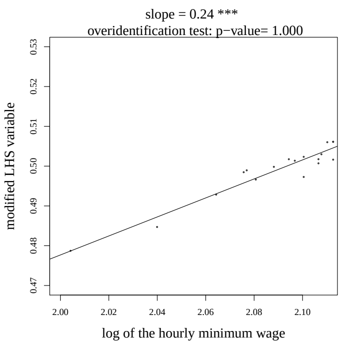
(1^{st} decile)



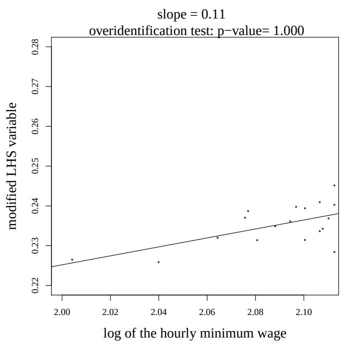
(2^{nd} decile)



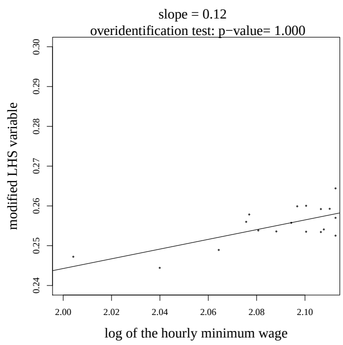
(3^{rd} decile)



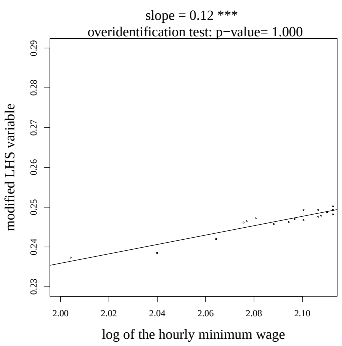
(4^{th} decile)



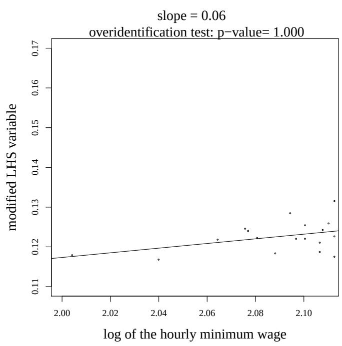
(5^{th} decile)



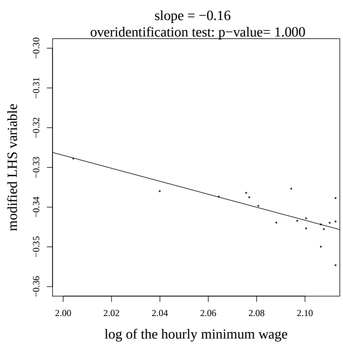
(6^{th} decile)



(7^{th} decile)



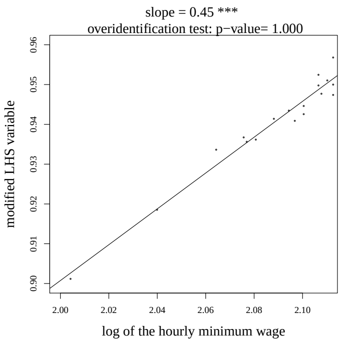
(8^{th} decile)



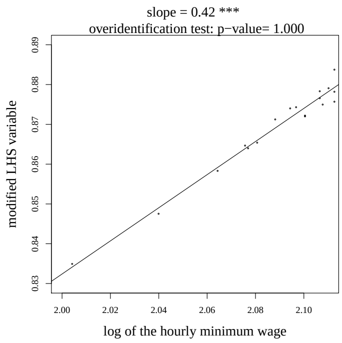
(9^{th} decile)

