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ABSTRACT

Labor Earnings Respond Differently to Income-Tax and to Payroll-Tax Reforms*

We estimate the responses of gross labor earnings with respect to marginal and average net-of-tax rates in France over the period 2003-2006. We exploit a series of reforms to the income-tax and the payroll-tax schedules that affect individuals who earn less than twice the minimum wage. Our estimate for the elasticity of gross labor earnings with respect to the marginal net-of-income-tax rate is around 0.2, while we find no response to the marginal net-of-payroll-tax rate. The elasticity with respect to the average net-of-tax rates is not significant for the income-tax schedule, while it is close to -1 for the payroll-tax schedule. A plausible explanation is the existence of significant labor supply responses to the income-tax schedule, combined with a short-term rigidity of the hourly taxable wage (i.e. the gross wage minus payroll taxes), casting doubts about public finance analysis that assumes perfect competition on the labor market. Finally, the effect of the net-of-income-tax rate seems to be driven by labor supply participation decisions, in particular those of females.

JEL Classification: H24, H31, J22, J38

Keywords: labor earnings, payroll tax, income tax

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I. Introduction

Labor income taxation is composed of several distinct schedules. According to the OECD tax database,[1] the total tax wedge for an average-wage worker amounts to 29.7% of labor costs to the employer in the U.S. in 2010. This tax wedge can be decomposed into 8.2% for transfers to the "central government", 5.8% for transfers to "sub-central" governments, and 15.8% for "social security contributions". In France at the same time, the total tax wedge for an average-wage earner amounts to 49.3%, with only 9.9% for transfers to the central government and 39.3% for social security contributions. Whether or not earnings respond identically to the different schedules is crucial to determine which type of tax should be used to finance public expenditures, including social security expenses and redistribution. In this work, we focus on the relative responsiveness of labor earnings to *payroll* (social contributions) taxation versus *income* taxation.

The most usual models of the labor market (including the standard labor supply model, the monopoly-union model under right-to-manage, or the individual wage bargaining model) predict identical earnings responses to payroll-tax and to income-tax schedules. By contrast, the empirical evidence is so far not conclusive because the existing literature never simultaneously considers the responses to payroll taxes and to income taxes,[2] one reason being the absence of simultaneous reforms to both schedules for similar individuals over the same period.

In contrast to the literature, we exploit a series of reforms to both the income-tax and the payroll-tax schedules that occurred in France over 2003-2006 for individuals in the bottom half of the wage distribution. In 2003, there existed two distinct schedules for the reduction in employers' payroll taxes for low-wage workers, depending on whether the firm had moved to the 35-hour workweek or had remained at 39 hours. A progressive convergence between the two schedules was implemented from 2003 and achieved in July 2005. This resulted in opposite effects for the two types of firms: an increase in the reduction in employers' payroll taxes for those remained at 39 hours and a decrease in the reduction for those that had moved to the 35-hour workweek. Over the same period of time, the *Prime pour l'Emploi*, a working tax credit for low-wage earners, has been substantially increased, the maximum amount of benefits being almost doubled from 2003 to 2006. Exploiting this rich set of reforms that affected workers who earn less than twice the minimum wage gives us the very unlikely opportunity to compare the responsiveness of labor earnings to income-tax and to payroll-tax reforms.

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[1] Authors' calculations from OECD tax data base available at http://www.oecd.org/dataoecd/44/3/1942514.xls

[2] Most of the papers focus on the distortions induced by income taxes (e.g. Feldstein (1995), Auten and Caroll (1999), Gruber and Saez (2002), Saez (2003), Blomquist and Selin (2010), Cabannes, Houdré and Landais (2011) among others. Another strand of the literature estimate the effects of payroll tax reforms (e.g. Gruber (1997), Kugler and Kugler (2009), Liebman and Saez (2006), and Saez, Matsaganis and Tsakloglu (2011) among others).

The dataset we use is the *Enquête Revenus Fiscaux*, which combines income tax records from the fiscal administration with the French Labor Force Survey (hereafter LFS). Whereas the information contained in income tax records allows us to compute the income tax schedule (including the tax credit for low-wage earners), the LFS provide all the variables necessary to reconstruct employer and employee payroll taxes. In particular, the labor market history and the usual weekly working time enable us to infer a monthly wage and an hourly wage, which are the basic inputs for calculating payroll taxes respectively before and after the reform. We are also able to infer whether the firm has moved to the 35-hour workweek or has remained at 39 hours, which determines which payroll tax schedule applies. To the best of our knowledge, this is the first work that uses income tax records matched to the Labor Force Survey to investigate the responsiveness of labor income to both income-tax and payroll-tax reforms.

We estimate the short-term gross wage (or labor cost) responses to the marginal and average net-of-tax rates[3] for both schedules. We find a significant elasticity (around 0.2) of labor earnings with respect to the marginal net-of-*income*-tax rate. By contrast, the elasticity of labor earnings with respect to the marginal net-of-*payroll*-tax rate is statistically insignificant and close to zero. Labor earnings thus respond differently to payroll-tax and to income-tax changes, at least in the short-run. This finding contrasts with the theorical predictions of the most common labor market models. We also find that the income effects of payroll-tax and income-tax changes are different. The elasticity with respect to the average net-of-*payroll*-tax rate does not differ significantly from minus one, while the elasticity with respect to the average net-of-*income*-tax rate is significantly lower in absolute term.

Our preferred interpretation for these findings is significant labor supply responses to the income-tax schedule, combined with a short-term rigidity of the hourly *taxable* wage (i.e. the gross wage minus payroll taxes) over this period. The effect of an income-tax reform works through rapid labor supply modifications. Further investigations suggest that this response is essentially due to female participation decisions. By contrast, the hourly taxable wage is probably determined through bargaining, which does not occur every year, implying that the response of gross wages to payroll-tax changes is delayed. Our results suggest in particular that, at least in the short-run, financing social security expenses and redistribution through payroll taxes is less distortive than through income taxes (e.g. through a rise in the amount of working tax credit). Our results also imply that, in France, reducing employer payroll taxes for low paid jobs actually decreases the labor cost to the employer, without any significant effect on the net wage. Hence, we do not find that the reduction in employer payroll taxes for low paid jobs has exacerbated a "low-wage trap" in France.[4] Finally, the plausibility of taxable wage rigidity suggests that the assumption of a competitive labor market might be

[3] The marginal (respectively average) net-of-tax rate is equal to one minus the marginal (average) tax rate.

[4] Our result is in line with two other studies using French data, Aeberhardt and Sraer (2009), and Bunel, Gilles, and L'Horty (2009), but differs from that of Lhommeau and Remy (2009) also for France.

inadequate for applied public finance, at least in the short-run and for economies such as France where the labor market is highly regulated. As the latter assumption is made in a large part of the optimal income tax literature (e.g. Mirrlees (1971)), one can question the relevance of the normative predictions derived by these models.

A large strand of the literature is interested in the response of taxable income to the marginal net-of-income-tax rate, following the idea of Lindsey (1987) and Feldstein (1995) that this elasticity summarizes all the deadweight losses due to taxation. We here compare our results with the main conclusions of this literature. *i*) The literature considers that the elasticity is hard to identify, in particular because of the difficulties to control for non-tax related changes in income. We better account for these changes by using additional control variables provided by the Labor Force Survey. *ii*) At a first look, our estimate of 0.2 for the elasticity with respect to the marginal net-of-income-tax rate belongs to the range of the most plausible values found in the literature (between 0.12 and 0.4 according to Saez, Slemrod and Giertz (2010) or around 0.33 for the intensive margin according to Chetty (2011)). However, we here estimate the response of labor earnings while most works are interested in the responses of taxable *total* income. Restricting the comparison to taxable *labor* income, our estimate is consistent with that of Blomquist and Selin (2010), who find significant responses, whereas differing from Saez (2003) who suggests that labor earnings do not significantly respond to taxation. *iii*) The existing literature suggests that the elasticity is presumably much higher for top income earners (e.g. Gruber and Saez (2002)). However, we obtain a significant elasticity of labor earnings with respect to the marginal net-of-income-tax rate by using reforms that affect individuals in the bottom half of the wage distribution. *iv*) Finally, our result that labor earnings are, at least in the short-run, insensitive to the marginal payroll-tax rate is consistent with those found by other studies (e.g. Liebman and Saez (2006), Saez et alii (2011)). More specifically, it is in line with different studies using French data. The novelty is that we exhibit simultaneously the unresponsiveness of taxable wages to payroll taxes and their responsiveness to income taxes.

The paper is organized as follows. Section II defines the different elasticity concepts and discusses theoretically whether labor earnings should respond identically to income taxes and payroll taxes. In section III, we present our empirical strategy and discuss the identification. Section IV describes the dataset used, which combines income tax records with the Labor Force Survey. In section V, we detail the institutional backgrounds and expose the main reforms that took place in France over the 2003-2006 period. Section VI presents results on the respective effects of payroll taxes and income taxes on labor earnings for all employees and for specific subsamples, and the last section concludes.

II. Theoretical backgrounds

II.1 Definitions

Because of the existence of taxes and transfers, the (annual) *net* wage c that a worker can consume differs from the *gross* wage (or labor cost) w paid by her employer. In most countries, labor earnings taxation can be decomposed into social security contributions or *payroll* taxes (that finance social security programs such as PAYG pensions, health insurance, unemployment insurance, etc.) and taxes to governments or *income* taxes.

We represent the payroll tax as a function of the gross wage and we denote it with a superscript P: $T^P(w)$. We denote by z the *taxable* wage, defined as the gross wage minus payroll taxes. On a linear part of the payroll tax schedule, we get:

$$z = \tau^P w + R^P \tag{1}$$

where $\tau^P = 1 - \frac{\partial T^P(w)}{\partial w}$ denotes the *marginal net-of-payroll-tax* rate and R^P is the virtual *taxable* income. We denote by $\rho^P(w) = \frac{z}{w} = \tau^P + \frac{R^P}{w}$ the *average net-of-payroll-tax* rate.

The income tax schedule, denoted with a superscript I, consists of income tax *per se* and of tax credits providing income subsidies to low-wage earners (e.g. *Earned Income Tax Credit* in the U.S., *Prime Pour l'Emploi* in France, etc.). The income tax is a function of the taxable wage z: $T^I(z)$. On a linear part of the income tax schedule, the net wage c equals the taxable wage z minus income taxes, implying that:

$$c = \tau^I z + R^I \tag{2}$$

where $\tau^I = 1 - \frac{\partial T^I(z)}{\partial z}$ denotes the *marginal net-of-income-tax* rate and R^I is the virtual *net* income. We denote by $\rho^I(z) = \frac{c}{z} = \tau^I + \frac{R^I}{z}$ the *average net-of-income-tax* rate. Combining Equations (1) and (2) gives the budget constraint:

$$c = \tau^I \tau^P w + \tau^I R^P + R^I \tag{3}$$

We thus define the *global* marginal net-of-tax rate as $\tau = \tau^P \tau^I$, the *global* virtual income as $R = \tau^I R^P + R^I$, and the *global* average net-of-tax rate c/w as $\rho = \rho^P \rho^I$.

The three wages w, z and c are endogenous and may depend on each of the four tax parameters τ^I, τ^P, R^I and R^P. More precisely, let assume that the gross wage w is determined by a function denoted $W(\tau^I,\tau^P,R^I,R^P)$. Partial derivatives of the function W then define a first set of behavioral parameters:

$$\frac{\Delta w}{w}=\left(\frac{\tau^P}{w}\frac{\partial W}{\partial \tau^P}\right)\frac{\Delta \tau^P}{\tau^P}+\left(\frac{\tau^I}{w}\frac{\partial W}{\partial \tau^I}\right)\frac{\Delta \tau^I}{\tau^I}+\left(\frac{\partial W}{\partial R^P}\right)\frac{\Delta R^P}{w}+\left(\frac{\partial W}{\partial R^I}\right)\frac{\Delta R^I}{w} \tag{4}$$

For $j=I,P$, the elasticity $(\tau^j/w)(\partial W/\partial \tau^j)$ captures the change in gross wages generated by a tax reform that not only modifies the slope of the budget constraint (3), but also the amount of tax paid. This is the reason why it is referred to as the *uncompensated* elasticity with respect to the marginal j net-of-tax rate. However, the theory of optimal taxation puts more emphasis on *compensated* elasticities. These ones capture the response of gross wages to tax reforms that modify the slope of the tax schedule while leaving unchanged the amount of tax paid at the pre-reform earnings level.[5] Formally, let w^* and z^* denote respectively the initial gross and taxable wages, with $z^*= \tau^P w^*+ R^P$. Then, the compensated elasticity with respect to the marginal payroll (respectively income) net-of-tax rate β_τ^P (respectively β_τ^I) captures the effect on gross wages of a payroll (respectively income) tax reform that varies the marginal net-of-payroll-tax rate τ^P (respectively marginal net-of-income-tax rate τ^I) and the virtual taxable income R^P (respectively virtual net income R^I) in such a way that the amount of payroll (resp. income) tax paid at w^* (resp. at z^*) is kept unchanged, i.e. $\Delta R^P = - w^* \Delta\tau^P$ (resp. $\Delta R^I = -z^* \Delta\tau^I$). The compensated elasticities are thus given by the "Slutsky-alike" equations:

$$\beta_\tau^P=\left(\frac{\tau^P}{w^*}\frac{\partial W}{\partial \tau^P}\right)-\tau^P\frac{\partial W}{\partial R^P} \quad \text{and} \quad \beta_\tau^I=\left(\frac{\tau^I}{w^*}\frac{\partial W}{\partial \tau^I}\right)-\tau^I\rho^P\frac{\partial W}{\partial R^I} \tag{5}$$

For a given level of the gross wage w^*, the change $\Delta\rho^P$ in the average payroll net-of tax rate equals $\Delta\tau^P+(\Delta R^P/w^*)$, while the change $\Delta\rho^I$ in the average income net-of tax rate equals $\Delta\tau^I+(\Delta R^I/z^*)-(R^I/w^*)(\Delta\rho^P/(\rho^P)^2)$. Combining (4) and (5) then leads to (See Appendix A.1):

$$\frac{\Delta w}{w^*}=\beta_\tau^P\frac{\Delta\tau^P}{\tau^P}+\beta_\tau^I\frac{\Delta\tau^I}{\tau^I}+\beta_\rho^P\frac{\Delta\rho^P}{\rho^P}+\beta_\rho^I\frac{\Delta\rho^I}{\rho^I} \tag{6}$$

[5] For instance, when the choice of w reflects a labor supply decision in a competitive labor market and the labor demand is infinitely elastic, the compensated elasticity captures only substitution effects while the uncompensated one captures both substitution and income effects. Then, the deadweight losses associated with taxation not being lump-sum are proportional to the compensated elasticity of the labor supply.

where the elasticity β_ρ^P (respectively β_ρ^I) with respect to the average payroll (respectively income) net-of-tax rate is defined by:

$$\beta_\rho^P = \rho^P \left(\frac{\partial W}{\partial R^P} + \frac{R^I}{z^*} \frac{\partial W}{\partial R^I} \right) \qquad \text{and} \qquad \beta_\rho^I = \rho^P \rho^I \frac{\partial W}{\partial R^I} \tag{7}$$

II.2 Foundations

In a large class of models of the labor market, the gross wage (or labor cost) is determined by the maximization of an objective function that depends negatively on the gross wage w to the firm and positively on the net wage c paid to the worker. This objective takes the general form $U(c,w)$ with $U_c^{'} > 0 > U_w^{'}$. We henceforth refer to this class of models as the "benchmark" ones. The textbook labor supply framework is typically one of them. In this framework, a worker of productivity p supplying L units of labor earns a gross wage $w = p\,L$. If her preferences over consumption and labor supply are described by the utility function $u(c,L)$, with $u_c^{'} > 0 > u_L^{'}$, one can define function U by $U(c,w) \equiv u(c,w/p)$. The objective U is decreasing in the gross wage w because earning a higher gross wage w requires working more intensively (i.e. higher L). The monopoly union model (under right-to-manage) is also a benchmark model (Hersoug, 1984). If the union's objective over net wages c and employment L is described by $u(c,L)$ and if the labor demand is described by the decreasing function $L=l^d(w)$, then function U is defined by $U(c,w) \equiv u(c,l^d(w))$. Here, U is decreasing in the gross wage because the labor demand depends negatively on the labor cost. Lastly, wage bargaining settings (e.g. Lockwood and Manning, 1993, Pissarides, 2000) are other examples of benchmark models. In these frameworks, function $U(c,w)$ is given by the generalized Nash product where the worker's (or union's) contribution to the Nash product is increasing in the net wage c, while the firm's contribution is decreasing in w, as higher gross wages lower profits. However, it is worth noting that, for the monopoly union model as well as for the wage bargaining model, the objective function takes the form $U(c,w)$ only if the wage setting concerns homogenous workers and firms, which implies that the wage and the tax schedules are unique. Hence, only bargaining models at the individual level (e.g. as in Mortensen and Pissarides, 1994) or at collective levels but for homogenous labor markets can be reduced to the maximization of this type of objective.

In any benchmark model, the gross wage is determined by the maximization of $U(c,w)$ subject to (3), the taxable wage z being an intermediate variable that is economically irrelevant. Formally, we have:

$$w = \arg\max_w U(\tau w + R, w) \equiv \Omega(\tau, R)$$

This implies that the various tax parameters influence the gross wage only through the *global* marginal net-of-tax rate, $\tau = \tau^P \tau^I$, and the *global* virtual income, $R = \tau^I R^P + R^I$. The function W then verifies the following restriction:

$$W(\tau^I, \tau^P, R^I, R^P) = \Omega(\tau^I \tau^P, \tau^I R^P + R^I) \tag{8}$$

Let β_τ and β_ρ define respectively the compensated elasticity with respect to the *global* marginal and average net-of-tax rates. Then applying the same Slutsky-alike decomposition as made in Equations (5) and (7), we obtain:[6]

$$\beta_\tau = \left(\frac{\tau}{w^*}\frac{\partial \Omega}{\partial \tau}\right) - \tau \frac{\partial \Omega}{\partial R} \tag{5'}$$

$$\beta_\rho = \rho \frac{\partial \Omega}{\partial R} \tag{7'}$$

$$\frac{\Delta w}{w^*} = \beta_\tau \frac{\Delta \tau}{\tau} + \beta_\rho \frac{\Delta \rho}{\rho} \tag{6'}$$

As shown in Appendix A.3, the second-order condition together with the assumption that the objective U is increasing in c ensures that β_τ is positive. Moreover, in the labor supply framework, assuming in addition the normality of leisure implies that β_ρ is negative. Finally, the restriction on W given by Equation (8) implies that:

$$\beta_\tau = \beta_\tau^P = \beta_\tau^I > 0 \qquad \text{and} \qquad \beta_\rho = \beta_\rho^P = \beta_\rho^I \tag{9}$$

Prediction (9) is obtained in the very large class of benchmark models. Therefore, if estimating Equation (6) leads to reject (9), it means that the above-mentioned models are rejected by the data. One should then look for alternative frameworks that can account for such departures. We have *a priori* three alternatives in mind that we now describe.

[6] The case with only one relevant tax schedule can be figured out by considering that τ^P=1 and R^P=0, implying that $\tau=\tau^P$, $R=R^P$ and $\rho=\rho^P$. Equations (5), (7) and (6) then directly give (5'), (7') and (6') (see Appendix A.2).

First, the "salience" (in the sense of Chetty, Looney, and Kroft (2009)) of income-tax reforms and of payroll-tax reforms may be different. For instance, one can argue that, as payroll taxes are paid on a monthly basis while income taxes are paid on an annual basis with one year lag in France, wages react more rapidly to changes in payroll taxes and more slowly to changes in income taxes. In this case, β_τ^I and β_ρ^I are expected to have the same sign but to be lower in absolute terms than β_τ^P and β_ρ^P respectively. Conversely, one may argue that individuals are much more aware of the income tax schedule than of the payroll tax schedule. This would imply that β_τ^I and β_ρ^I have the same sign but are larger in absolute terms than respectively β_τ^P and β_ρ^P.

Second, payroll taxes finance various social programs. For some of these programs, both the eligibility and the benefit level are related to the payroll taxes paid. The most illustrative example is the pension system, where the benefit level received upon retirement depends explicitly on both the level and duration of contributions. Unemployment insurance and health insurance also exhibit this contribution-related property. In case of job loss, the maximum duration of UI benefits depends on the duration of contributions. In case of sick leave, health insurance gives benefits that again depend on the level of past contributions. When payroll taxes *per se* generate benefits with some probability, the objective to maximize must be modified by adding a function of the level of payroll taxes into consumption. Therefore, the gross wage solves:

$$w = W\left(\tau^P, \tau^I, R^I, R^P\right) \equiv \arg\max_{w} U\left(\tau w + R + k\left(\left(1-\tau^p\right)w - R^P\right), w\right)$$

In this specification, the parameter k captures how the overall level of consumption depends on the level of payroll taxes through the payments of various social benefits. Different arguments suggest that k is small. First, benefits depend on wages more *directly* than *indirectly* through the amount of payroll taxes. Second, as the level of benefits paid depends on the whole labor market history (in particular for pensions), current contributions determine only partially this level. Third, these benefits will be given only in the future, and with some probability, which generates discounting. We hence assume that $k < \tau^I$ and $k < \rho^I$. As the objective to be maximized does no longer verify restriction (8), prediction (9) is not verified as well. Actually, when the benefit level depends on the payroll taxes paid, the elasticity with respect to the marginal (average) net-of-payroll-tax rate is lower (in absolute term). We show in Appendix A.4 that (9') is verified instead of (9):

$$0 < \beta_\tau^P < \beta_\tau^I \qquad \text{and} \qquad \left|\beta_\rho^P\right| < \left|\beta_\rho^I\right| \tag{9'}$$

Third, consider a labor supply model where individuals have preferences $u(c,L)$ over consumption c and labor supply L. Under perfect competition, an individual's hourly gross wage equals her marginal product of hours worked. Now, assume instead that the hourly *taxable* wage (denoted s) is rigid. Then, when a worker supplies L units of labor, she receives the taxable wage $z=sL$. Given the budget constraint (2), she thus chooses her labor supply to maximize $U(c,z) = u(c,z/s)$. Therefore, the taxable wage chosen does not depend on the payroll tax parameters, implying that:

$$\frac{\Delta z}{z^*} = \beta_\tau^I \frac{\Delta \tau^I}{\tau^I} + \beta_\rho^I \frac{\Delta \rho^I}{\rho^I}$$

Given that $z = \rho\, w$, this leads to $\Delta w/w = \Delta z/z - \Delta\rho^P/\rho^P$. Therefore, we get instead of (9):

$$\beta_\tau^P = 0 \qquad \beta_\tau^I > 0 \qquad \text{and} \qquad \beta_\rho^P = -1 \tag{9''}$$

III. Empirical strategy

Our objective is to evaluate separately the responses of annual gross wages to reforms to the payroll-tax and to the income-tax schedules. We then estimate the following empirical counterpart of Equation (6) for an individual i employed at t-1 and t:

$$\Delta \log w_{i,t} = \alpha + \beta_\tau^P \, \Delta \log \tau_{i,t}^P + \beta_\tau^I \, \Delta \log \tau_{i,t}^I + \beta_\rho^P \, \Delta \log \rho_{i,t}^P + \beta_\rho^I \, \Delta \log \rho_{i,t}^I + \gamma \cdot X_{i,t-1} + u_{i,t} \tag{10}$$

where Δ is the time-difference operator between dates t and t-1, $X_{i,t-1}$ is a vector of observed individual and firm characteristics at t-1, and $u_{i,t}$ is an error term that captures unobserved heterogeneity. In specifying the empirical setup, we are aware that heterogeneous individuals may respond to tax changes differently. Hence, we only provide evidence on the average of these behavioral elasticities, i.e. on the Local Average Treatment Effect (LATE).