Note: Each graph corresponds to an estimation on grouped data for a given decile. Each point corresponds to a FMGR×year cluster. The x-axis corresponds to the log hourly minimum wage of the FGMR in July of the previous year. The y-axis corresponds to $c_{1,\tau}$ *multiplied by the proportion of individuals in the cluster whose hourly wage is above the given decile, purged from FGMR fixed effects, year fixed effects and also purged of the effect of other covariates including age, socio-economic position, industry, size of the firm, seniority. The slope can be compared to the coefficients obtained in the previous graphs. The differences are due to the choice of weights in the grouped estimation.*

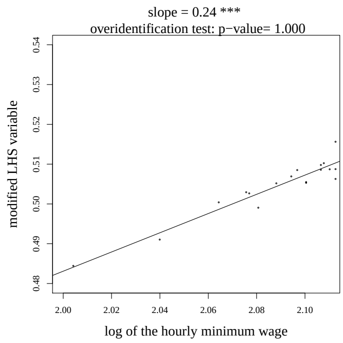
Figure 9: Estimation on clustered data - Women



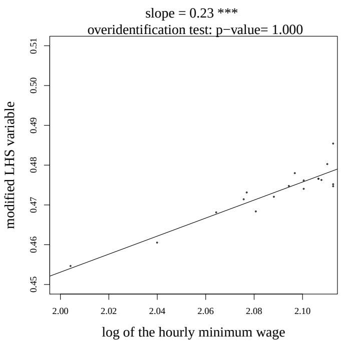
(1^{st} decile)



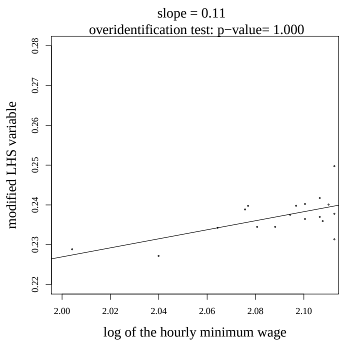
(2^{nd} decile)



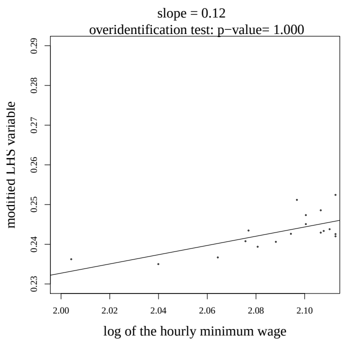
(3^{rd} decile)



(4^{th} decile)

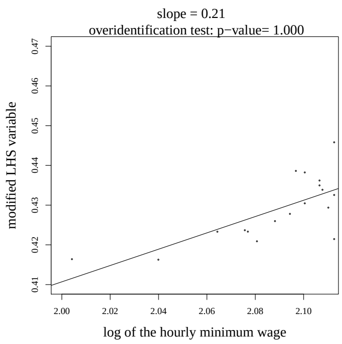


(5^{th} decile)

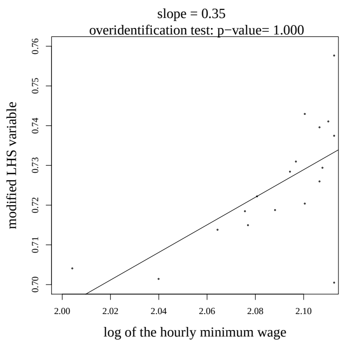


(6^{th} decile)

(7^{th} decile)



(8^{th} decile)



(9^{th} decile)

Note: Each graph corresponds to an estimation on grouped data for a given decile. Each point corresponds to a FMGR×year cluster. The x-axis corresponds to the log hourly minimum wage of the FGMR in July of the previous year. The y-axis corresponds to $c_{1,\tau}$ *multiplied by the proportion of individuals in the cluster whose hourly wage is above the given decile, purged of FGMR fixed effects, year fixed effects and also purged of the effect of other covariates including age, socio-economic position, industry, size of the firm, seniority. The slope can be compared to the coefficients obtained in the previous graphs. The differences are due to the choice of weights in the grouped estimation.*

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