Various methodological issues complicate the estimation. A first issue is the existence of non-tax related changes in gross wages. For example, technical progress and international trade generate changes in gross wages, that are likely to be different across firm size and industry, age category, level of education, etc., and presumably lead to a widening of the wage distribution. The risk when evaluating a tax reform that reduces the marginal tax rate for top income earners, such as TRA86 in the U.S., is to attribute changes in wages to the reform rather than to "non-tax" causes, thereby causing an upward bias in the elasticity estimate (Gruber and Saez (2002)). To better account for these non-tax related changes, we include in $X_{i,t-1}$ additional control variables provided by the LFS, which describe individual and firm characteristics.

Reversion-to-the-mean constitutes another source of non-tax factors. An individual with an unusually low (respectively high) labor income in period t-1 is very likely to have a higher (lower) one at t. This is typically what happens when an individual enters unemployment (or involuntary part-time work) during year t-1. Her labor income is then unusually low and increases substantially in year t if she finds a permanent (or full-time) job. In order to control for reversion-to-the-mean, the standard procedure in the literature is to include a flexible function of base-year income, $f(\log w_{i,t-1})$, in the vector of controls $X_{i,t-1}$. Auten and Caroll (1999) use a linear function while Gruber and Saez (2002) use a very flexible 10-piece spline. As a flexible specification allows one to better control for reversion-to-the-mean but may destroy identification (Saez, Slemrod and Giertz (2010)), we choose to include a fifth-order polynomial of the log of base gross wages $log(w_{i,t-1})$ in the list of controls.

However, including $f(\log w_{i,t-1})$ and individual and firm characteristics in $X_{i,t-1}$ may be insufficient to control for mean reversion, trends in the gross wage distribution or unobserved heterogeneity.[7] In particular, $\log w_{i,t-1}$ is likely to be correlated with $u_{i,t}$ if the residuals of Equation (10) are serially correlated. Formally, if $u_{i,t}$ is not a white noise process, then $\mathbf{E}\left(u_{i,t} \middle| \log w_{i,t-1}\right)$ is different from zero (Holmlund and Söderström, (2008), Blomquist and Selin (2010)), implying that $\log w_{i,t-1}$ (and the subsequent terms of the polynomial) must be instrumented. As our dataset provides information on gross wages in year t-2, a natural instrument for $w_{i,t-1}$ is $\overline{w}_{i,t-2} = w_{i,t-2} \times \pi_{t-2}$, where π_{t-2} denotes the average growth rate of (nominal) gross wages between years t-2 and t-1. This instrument is valid, provided that the error term $u_{i,t}$ follows a MA(1) process.

Another issue concerns the potential simultaneity bias. Because of the nonlinearity of, respectively, the payroll-tax and the income-tax schedules, the marginal net-of-tax rates $\tau_{i,t}^{P}$ and $\tau_{i,t}^{I}$ are functions of the gross wage level. Similarly, the average net-of-tax rates $\rho_{i,t}^{P}$ and $\rho_{i,t}^{I}$ depend on the gross wage as well, the reason being that the marginal and the average rates differ. In order to isolate the impact of taxes on gross wages, we need instruments for $\Delta \log \tau_{i,t}^{j}$ and $\Delta \log \rho_{i,t}^{j}$, with $j=P,I$. In the literature, the standard procedure, proposed by Auten and Caroll (1999), uses the predicted change in the log of the net-of-tax rate would the real wage do not change from year t-1 to year t. By construction, the instrument captures changes in the tax rate absent any behavioral response. We apply this method separately to the marginal and the average net-of-tax rates associated with the two tax schedules. Let $\overline{w}_{i,t-1} = w_{i,t-1} \times \pi_{t-1}$ denote the base-year inflation-adjusted gross wage, where

[7] In the literature, there is no proof that this procedure is sufficient.

π_{t-1} is the average growth rate of gross wages between years t-1 and t. We note with an upper bar the net-of-tax rates obtained by applying the year-t tax schedule to $\overline{w}_{i,t-1}$. We then define for j =P, I:

$$\overline{\tau}^{j}_{i,t} = 1 - \frac{\partial T^{j}\left(\overline{w}_{i,t-1};t\right)}{\partial w} \quad \text{and} \quad \overline{\rho}^{j}_{i,t} = 1 - \frac{T^{j}\left(\overline{w}_{i,t-1};t\right)}{\overline{w}_{i,t-1}}$$

For j = P, I and $\varphi = \tau, \rho$, we define the "type-I" instrument as $\Delta \log \overline{\varphi}^{j}_{i,t} = \log \overline{\varphi}^{j}_{i,t} - \log \varphi^{j}_{i,t-1}$.

Moreover, $\mathbf{E}\left(u_{i,t} \middle| \log w_{i,t-1}\right) \neq 0$ may also (but not necessarily[8]) imply that $\Delta \log \overline{\varphi}^{j}_{i,t}$ is not a valid instrument for j =P, I and $\varphi = \tau$, ρ. This leads us to propose a second group of instruments based on year t-2 inflation adjusted gross wages $\overline{\overline{w}}_{i,t-2} = w_{i,t-2} \times \pi_{t-2} \times \pi_{t-1}$ for year t, and on $\overline{w}_{i,t-2} = w_{i,t-2} \times \pi_{t-2}$ for year t-1. We then define, for j =P, I:

$$\overline{\overline{\tau}}^{j}_{i,t} = 1 - \frac{\partial T^{j}\left(\overline{\overline{w}}_{i,t-2};t\right)}{\partial w} \quad \text{and} \quad \overline{\overline{\rho}}^{j}_{i,t} = 1 - \frac{T^{j}\left(\overline{\overline{w}}_{i,t-2};t\right)}{\overline{\overline{w}}_{i,t-2}}$$

$$\overline{\tau}^{j}_{i,t-1} = 1 - \frac{\partial T^{j}\left(\overline{w}_{i,t-2};t-1\right)}{\partial w} \quad \text{and} \quad \overline{\rho}^{j}_{i,t-1} = 1 - \frac{T^{j}\left(\overline{w}_{i,t-2};t-1\right)}{\overline{w}_{i,t-2}}$$

Using the above definitions, type-II instrument for $\Delta \log \varphi^{j}_{i,t}$ is given by $\Delta \log \overline{\overline{\varphi}}^{j}_{i,t} = \log \overline{\overline{\varphi}}^{j}_{i,t} - \log \overline{\varphi}^{j}_{i,t-1}$, with j=P, I and $\varphi = \tau$, ρ. Having two instruments for $\Delta \log \varphi^{j}_{i,t}$ allows us to test their validity.[9]

IV. The data

The existing empirical literature uses either administrative income tax records (e.g. Feldstein, 1995, Auten and Caroll, 1999, Gruber and Saez, 2002) or payroll tax records (e.g. Saez *et alii*, 2011). Although administrative tax records have the advantage of providing exhaustive and longitudinal data, they contain limited information on individual characteristics and no information on labor market history and firms characteristics. Since the main goal for collecting these data is policy-oriented, only

[8] The endogeneity of $w_{i,t-1}$ does not necessarily imply the endogeneity of the instruments, because $\overline{\tau}^{P}_{i,t}$, $\overline{\tau}^{I}_{i,t}$, $\overline{\rho}^{P}_{i,t}$ and $\overline{\rho}^{I}_{i,t}$ are non-monotonic functions of $w_{i,t-1}$.

[9] Holmlund and Söderström (2008) use the same instrument as our type-II one. Blomquist and Selin (2010) propose another strategy. They regress both $w_{i,1991}$ and $w_{i,1981}$ on $w_{i,1986}$ and on a set of controls and use the predicted values to build their instruments. In any case, instruments are valid only under some assumption on the serial correlation of the residuals.

the variables necessary to compute taxes are provided. In contrast to the existing literature, we use a research-oriented dataset, the *Enquête Revenus Fiscaux* (hereafter ERF), resulting from the match of the French Labor Force Survey to administrative income tax records. Specifically, the individuals interviewed at the 4th quarter of year *t* in the LFS are matched with their *t*+1 administrative income tax records to generate the year-*t* wave of the ERF dataset. The LFS is a rotating 18-month panel that starts a new 18-month wave every quarter. As individuals are interviewed during six consecutive quarters, they are at best present during two consecutive years in the ERF dataset. We use the 2003-2006 waves of the ERF because reforms to both the payroll-tax and the income-tax schedules occurred during this period for similar individuals. The individuals sampled thus appear either in 2003 and 2004, in 2004 and 2005, or in 2005 and 2006. As the LFS contains detailed information on personal characteristics (in particular education), on labor market history and on the characteristics of the job (in particular usual weekly hours of work, industry), we are able to control in a rich way for mean reversion and for other trends in the gross wage distribution.

We now describe the wage variable we use. The year-*t* administrative income tax records contains information on the annual *taxable* wage earned by each member of the household (i.e. z defined as the gross wage minus payroll taxes) at dates t-2, t-1, and t. The variable is reported by the employer and controlled by the fiscal administration, and as such is reliable. However, our aim in this work is to evaluate the respective effects of payroll taxes and income taxes on *gross* wages. While the ERF does not directly provide gross wages, it contains information useful to reconstruct this variable by applying the legislation on employer and employee payroll taxes. Payroll taxes are paid each month and are calculated as a function of the monthly and the hourly taxable wages. They depend on the firm size, the type of work,[10] and whether or not the firm has moved to the 35-hour workweek. As the LFS contains all these variables (with the exception of the 35-hour workweek, which is inferred), we can reconstruct payroll taxes, and thereby gross wages. The monthly taxable wage is obtained by matching the annual amount drawn from tax records to the labor market history described by the LFS.[11]

We build our own simulator for the payroll tax system. For the income tax schedule, we use a tax simulator adapted from the INES (INsee Etudes Sociales) micro-simulation model provided by INSEE and DREES. Using these simulators, we compute the average payroll and income taxes, and simulate the marginal payroll and income tax rates generated by a 5% increase in taxable wages. As administrative tax records provide also information on taxable wages at t-1 and t-2, we are able to compute our two types of instruments: instrument I based on $w_{i,t-1}$ and instrument II based on $w_{i,t-2}$.
We restrict the sample to individuals who experienced no change in their marital status between dates t-1 and t, since those who marry, divorce, or become widowed have to report several fiscal returns. In

[10] Engineers, managers and professionals face a specific payroll tax code.

[11] Each quarter, the individuals interviewed in the LFS report their taxable monthly wage. As this information is self-declared, it is less reliable than the annual taxable wages provided by the fiscal administration, and we do not use it.

addition, we exclude public sector workers, as they face up very specific labor market regulations, and the self-employed. We also exclude the individuals for whom we cannot determine with a high degree of certainty whether or not they work under the 35-hour arrangement.[12] Finally, we restrict the sample to employees who report a positive labor income at dates *t*-1 and *t*. Our final sample comprises 12,512 individuals observed at two consecutive years.

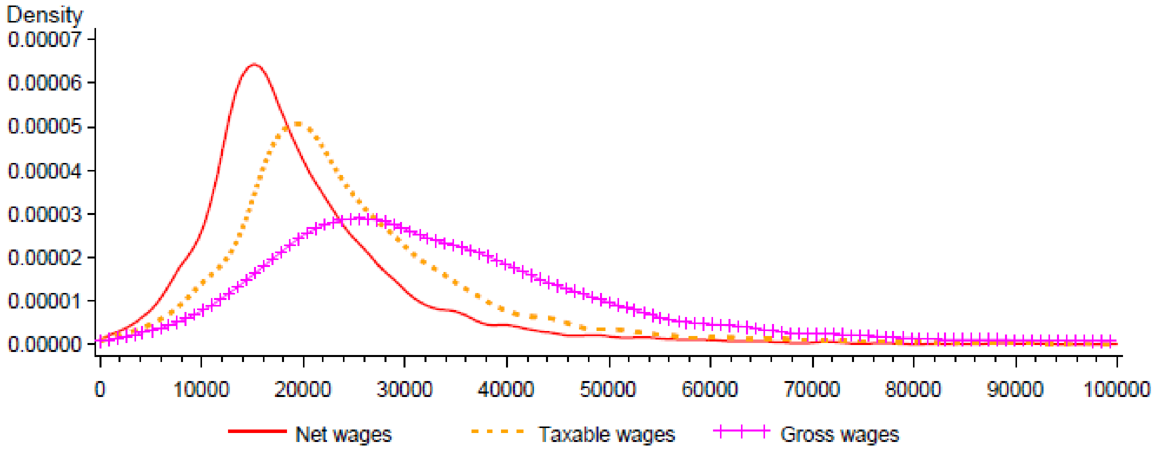

Figure 1: **The distribution of net, taxable, and gross wages in 2004**.

The distributions of the annual gross wage (*w*), taxable wage (*z*), and net wage (*c*) on our sample in 2004 are displayed in Figure 1. The three distributions appear hump-shaped, with a fat upper tail (particularly for gross wages). Due to the high level of payroll taxes in France, the distribution of taxable wages lies far to the right of the distribution of taxable wages, which itself lies slightly to the right of the distribution of net wages.

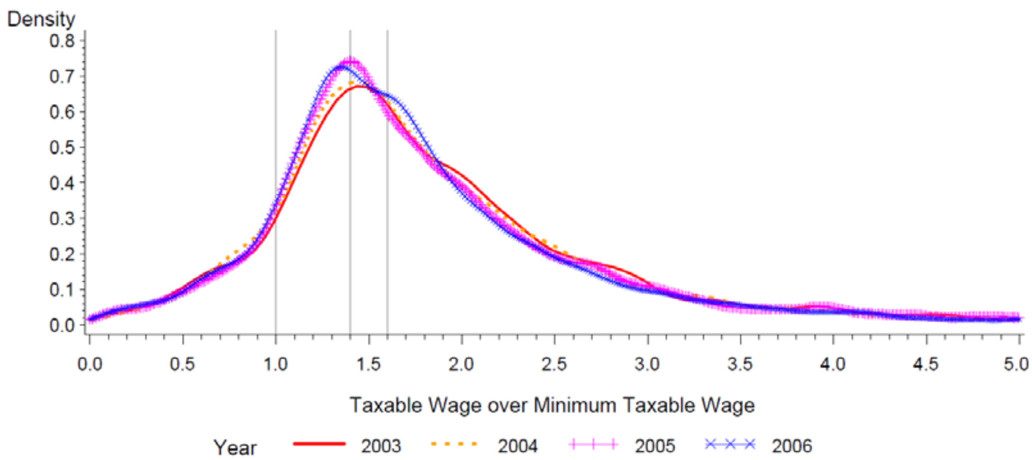

Figure 2: **The distributions of the ratio of taxable wages to the minimum wage, 2003-2006**.

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[12] Using the legal and usual weekly hours of work, and the existence of additional vacations (RTT), we are able to infer whether the individual is working in a 35-hour firm or in a 39-hour firm at dates *t* and *t*-1. Given that, since 2003, few firms have moved to the 35-hour workweek, we are confident about the firm working time when it is the same at dates *t* and *t*-1. As a result, we restrict the sample to employees who work either 35 hours or 39 hours at *t* and *t*-1.

Figure 2 describes the distributions in 2003-2006 of the ratio of the individual's taxable wage to the taxable wage of an individual working full-time for the full-year at the minimum wage, hereafter the "annual minimum wage". For each year, the mode of the distribution of the ratio is around 1.4. However, the proportion of employees between 1 and 1.4 times the minimum wage increased, essentially between 2004 and 2005, while the proportion around 2 times the minimum wage decreased slightly.

Age		**Economic activity**	
< 20 years	0.1 %	Agriculture	1.5 %
20 - 29 years	13.4 %	Manufacturing	26.8 %
30 - 39 years	29.4 %	Construction	7.2 %
40 - 49 years	33.4 %	Energy	1.6 %
50 - 59 years	22.9 %	Education and social activities	9.9 %
≥ 60 years	0.8 %	Trade and repair	17.0 %
Gender		Other tertiary	35.9 %
Women	42.1 %	**Job tenure**	
Men	57.9 %	< 1 year	6.0 %
Household composition		1 - 5 years	25.4 %
Single individual	11.1 %	5 - 10 years	18.6 %
Single parent	6.3 %	≥ 10 years	50.0 %
Couples without children	20.3 %	**Firm size**	
Couples with children	59.5 %	< 10 employees	13.6 %
Other households	2.8 %	10-19 employees	7.0 %
Change in the number of children		≥ 20 employees	79.4 %
Birth of a child between *t* and *t*-1	5.5 %	35-hour workweek	76.0 %
Departure of a child between *t* and *t*-1	6.2 %	35-hour workweek and < 20 employees	8.6 %
No change	88.3 %	35-hour workweek and ≥ 20 employees	67.4 %
Level of education			
College (> 2 years)			11.1 %
College (≤ 2 years)			17.5 %
High school graduate			16.0 %
High-school drop-out or vocational diploma			38.3 %
Junior high school or basic vocational			7.5 %
No diploma or elementary school			9.6 %
N° observations			12 512

Table 1: **Descriptive statistics**

Some summary statistics are presented in Table 1. Due to the selection criteria, those who are under the age of 30 and over the age of 60, as well as women, are under-represented in the sample. Only 3 over 4 employees work in a firm under the 35-hour workweek regulation, even though the working time reduction is legally enforced since 2000 for large firms and 2002 for small firms. If employees in large firms are more likely to work in a firm under the 35-hour workweek regulation than employees in small ones, a significant part of them continue to work in a firm under the 39-hour workweek regulation (15% of all individuals working in a firm with more than 20 employees). Family events like the birth of a child or a child leaving the fiscal household occur for respectively 5.5% and 6.2% of the individuals.

Figure 3: **Means of the growth rate of gross wages for each percentile of the wage distribution**

Figure 3 describes for each base year the growth rate of gross wages ($\Delta \log w_{i,t}$) along the wage distribution. To make the curves comparable across time, we represent the growth rate as a function of the ratio of gross wages to the annual full-time minimum wage. Given the variability of growth rates among individuals with the same earnings level, we compute the means within each percentile of the taxable wage for each year. Figure 3 stresses the reversion-to-the-mean phenomenon at the bottom end of the wage distribution. The most plausible explanation for this fact is exit from unemployment/entry into stable employment between years *t*-1 and *t*.

V. Institutional backgrounds

We now describe the reforms to the payroll-tax and the income-tax schedules that occurred in France during the 2003-2006 period.

V.1 Income tax reforms

By "income tax" we designate the sum of the income tax *per se* and of a tax credit for low-paid earners (*Prime pour l'emploi*, hereafter PPE). Income tax *per se* in France is calculated at the fiscal household level, which differs from the usual notion of household: two persons who live as a couple are considered by the administration as a single fiscal household only if they are married or linked by a civil contract[13]. The income-tax schedule is a function of the ratio of the total income earned by the fiscal household to a weighted sum of its members. The amount of tax paid then equals

[13] Pacte civil de solidarité (PACS).

the income tax that would be paid by a single individual whose income is equal to this ratio, multiplied by the weighted sum. This implies that both the marginal and average net-of-tax tax rates of a given individual change with marital status, spouse's income, at the birth of a child, or when a child becomes adult and exits the household. However, these events are likely to affect the labor supply decisions, the only exception being the exit of a child which generates an instantaneous change in the tax schedule, while the change in the labor supply, if any, is likely to be smoothed over time. Therefore, income tax reforms provide more convincing sources of identification than these family events. Nevertheless, thanks to the complexity of the tax schedule, the very large range of income tax rates that can face different individuals with the same income improves the identification possibilities associated with a fiscal reform.

Over the 2003 – 2006 period covered by our dataset, there are several changes in the income tax code *per se*. In 2004 and 2005, tax brackets are indexed for consumer price inflation. This generates a form of bracket 'creeps' (Saez (2003)), as labor earnings have increased more rapidly than inflation over this period. A more substantial reform in 2006 reduces the number of brackets from seven to five and modifies the rates.

However, the reform that generates the larger changes in tax rules over 2003-2006 is the increase in the *Prime pour l'Emploi*, a tax credit conditional on working that was created in 2001. Both eligibility for the tax credit and the amount of subsidy paid depend essentially on the individual full-time equivalent annual wage, but the total income earned by the household and its composition also intervene. More precisely, eligibility requires that the individual's annual wage is above 0.3 and that her full-time equivalent annual wage is lower than 1.4 times the annual minimum wage for a single worker without children (up to 2.1 times the annual minimum wage for some household compositions). Considering the wage distribution in France,[14] this implies that one-third of the employees are eligible for the working tax credit.[15] We now describe the scheme for a single individual without children. There is a phase-in range between 0.3 and 1 time the annual minimum wage in which the tax credit is proportional to the wage. Unlike the EITC in the US, there is no plateau range: the phasing-out range is between 1 and 1.4 times the annual minimum wage. The entrance in the phase-in income range leads to a reduction in both marginal and average tax rates. The entrance in the phase-out income range is associated with a rise in the marginal tax rate, since a higher wage reduces the tax credit. The average tax rate is minimal at the minimum wage level and then increases. While the PPE scheme remains essentially unchanged in 2004 with respect to 2003, major changes occur both in 2005 and 2006, as described in Figure 4. As a result, the maximum level of the

[14] In France in 2006, 22% of the employed earn a wage between 0.3 and 1.4 times the minimum wage, and 50% get a wage between 0.3 and 2.1 times the minimum wage.

[15] Compared to the EITC or the WFTC, the French tax credit thus differs on two points: a much larger share of the population is eligible; the presence of children has a very limited effect on the amount of benefit.

subsidy increases from 4.6% of the annual minimum wage (i.e. 517€ per year) in 2003 to 7.7% in 2006 (i.e. 948€ per year).

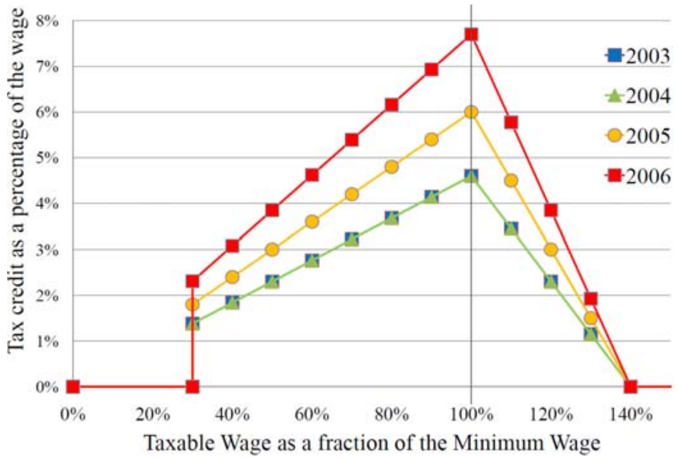

Figure 4: **Reforms to the French income tax credit, 2003-2006**

Note: the amount of PPE is expressed as a percentage of yearly labor earnings. The figure describes the scheme for a single worker without children. For couples or singles with children, the "phasing-out" of the PPE scheme can go up to 2.1 times the minimum wage.

Figure 5a (respectively 5b) depicts the evolution of the marginal 1-τ^I (respectively average 1-ρ^I) income-tax rate simulated on our sample through the years 2003-2006. Even though the rates are very noisy, especially for part-time workers below the full-time minimum wage, for each year we observe that the marginal rate is much higher between 1 and 1.4 times the annual minimum wage than elsewhere. Moreover, as expected, the increase in the tax credit from 2003 to 2006 leads the marginal income-tax rate to increase significantly in this phase-out range. It also reduces the average income-tax rate, especially at the minimum wage level where the PPE is maximal. The tax reforms generated by the income-tax *per se* are conversely much less apparent, except for the reduction in the average tax rate between 2005 and 2006 for gross wages above two times the annual minimum wage.

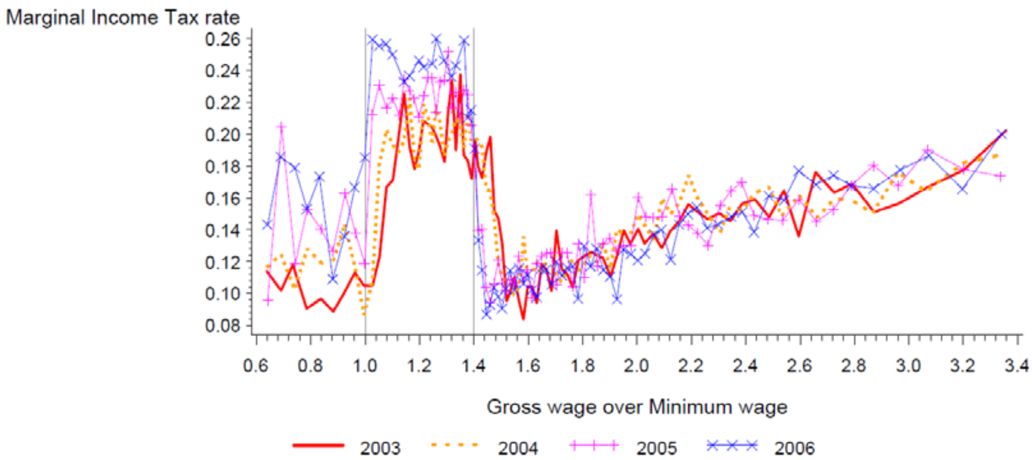

Figure 5a: **Means of Marginal Income tax rates for each percentile of the wage distribution**

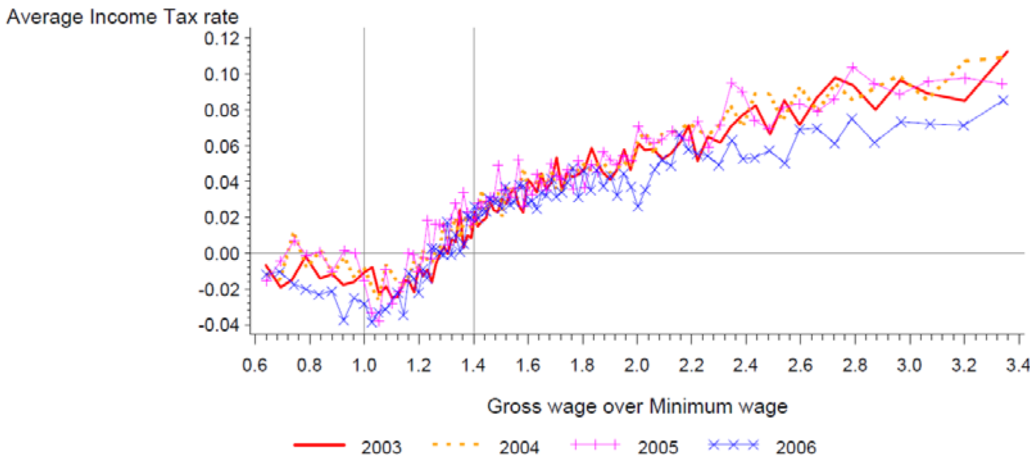

Figure 5b: **Means of Average Income tax rates for each percentile of the wage distribution**
Sample: Employees present two consecutive years. Source: ERF survey, Insee, 2003-2006.

V.2 Payroll tax reforms

In almost all countries, payroll taxes are flat, apply only to earnings below a given cap and are roughly invariant over time. As they are flat and time-invariant, their effect is difficult to identify, which explains why there are few empirical works on how payroll taxes affect labor supply and labor demand. By contrast, in France, the rate of payroll taxes paid is a function of wages since the introduction in 1993 of a reduction in employer payroll taxes for low-wage employees (see e.g. Kramarz and Phillipon (2001) for a description of the policy and an evaluation of its effects on employment).

Since June 1998, the situation was complicated further by the existence of two reduction schedules, depending on whether or not the employee is working under the 35-hour arrangement. At this date, a law implemented by a left-wing government initiated the move to a 35-hour workweek, a process that became in principle mandatory for large firms (more than 20 employees) in January 2000 and small firms in January 2002. The firms moving from a 39-hour to a 35-hour workweek (hereafter the “35-hour firms”) were given an additional reduction in employer payroll taxes compared to the firms remained at 39 hours (hereafter the “39-hour firms”). As all firms were intended to move to the 35-hour workweek, the existence of two types of tax subsidies was not a long-term problem. However, the process of reduction in the workweek was stopped in June 2002 when a right-wing government came into power. At this date, a non-negligible proportion of firms had not moved to the 35-hour workweek and had no intention to move later (Table 1). As a result, a law in January 2003 was planning the creation of a unique tax subsidy schedule, that applied whether or not the firm had moved to the 35-hour workweek. The convergence process lasted three years.

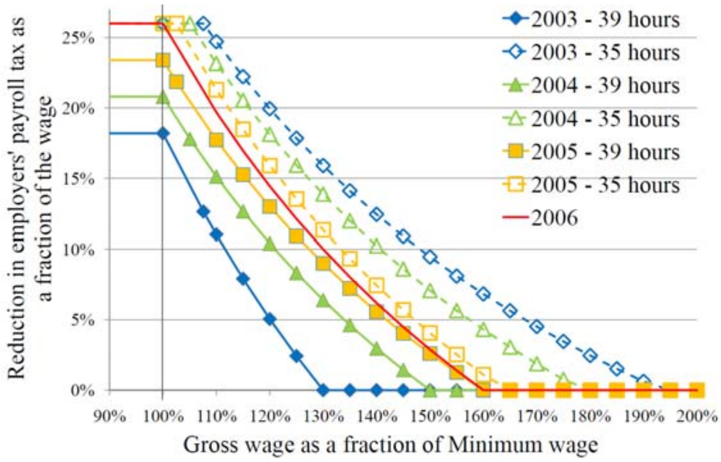

Figure 6: **Changes in the reduction in employer payroll taxes for low paid earners**

Figure 6 presents the changes in the tax subsidy from 2003 to 2006 for the two types of firms. At the beginning, in January 2003, the two subsidy schedules differ substantially. For a 39-hour firm (solid curves), the reduction in employer payroll taxes reaches a maximum of 18.2 percentage points at 1 times the hourly minimum wage, then decreases up to 1.3 times the minimum wage. For a 35-hour firm (dashed curve), the reduction amounts to a maximum of 26 percentage points at 1.076 times the hourly minimum wage, then decreases up to 1.94 times the minimum wage.[16] For the 39-hour firms, the maximum percentage points of reduction increases from 18.2 in 2003 to 26 in 2006. Moreover, the phase-out income range widens from an initial 1 to 1.3 times the minimum wage to 1 to 1.6 times the minimum wage. For the 35-hour firms, the maximum percentage points of reduction remains unchanged, whereas the phasing-out part of the schedule shifts to the left from 1.076 to 1.94 times the minimum wage to 1 to 1.6 times the minimum wage. In average over the period 2003-2006, the tax subsidy decreases for the 35-hour firms while increasing for the 39-hour ones.

Simulating the payroll taxes on our sample of individuals, we find that the marginal payroll tax rate is very high on the phase-out part of the schedule, amounting in 2006 to 57 % between 1 to 1.6 times the minimum wage versus 43 % above 2 times the minimum wage. For those working 35 hours, Figure 7a shows that the income range with very high marginal tax rates shrinks from 2003 to 2006 but that these marginal rates are still higher, as expected from the description of the reform. Turning to average tax rates (Figure 7b), we observe that they do not change over time at the minimum wage level and above 2 times the minimum wage. However, the gross wage above which the average payroll tax rate is the highest diminishes, as time goes on.

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[16] In order to prevent the reduction in the work week to lower the monthly wage, the (hourly) minimum wage regulation (SMIC for *Salaire Minimum Interprofessionnel de Croissance*) was supplemented by a system of monthly minimum wage regulation called GMR (*Guarantie Mensuelle de Rémunerations)* which depends also on the date at which firms adopted the 35-hour work week. In 2003 for a firm having moved to the 35-hour in 2000, the GMR2 is equal to 1.076 times the minimum wage. The reduction is maximal at 1 times the GMR and decreases up to 1.8 times the GMR. The GMR has decreased with respect to the minimum wage from 2003 to 2006.

Figure 7a: **Means of Marginal Payroll tax rates for each percentile of the wage distribution in the 35-hour firms subsample**

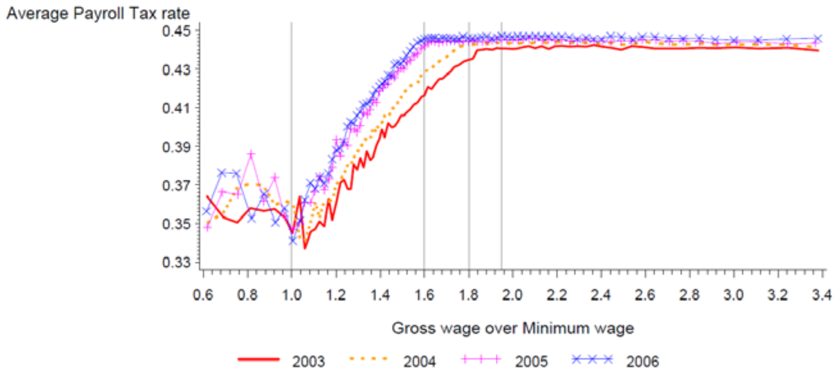

Figure 7b: **Means of Average payroll tax rates for each percentile of the wage distribution in the 35-hour firms subsample**

Sample: Employees present two consecutive years. Source: ERF survey, Insee, 2003-2006.

For those working 39 hours, we observe a rise over time in the marginal payroll tax rate for wages comprised between 1.3 and 1.6 times the minimum wage, as expected from the widening of the phase-out part of the schedule (Figure 8a). By contrast, the average payroll tax rate is significantly reduced at the minimum wage level, following the increase in the maximum percentage points of reduction. The decrease in average tax rates vanishes progressively as one moves rightwards along the wage distribution.

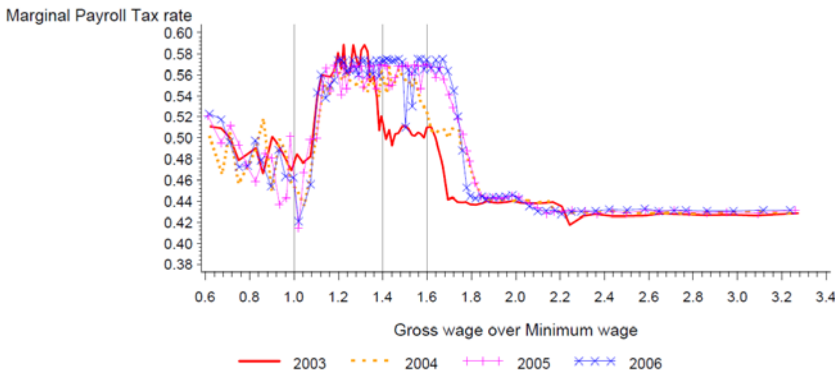

Figure 8a – **Means of Marginal Payroll tax rates for each percentile of the wage distribution in the 39-hour firms subsample**

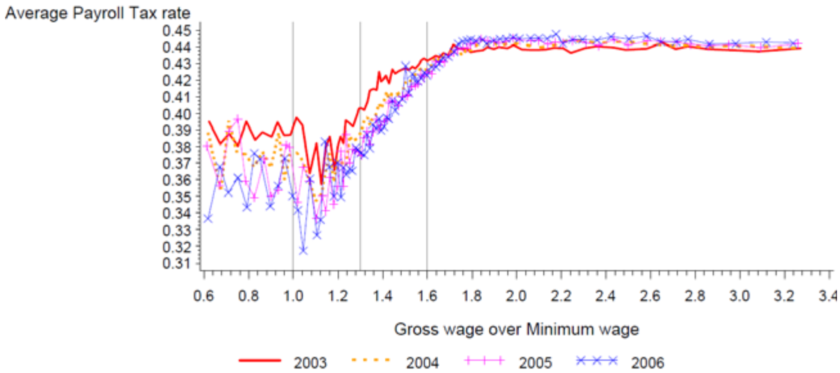

Figure 8b - **Means of Marginal Payroll tax rates for each percentile of the wage distribution in the 39-hour firms subsample**

Source: ERF survey, Insee, 2003-2006. Sample: Employees present two consecutive years.

VI. Results

VI.1 Effects of payroll and income taxes

We estimate Equation (10) using the 2SLS approach. The Sargan test leads us to conclude that type-I and type-II instruments are valid. The first-stage regressions are displayed in Table B.1 of Appendix B. The F-statistics of the excluded instruments are always high, meaning that the instruments are strongly correlated with the instrumented regressors. As type-I instruments are better for marginal net-of-tax rates, and type-II instruments better for average net-of-tax rates, we choose to use both of them.

Table 2 displays our estimates of the gross wages responses for various sets of controls. The full results are presented in Table B.2 in Appendix B. In Column 1 of Table 2, the only covariates are time dummies and a fifth-order polynomial of the log of base year gross wages. In Column 2, we add socio-demographic covariates drawn from tax records concerning age, gender, and the composition of the household. In Column 3, the set of covariates furthermore includes variables drawn from the LFS. These variables describe the educational level, type of occupation, firm size, industry, job tenure, and whether the firm has moved to the 35-hour workweek. We also introduce dummies interacting the working time regulation of the firm and the year. Our objective in including these dummies is to account for the differences in minimum wage regulations between the two types of firms over 2003-2006. In 2002, the binding minimum wage regulation (GMR) for 35-hour firms implied a minimum wage 8.2% higher than for 39-hour firms. Similarly as with the payroll tax schedule, a progressive convergence between the two minimum wage regulations was implemented in 2003 and achieved in July 2005.

We first examine the elasticity with respect to the marginal net-of-income-tax rate. The elasticity estimate is slightly above 0.2, significant, and rather robust to changes in the set of covariates. Our results are thus consistent with those of Blomquist and Selin (2010) for males in Sweden, who find an elasticity of taxable labor income around 0.25. Our estimate also belongs to the [0.12; 0.4] interval that Saez, Slemrod and Giertz (2010) argue is the most plausible. It is worth noting, however, that they consider taxable income, not taxable *labor* income, and believe the former to be more responsive to taxation. In particular, Saez (2003), comparing the effect of taxes on different definitions of income in the U.S. over the period 1979-1981, finds significant responses of both taxable income and adjusted gross income, while the elasticity of labor income is close to zero. Another difference with the existing literature is that we use reforms whose stronger effects occur for individuals below the median of the wage distribution, whereas most of the literature uses reforms occurring in the top of the distribution. High-income individuals are expected to be more sensitive to taxes, in particular because they have more access to avoidance opportunities. The literature usually considers that responses to marginal tax rates are essentially driven by those individuals (e.g. Gruber and Saez (2002)). Our results thus show that significant responses may also arise for low or median-income individuals.

	No covariate (1)	Tax records covariates (2)	Tax records & LFS covariates (3)
β_τ^I	0.240*** (0.085)	0.238*** (0.086)	0.225*** (0.085)
β_τ^P	-0.010 (0.095)	0.001 (0.096)	0.016 (0.096)
β_ρ^I	-0.728** (0.329)	-0.78** (0.385)	-0.694* (0.395)
β_ρ^P	-1.437*** (0.289)	-1.390*** (0.292)	-1.749*** (0.511)
$(9):\ \beta_\tau^I = \beta_\tau^P$ and $\beta_\rho^I = \beta_\rho^P$	5.15*** [0.6%]	3.73** [2.4%]	4.50** [1.1%]
$(9''):\ \beta_\tau^P = 0$ and $\beta_\rho^P = -1$	1.15 [31.7%]	0.89 [41.03%]	1.07 [34.22%]
Over-identification Sargan test	6.13 [18.9%]	5.53 [23.65%]	3.71 [44.6%]
Tax records variables	No	Yes	Yes
LFS variables	No	No	Yes
N° of Observations	12,512	12,512	12,512

Table 2 – **Estimates of the elasticities with respect to the net-of-tax rates**

Notes: standard errors are in parentheses and p-values in brackets. * denotes significance at 10%, ** significance at 5% and *** significance at 1%. Estimation by 2SLS using instruments I and II. All regressions include time dummies and a fifth-order polynomial of the log of base year gross wages.
Sample: employees present two consecutive years.
Source: ERF survey, Insee, 2003-2006.

Conversely, our estimate for the effect of marginal net-of-payroll-tax rates on gross wages β_τ^P is close to zero and not significant. This finding is obtained whatever the set of controls is. This finding is in line with Saez, Matsaganis and Tsakloglu (2010) for Greece, and with Aeberhardt and Sraer (2009) for France. However, it differs from Lhommeau and Remy (2009), also for France, who find that the progressivity of payroll taxes has a negative effect on wage growth. Our result that gross wages do not respond to marginal payroll tax rates suggests that, at least in the short run, the efficiency costs of financing social security expenses and redistribution through payroll taxes are lower than through income taxes.

We now turn to income effects. The elasticity with respect to the average income net-of tax rate is negative and significant (only at the 10% level in column 3). This contrasts with the literature which usually does not find significant income effects (e.g. Gruber and Saez (2002)). We also find that the elasticity with respect to the average net-of-payroll-tax rate is negative and significant. The parameter varies between -1.4 and -1.7, and are reinforced when adding LFS controls. However, we cannot reject that it is equal to -1, which suggests that wages are negotiated net-of employer payroll taxes. A decrease in employer payroll taxes seems almost entirely absorbed by employers and thus actually reduces the labor cost, without any significant effect on taxable wages.

Our result that gross wages are insensitive to marginal payroll tax rates, whereas they respond to income tax rates has important implications. Section II shows that a large class of theoretical models of the labor market predicts identical elasticities, as expressed by (9). This class includes the textbook labor supply model where the hourly gross wage equals the marginal productivity of labor. According to F-tests, the evidence for France is that Equation (9) is strongly rejected (at the 1% level for model (1) and at the 5% level for models (2) and (3)). This therefore suggests that the textbook "competitive – labor supply" model that underlies most of public finance analysis (e.g. Mirrlees (1971)) may be inadequate.

Moreover, our estimates also lead us to reject prediction (9') since gross wages respond more to marginal income than to marginal net-of-payroll-tax rates, but less to average income than to average net-of-payroll-tax rates. As argued in Section II, our results are not supportive of the idea that people understand differently the consequences of income taxes and payroll taxes. Our findings also reject models where payroll taxes generate benefits that are internalized in the formation of the gross wages.

We also test Equation (9'') which is obtained under the assumptions that individuals choose their labor supply and that hourly taxable wages are rigid. These assumptions imply in particular that the hourly gross wage differs from the marginal productivity of labor, a key departure from a perfectly competitive labor market. The F-tests indicate that (9'') is very far from being rejected by the data. Plausibly, the hourly taxable wage is determined through bargaining, which does not occur every year. This implies that the responses of gross wages to reforms that took place between years *t*-1 and *t* are delayed, and thus are not reflected in our estimates of β_τ^P.

VI.2 Heterogeneous effects

In this Section, we are interested in the effects of taxes on gross wages across subgroups. This will also help us to understand the economic behaviors behind our results. First, as mentioned in section V, those working 35 hours a week and those working 39 hours a week have been subjected to very different payroll tax changes. In order to check the robustness of our main result that wages respond differently to income taxes and to payroll taxes, we run separate analyses for the two types of employees. The results are reported in Table 3. Because of the subsamples' size, we keep a subset of the LFS and tax records control variables included in column 3 of Table 2.[17] This is the reason why the results reported in the first column of Table 3 for the whole sample differ very slightly compared to those reported in column 3 of Table 2.

	Whole sample (1)	35-hour workweek (2)	39-hour workweek (3)
β_τ^I	0.194** (0.081)	0.183* (0.099)	0.176 (0.128)
β_τ^P	-0.049 (0.092)	-0.019 (0.134)	-0.010 (0.161)
β_ρ^I	-1.068*** (0.341)	-0.825 (0.625)	-0.557 (0.376)
β_ρ^P	-1.458*** (0.267)	-2.07*** (0.678)	-1.751** (0.710)
$(9): \quad \beta_\tau^I = \beta_\tau^P \quad \text{and} \quad \beta_\rho^I = \beta_\rho^P$	3.77** [2.3%]	4.77 [0.8%]	1.05 [34.9%]
$(9''): \quad \beta_\tau^P = 0 \quad \text{and} \quad \beta_\rho^P = -1$	1.57 [20.7%]	2.05 [12.9%]	0.69 [50.0%]
Over-identification Sargan test	3.80 [28.4%]	3.0 [22.2%]	2.91 [40.6%]
Tax records variables	Yes	Yes	Yes
LFS variables	Yes	Yes	Yes
N° of Observations	12,512	9,509	3,003

Table 3 - **Elasticities for employees working 35-hour and 39-hour weeks**

Notes: standard errors are in parentheses and p-values in brackets. * denotes significance at 10%, ** significance at 5%, and *** significance at 1%. Estimation by 2SLS using instruments I and II.
Sample: employees present two consecutive years.
Source: ERF survey, Insee, 2003-2006.

The main results found for the whole population remains qualitatively unchanged, implying that Equations (9) are rejected while Equation (9'') is consistent with the data. However, the small size of the 39-hour subsample (column 3) makes the estimates more imprecise. For instance, although the

[17] We take a third-order polynomial of base year gross wages instead of a fifth-order one. We drop the indicators of household' type, industry, firm size, and the interaction between firm working time and the year , which are no longer significant.

estimate for β_τ^I on the 39-hour subsample is very close to that estimated on the 35-hour subsample (column 2), it is statistically not significant. Consequently, Equation (9) can no longer be rejected for employees working 39-hour weeks; however, Equation (9'') is much more easily accepted than Equation (9).

We next present the results for females and males separately in Table 4. In both subsamples, the marginal net-of-payroll-tax rate does not significantly affect gross wages, and the elasticity of gross wages with respect to the average net-of-payroll-tax rate is negative, significant, and does not significantly differ from -1. Consequently, the F-test of Equation (9'') on both subsamples does not reject the nul hypothesis that β_τ^P=0 and β_ρ^P=-1. The responses to income taxation are however very different for females and for males. The elasticity β_τ^I of gross wages with respect to the marginal net-of-income-tax rate is significantly positive for females and much higher than on the whole sample. Conversely, for males, gross wages do not seem to respond to changes in income taxation. That males and females exhibit different responses to income taxation suggests that the responses to income taxation highlighted in Table 2 for the whole sample might be the result of labor supply decisions, which are notoriously much more important for females than for males. If these responses were due to wage negotiation effects, then behavioral responses of males and females would very likely be similar.

	Whole sample (1)	Females (2)	Males (3)
β_τ^I	0.194** (0.081)	0.879*** (0.234)	-0.024 (0.113)
β_τ^P	-0.049 (0.092)	0.120 (0.176)	-0.108 (0.111)
β_ρ^I	-1.068*** (0.341)	-0.560 (0.530)	-0.791 (2.037)
β_ρ^P	-1.458*** (0.267)	-1.499*** (0.525)	-1.250*** (0.342)
$(9): \beta_\tau^I = \beta_\tau^P$ and $\beta_\rho^I = \beta_\rho^P$	3.77** [2.3%]	8.83*** [0.0%]	0.32 [72.7%]
$(9''): \beta_\tau^P = 0$ and $\beta_\rho^P = -1$	1.57 [20.7%]	0.54 [58.14%]	0.64 [52.8%]
Over-identification Sargan test	3.80 [28.4%]	1.83 [60.7%]	0.31 [57.7%]
Tax records variables	Yes	Yes	Yes
LFS variables	Yes	Yes	Yes
N° of Observations	12,512	5,266	7,246

Table 4 - **Elasticities for females and males**

Notes: standard errors are in parentheses and p-values in brackets. * denotes significance at 10%, ** significance at 5%, and *** significance at 1%. Estimation by 2SLS using instruments I and II.
Sample: employees present two consecutive years.
Source: ERF survey, Insee, 2003-2006.

In order to better interpret the mechanisms underlying our findings, Table 5 displays the results on different subsamples more particularly affected by the reforms. Column 1 reports the results over the whole sample. In column 2, we keep the employees whose earnings at *t*-1 are lower than 2.2 times the earnings of an individual working full-time and for the full-year at the minimum wage. As explained in Section V, the most important reforms that took place over the period 2003-2006 concern those individuals. As expected, the responses to marginal and average net-of-income-tax rates are reinforced. However, the response to marginal net-of-payroll-tax rates remains close to zero and insignificant. Equation (9) is now rejected at the 1% level, while equation (9'') is more easily accepted. This strengthens our interpretation that the hourly taxable wage is rigid, while individuals respond to change in income taxation through labor supply.

	Whole sample (1)	< 2.2 times the Minimum wage at *t*-1 (2)	(2) & employed full-year at *t*-1 (3)	(3) & employed full-year at *t* (4)
β_τ^I	0.194** (0.081)	0.467*** (0.148)	0.275** (0.131)	0.033 (0.122)
β_τ^P	-0.049 (0.092)	-0.003 (0.095)	-0.048 (0.081)	-0.073 (0.073)
β_ρ^I	-1.068*** (0.341)	-1.581* (0.938)	-0.788 (0.803)	-0.516 (0.734)
β_ρ^P	-1.458*** (0.267)	-1.524*** (0.312)	-1.229*** (0.277)	-1.323*** (0.248)
$(9): \quad \beta_\tau^I = \beta_\tau^P \quad \text{and} \quad \beta_\rho^I = \beta_\rho^P$	3.77** [2.3%]	5.67 [0.4%]	5.05 [0.7%]	2.44* [8.7%]
$(9''): \quad \beta_\tau^P = 0 \quad \text{and} \quad \beta_\rho^P = -1$	1.57 [20.7%]	1.41 [24.4%]	0.50 [60.9%]	1.35 [25.9%]
Over-identification Sargan test	3.80 [28.4%]	4.29 [36.8%]	1.31 [72.6%]	0.95 [81.2%]
Tax records variables	Yes	Yes	Yes	Yes
LFS variables	Yes	Yes	Yes	Yes
N° of Observations	12,512	9,979	9,320	9,200

Table 5 - **Elasticities for specific subsamples in the first middle of the wage distribution**

Notes: standard errors are in parentheses and p-values in brackets. * denotes significance at 10%, ** significance at 5%, and *** significance at 1%. Estimation by 2SLS using instruments I and II.
Sample: employees present two consecutive years.
Source: ERF survey, Insee, 2003-2006.

The literature surveyed by Saez, Slemrod and Giertz (2010) is concerned with potential biases due to reversion to the mean. When using reforms concentrated on top-income earners, one has to care that getting a very high income at *t*-1 may be accidental, thereby leading to a negative change in earnings between *t* and *t*-1. A symmetrical issue can occur at the bottom of the income distribution when using reforms concentrated on bottom-income earners, which is our case. An individual not working full-year at *t*-1 (for example a young worker entering the labor market) is more likely to be employed full-year at *t*, thereby leading to a rise in the gross wage that should not be attributed to tax changes. Figure 3 suggests that the reversion-to-the-mean phenomenon is very important among bottom-wage earners. The peak in the change in gross wages at the bottom of the distribution becomes much lower when the sample is restricted to those employed full-year. Therefore, in order to verify that our results are not due to reversion to the mean at the bottom of the distribution, we restrict the subsample used in column 2 to those employed 12 months in year *t*-1. The results displayed in column 3 indicate that the elasticity with respect to changes in the marginal net-of-payroll tax rate is unaffected, while the elasticity with respect to changes in the marginal net-of-income tax rate is reduced by a half, while remaining highly significant.

Interestingly, the reponse to changes in average net-of-income tax rates becomes insignificant. This suggests that an important channel through which employees respond to income taxation is the choice of the number of working months by year, i.e. an extensive or participation margin. In order to test this hypothesis we drop from the subsample used in column 3 those not employed full-year at *t*. Hence column 4 corresponds to the results for the subsample of column 2 of those employed full-year at *t* and *t*-1. Again, the elasticities with respect to marginal and average net-of-payroll-tax rates are unaffected and in line with restriction (9''). The novelty is that the elasticity with respect to the marginal income-net-of tax rate becomes very close to zero and not significant. This confirms our presumption that our responses to changes in income taxation essentially reflect participation decisions of individuals, rather than hours-of-work intensive decisions of those remaining employed.

VII. Conclusion

In this paper, we estimate the behavioral responses of the gross labor income to marginal and average tax rates of both the income and the payroll tax schedules in France over the period 2003-2006. We use the changes in the schedules of reduction of employers' social security contribution for low paid jobs to identify the responses to the payroll tax schedules. As for identifying the responses to the income tax schedule, we use the increase in income tax credit for low wage earners that took place over the same period. We find a significant elasticity of the gross labor income with respect to the marginal net-of-income-tax rate around 0.2 and our results suggest that this effect is driven by females labor supply decisions. Conversely, we find no significant effect of marginal net-of-payroll-tax rate. This discrepancy appears robust to different specifications and selections. It is in contradiction with a prediction that is common to a large class of models of the labor market, in particular the competitive labor supply framework, which plays a central role in the theoretical literature of optimal income taxation.

We also find a significant elasticity of gross labor income to the average payroll net-of tax rate, which is not significantly different than minus one. Conversely, the elasticity with respect to the average net-of-income-tax rate is much weaker and generally not significant. Among the different theories that can account for different behavioral responses to payroll and income taxation, the most plausible one is based on the rigidity of the hourly taxable wage, together with labor supply responses to the income tax schedule.

This suggests that the payroll tax schedule may be less distortive than income tax schedule, at least over the short horizon of our study. Another implication is that reduction of employers' social security contributions that play a very important role in France seems to benefit exclusively to the employer, so presumably to the labor demand of low skilled jobs. Moreover, our estimates does not support the view that these programs exacerbated a low-wage trap.

This work can be extended in different directions. A first direction would be to consider a longer panel of observations to investigate the long run responses to taxation. This in particular would enable us to test whether the irresponsiveness of taxable wages to payroll taxation is only a short run result or whether the responses of gross wages to payroll taxation in the long run are similar to the responses to income taxation. Another extension is to disentangle the responses we obtain in terms of wage formation, labor demand effects, participations decisions effects and intensive labor supply effects.

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Appendix A

A.1) Proof of Equation (6)

Combining Equations (4) and (5) leads to:

$$\frac{\Delta w}{w^*} = \beta_\tau^P \frac{\Delta \tau^P}{\tau^P} + \beta_\tau^I \frac{\Delta \tau^I}{\tau^I} + \frac{\partial W}{\partial R^P}\left[\Delta \tau^P + \frac{\Delta R^P}{w^*}\right] + \frac{\partial W}{\partial R^I}\left[\rho^P \Delta \tau^I + \frac{\Delta R^I}{w^*}\right]$$

Using $\Delta\rho^P = \Delta\tau^P + (\Delta R^P / w^*)$ and $\Delta\rho^I + (R^I/z^*)(\Delta\rho^P/\rho^P) = \Delta\tau^I + (\Delta R^I / z^*)$ gives:

$$\frac{\Delta w}{w^*} = \beta_\tau^P \frac{\Delta \tau^P}{\tau^P} + \beta_\tau^I \frac{\Delta \tau^I}{\tau^I} + \frac{\partial W}{\partial R^P}\Delta\rho^P + \frac{\partial W}{\partial R^I}\rho^P\left[\Delta\rho^I + \frac{R^I}{z^*}\frac{\Delta\rho^P}{\rho^P}\right]$$

Rearranging terms using Equation (7) gives Equation (6).

A.2) Proof of Equations (5'), (6') and (7')

Differentiating both sides of (8) gives:

$$\frac{\partial W}{\partial \tau^I} = \tau^P \frac{\partial \Omega}{\partial \tau} + R^P \frac{\partial \Omega}{\partial R} \qquad \frac{\partial W}{\partial \tau^P} = \tau^I \frac{\partial \Omega}{\partial \tau}$$

$$\frac{\partial W}{\partial R^I} = \frac{\partial \Omega}{\partial R} \qquad \frac{\partial W}{\partial R^P} = \tau^I \frac{\partial \Omega}{\partial R}$$

Then, using (5) and (5') gives:

$$\beta_\tau^P = \left(\frac{\tau^P}{w} \frac{\partial W}{\partial \tau^P} \right) - \tau^P \frac{\partial W}{\partial R^P} = \left(\frac{\tau^P \tau^I}{w} \frac{\partial \Omega}{\partial \tau} \right) - \tau^P \tau^I \frac{\partial \Omega}{\partial R} = \left(\frac{\tau}{w} \frac{\partial \Omega}{\partial \tau} \right) - \tau \frac{\partial \Omega}{\partial R} = \beta_\tau$$

$$\beta_\tau^I = \left(\frac{\tau^I}{w} \frac{\partial W}{\partial \tau^I} \right) - \tau^I \rho^P \frac{\partial W}{\partial R^I} = \frac{\tau^P \tau^I}{w} \frac{\partial \Omega}{\partial \tau} + \frac{\tau^I}{w} R^P \frac{\partial \Omega}{\partial R} - \tau^I \rho^P \frac{\partial \Omega}{\partial R}$$

$$= \left(\frac{\tau}{w} \frac{\partial \Omega}{\partial \tau} \right) - \tau^I \left(\tau^P + \frac{R^P}{w} - \frac{R^P}{w} \right) \frac{\partial \Omega}{\partial R} = \left(\frac{\tau}{w} \frac{\partial \Omega}{\partial \tau} \right) - \tau \frac{\partial \Omega}{\partial R} = \beta_\tau$$

Using (7) and (7') gives:

$$\beta_\rho^P = \rho^P \frac{\partial W}{\partial R^P} + \rho^P \frac{R^I}{z^*} \frac{\partial W}{\partial R^I} = \rho^P \left(\tau^I + \frac{R^I}{z^*} \right) \frac{\partial \Omega}{\partial R} = \rho^P \rho^I \frac{\partial \Omega}{\partial R} = \rho \frac{\partial \Omega}{\partial R} = \beta_\rho$$

$$\beta_\rho^I = \rho^P \rho^I \frac{\partial W}{\partial R^I} = \rho \frac{\partial \Omega}{\partial R} = \beta_\rho$$

A.3) Sign of β_τ in Equation (9)

The first-order condition of the maximization of $U(\tau w + R, w)$ writes: $F(w,\tau,R)=0$, where function $F(.,.,.)$ is defined by: $F(w,\tau,R) \equiv \tau \cdot U_1'(\tau\, w + r, w) + U_2'(\tau\, w + r, w)$. Assuming that the second-order condition $F_w' < 0$ holds with a strict inequality (which is the case if for instance U is strictly concave), the implicit function theorem enables us to compute the partial derivatives of function $\Omega(.,.)$ at $w^*=\Omega(\tau,R)$ through:

$$\Omega_\tau' = -\frac{F_\tau'}{F_w'} = -\frac{U_1' + w^*\left(\tau \cdot U_{11}'' + U_{12}''\right)}{F_w'} \quad \text{and} \quad \Omega_R' = -\frac{F_R'}{F_w'} = -\frac{\tau \cdot U_{11}'' + U_{12}''}{F_w'}$$

where the partial derivatives of U are computed at $w = w^*$ and $c=\tau w^*+R$. Applying (5') gives:

$$\beta_\tau = \frac{\tau}{w^*}\left(\Omega_\tau' - w^* \Omega'_R\right) = -\frac{\tau\left(F_\tau' - w^* F_R'\right)}{w \cdot F_w'} = -\frac{\tau}{w \cdot F_w'} \cdot U_1'$$

$\beta_\tau>0$ then follows the assumption that the objective U is increasing in c, i.e. $U_1'>0$.

A.4) Proof of (9')

Let us write $F(w,\tau^P,\tau^I,R^P,R^I)=0$ the first-order condition of the maximization of $U\left(\tau^P\tau^I w+\tau^I R^P+R^I+k\left(\left(1-\tau^P\right)w-R^P\right),w\right)$, where function $F(.,.,.)$ is now defined by:

$$
\begin{aligned}
F\left(w,\tau^P,\tau^I,R^P,R^I\right) &\equiv \left(\tau^P\cdot\tau^I+k\left(1-\tau^P\right)\right)\cdot U_1^{'}\left(\tau^P\tau^I w+\tau^I R^P+R^I+k\left(\left(1-\tau^P\right)w-R^P\right),w\right) \\
&+U_2^{'}\left(\tau^P\tau^I w+\tau^I R^P+R^I+k\left(\left(1-\tau^P\right)w-R^P\right),w\right)
\end{aligned}.
$$

Assuming that the second-order condition $F_w^{'}<0$ holds with a strict inequality (which is the case if for instance U is strictly concave), the implicit function theorem enables us to compute the partial derivatives of function $W(.,.,.,.)$ at $w^*=W(\tau^P,\tau^I,R^P,R^I)$ through:

$$
\begin{aligned}
W_{\tau^P}^{'} &= -\frac{F_{\tau^P}^{'}}{F_w^{'}} = -\frac{\left(\tau^I-k\right)\cdot U_1^{'}+\left(\tau^I-k\right)\cdot w^*\cdot\left(\left(\tau^P\cdot\tau^I+k\left(1-\tau^P\right)\right)\cdot U_{11}^{''}+U_{12}^{''}\right)}{F_w^{'}} \\
W_{\tau^I}^{'} &== -\frac{F_{\tau^I}^{'}}{F_w^{'}} = -\frac{\tau^P\cdot U_1^{'}+\left(\tau^P\cdot w^*+R^P\right)\cdot\left(\left(\tau^P\cdot\tau^I+k\left(1-\tau^P\right)\right)\cdot U_{11}^{''}+U_{12}^{''}\right)}{F_w^{'}} \\
W_{R^P}^{'} &= -\frac{F_{R^P}^{'}}{F_w^{'}} = -\frac{\left(\tau^I-k\right)\cdot\left(\left(\tau^P\cdot\tau^I+k\left(1-\tau^P\right)\right)\cdot U_{11}^{''}+U_{12}^{''}\right)}{F_w^{'}} \\
W_{R^I}^{'} &= -\frac{F_{R^I}^{'}}{F_w^{'}} = -\frac{\left(\tau^P\cdot\tau^I+k\left(1-\tau^P\right)\right)\cdot U_{11}^{''}+U_{12}^{''}}{F_w^{'}}
\end{aligned}
$$

where the various partial derivatives are computed at w = w^* and $c=\tau w^*+R$. Applying (5) while taking $z^*=\tau^P w^*+R^P$ into account leads to:

$$
\begin{aligned}
\beta_{\tau^I} &= \frac{\tau^I}{w^*}\cdot\left(W_{\tau^I}^{'}-z^*\cdot W_{R^I}^{'}\right) = -\frac{\tau^I\left(F_{\tau^I}^{'}-z^*\cdot F_{R^I}^{'}\right)}{w\cdot F_w^{'}} = -\frac{\tau^P\cdot\tau^I}{w\cdot F_w^{'}}\cdot U_1^{'}>0 \\
\beta_{\tau^P} &= \frac{\tau^P}{w^*}\cdot\left(W_{\tau^P}^{'}-w^*\cdot W_{R^P}^{'}\right) = -\frac{\tau^P\left(F_{\tau^P}^{'}-w^* F_{R^P}^{'}\right)}{w\cdot F_w^{'}} = -\frac{\tau^P\left(\tau^I-k\right)}{w\cdot F_w^{'}}\cdot U_1^{'} = \beta_{\tau^I}\cdot\left(1-\frac{k}{\tau^I}\right)
\end{aligned}
$$

where the inequalities follow the second-order conditions $F_w^{'}<0$ and $U_1^{'}>0$. Applying (7) then gives:

$$\beta_\rho^I = \rho^P \cdot \rho^I \cdot W_{R^I}^{'} = -\frac{\rho^P \cdot \rho^I \cdot \left(\left(\tau^P \cdot \tau^I + k\left(1-\tau^P\right)\right) \cdot U_{11}^{"} + U_{12}^{"}\right)}{F_w^{'}}$$

$$\beta_\rho^P = \rho^P \cdot \left[W_{R^P}^{'} + \frac{R^I}{z^*} W_{R^I}^{'}\right] = -\frac{\rho^P \cdot \left(\tau^I - k + \frac{R^I}{z^*}\right) \cdot \left(\left(\tau^P \cdot \tau^I + k\left(1-\tau^P\right)\right) \cdot U_{11}^{"} + U_{12}^{"}\right)}{F_w^{'}}$$

$$= -\frac{\rho^P \cdot \left(\rho^I - k\right) \cdot \left(\left(\tau^P \cdot \tau^I + k\left(1-\tau^P\right)\right) \cdot U_{11}^{"} + U_{12}^{"}\right)}{F_w^{'}} = \beta_\rho^I \cdot \left(1 - \frac{k}{\rho^I}\right)$$

Therefore we get (9') as $k < \tau^I$ and $k < \rho^I$.

Appendix B: additional empirical results

Table B.1 reproduces the results of the first stage equations, using the specification and method that correspond to the results in Table 2, column 3. Unlike for the other results, we report T statistics in brackets instead of standard deviation. In bold are estimates of the direct effects of $\Delta \log \overline{\varphi}_{i,t}^{j}$ and $\Delta \log \overline{\overline{\varphi}}_{i,t}^{j}$ on $\Delta \log \varphi_{i,t}^{j}$:

	$\Delta\log\tau_{i,t}^I$	$\Delta\log\tau_{i,t}^P$	$\Delta\log\rho_{i,t}^I$	$\Delta\log\rho_{i,t}^P$
$\Delta\log\overline{\tau}_{i,t}^I$	**0.445***** **(25.74)**	-0.092*** (-6.31)	0.022*** (4.29)	0.002 (0.36)
$\Delta\log\overline{\tau}_{i,t}^P$	-0.031 (-0.97)	**0.517***** **(19.38)**	0.009 (0.96)	-0.003 (2.99)
$\Delta\log\overline{\overline{\tau}}_{i,t}^I$	**0.026*** **(1.70)**	-0.008 (-0.65)	0.006 (1.32)	-0.003 (-0.71)
$\Delta\log\overline{\overline{\tau}}_{i,t}^P$	0.039 (1.19)	**0.152***** **(5.46)**	-0.020** (-2.03)	-0.010 (-1.05)
$\Delta\log\overline{\rho}_{i,t}^I$	0.304 (2.05)	-0.009 (-0.07)	**0.182***** **(4.20)**	-0.292*** (-6.14)
$\Delta\log\overline{\rho}_{i,t}^P$	-0.447 (-1.89)	-0.995*** (-5.01)	-0.009 (-0.14)	**0.263***** **(3.79)**
$\Delta\log\overline{\overline{\rho}}_{i,t}^I$	-0.079 (-0.67)	0.042 (0.43)	**0.352***** **(10.19)**	0.124*** (3.60)
$\Delta\log\overline{\overline{\rho}}_{i,t}^P$	0.468* (1.84)	0.623*** (2.92)	-0.073 (-0.98)	**0.365***** **(4.89)**
Tax records variables	Yes	Yes	Yes	Yes
LFS variables	Yes	Yes	Yes	Yes
No. of individuals	12,512	12,512	12,512	12,512
F-Statistic	18.26***	17.80***	26.15***	16.97***

Table B.1 - **First-stage regressions**
(Student test in brackets)

Table B.2 displays the complete estimates presented in Table 2.

	Parameter Estimate	Standard Error	Parameter Estimate	Standard Error	Parameter Estimate	Standard Error
Constant	-148.177	62.4073	-162.267	62.77584	-160.196	64.47044
β_{τ}^{I}	0.239967	0.084666	0.237658	0.08603	0.225061	0.085914
β_{τ}^{P}	-0.00995	0.094904	0.001442	0.096234	0.015545	0.096
β_{ρ}^{I}	-0.72796	0.328761	-0.78578	0.385905	-0.69477	0.395067
β_{ρ}^{P}	-1.4372	0.288757	-1.38961	0.292004	-1.74857	0.511287
2003-2004	-0.02522	0.007704	-0.02636	0.008335	-0.02945	0.014526
2004-2005	-0.01504	0.007953	-0.01643	0.008759	-0.02147	0.013964
Terms of the 5-th order polynomial in log $w_{i,t-1}$	78.68947	31.297	85.71767	31.51517	84.57709	32.27043
	-16.1052	6.202628	-17.481	6.2519	-17.2832	6.382864
	1.603197	0.607763	1.735336	0.613139	1.723013	0.624238
	-0.07806	0.029459	-0.08429	0.029745	-0.08418	0.030204
	0.001493	0.000565	0.001609	0.000571	0.001617	0.000579
≤ 29 years			0.026723	0.008623	0.008019	0.00945
30 - 39 years			0.01576	0.006287	0.009712	0.006543
50 - 59 years			-0.02746	0.006974	-0.02198	0.007094
≥ 60 years			-0.11095	0.028075	-0.09784	0.028152
Women			-0.00327	0.006599	0.004896	0.006917
Women with a new child since t-1			-0.07767	0.021859	-0.07334	0.021707
New child since t-1			0.017167	0.015011	0.013593	0.014941
Women with a child exiting			-0.02671	0.019924	-0.02526	0.019849
Exit of a child since t-1			0.009469	0.012893	0.008312	0.01282
Women and child under 18 months			0.005188	0.018167	0.007338	0.018401
Women and child under 3 years old			0.06572	0.017898	0.05849	0.018216
Women and child under 6 years old			-0.00849	0.012336	-0.00908	0.012302
Women and child under 18 years old			0.011144	0.008579	0.010866	0.008592
Single individual			-0.01328	0.008408	-0.01377	0.008445
Single parent			0.008358	0.010016	0.006936	0.009985
Couple with children			-0.00037	0.007284	-0.00174	0.007252
"Complex" household			0.011143	0.014617	0.015207	0.014549
College (> 2 years)					0.034023	0.012716
College (≤ 2 years)					0.023488	0.010968
High school graduate					0.023217	0.010638
High-school drop-out or vocational diploma					0.021058	0.009026
Junior high school or basic vocational					0.016832	0.011743
Manufacturing					0.004314	0.006169
Agriculture					-0.02141	0.020218
Construction					0.001379	0.010221
Energy					-0.00616	0.019172
Education and social activities					-0.01762	0.009283
Trade and repair					-0.00133	0.007194
Engineers, managers and professionals					-0.00766	0.009623
< 10 employees					-0.00552	0.007982
10-19 employees					-0.00208	0.009801
35-hour and 2003-2004					-0.00924	0.013842
35-hour and 2004-2005					-0.01127	0.014541
35-hour and 2005-2006					-0.01906	0.010895
< 1 year					-0.06032	0.023244
1-5 years					-0.16661	0.021435
5 - 10 years					-0.16861	0.022446
≥ 10 years					-0.16409	0.02254

Table B.2 - **Complete results of the model 1 inTable 3**