Do Firms Really Share Rents With Their Workers?

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Abstract

We use matched firm - worker panel data from France and Norway to consider observationally equivalent alternatives to the hypothesis that firms share product market rents with their workers in the form of higher wages. After documenting the main stylized fact, we find that neither the main statistical explanations (group effects in residuals and measurement error) nor sectoral shocks seem to be responsible for the observed correlation. Statistical-economic explanations (endogeneity of profits, omitted variable biases in terms of individual productive characteristics) are slightly more successful, as instrumentation reduces the significance level in France to 89% (via an increase in the stardard error of the estimate). The most complete model, with unobserved heterogeneity in both time-invariant firm compensation policy and time-invariant individual characteristics, instrumental variables and a complete set of controls for worker observables and sectoral shocks renders the coefficient insignificant for France and weakens its significance for Norway. This result is coherent with the multi-level bargaining structure found in Norway and the presence of a more mobile labor force in France, although it may also be due to insufficient degrees of freedom.

Resumé

Nous utilisons des données longitudinales appareillées employeur-employé pour évaluer des explications alternatives à l'hypothèse que les entreprises avec plus de rentes dans leur marché de produit les partagent avec leurs employés sous forme d'un salaire plus élevé. Après avoir démontré le fait stylisé nous trouvons que ni les explications statistiques principales (erreur de mesure et effets de groupe) ni les chocs sectoriels ne semblent êtres les responsables de cette corrélation. Les alternatives statistique-économique (endogénéité des profits, biais de variables omises pour les caractéristiques productives individuelles) ont légèrement plus de succès, dans la mesure que l'instrumentation réduit la significativité des profits par travailleur en France à 89% (via une augmentation de l'écart type). Le modèle le plus complet, avec de l'hétérogénéité inobservée du côté de l'employeur et de l'employé, variables instrumentales, et un ensemble complet de variables de contrôle pour les observables individuelles et chocs sectoriels rend le coefficient non-significatif pour la France et réduit la significativité pour la Norvège. Ce résultat est cohérent avec les négociations collectives à plusieurs niveaux en Norvège et la présence d'une force de main-d'oeuvre plus mobile en France, mais une explication en termes de manque de degrés de liberté n'est pas à exclure.

Key Words: Rent Sharing, Matched Employer-Employee Panel Data, International Comparisons.
Mots clés: Partage des rentes, données longitudinales appareillées employeur-employé, comparaisons internationales.
JEL Codes / Codes JEL: J31, C23

1 Introduction

The recent development of matched employer-employee data has opened enormous possibilities in terms of the ability to test previously-existing economic theories.[1] One of the theories that seems the most ripe for investigation, and for which matched firm-worker panel data seems ideally suited, is the theory of rent sharing in the labor market.[2] According to this theory, firms that earn positive economic rents on the product market share them with their employees in relation to the parties' relative bargaining powers. As a result, workers employed by the most profitable firms, ceteris paribus, will tend to earn more than those in less profitable firms and, in particular, fluctuations in profits over time for a firm should be reflected in the earnings of workers that the firm employs.

Using matched firm-worker panel data from the manufacturing sectors of France and Norway, we find that there is indeed a positive, significant relation between profits per worker and log annual earnings, measured at the individual worker level.[3] Nevertheless, there are many other possible explanations, besides rent sharing, that could explain the significant coefficient. The main objective of this paper is to consider each of these alternative explanations sequentially, and to see if any of them is capable of reducing the coefficient on profits per worker to insignificance.

The structure of this paper is as follows. After providing a rapid discussion of previous research designed to address the rent sharing model in section 2, we begin in section 3 by presenting a simple model of bargaining which underlies much of the literature on rent sharing. We then present the main (stylized) fact; real profits per worker in the manufacturing sectors of France and Norway are significantly, and positively, related to real full-year equivalent gross earnings for full-time employed men in an ordinary least squares regression.

Our first concern is to rule out the bias in the estimated standard errors of the coefficients due to the estimation being carried out at the individual level while profits are measured at the firm level (and we have multiple individuals per firm). This question, addressed in detail by Greenwald (1983) and demonstrated dramatically by Moulton (1990) , is considered in section 4. Section 4 also discusses the impact that measurement error, particularly mean reversion due to tax incentives to defer profits and spread losses across several years, might have on our estimates.

Section 5 then considers three alternative explanations, namely the presence of inter-industry wage differentials that are unrelated to idiosyncratic profitability variation, business

[1]See Abowd and Kramarz (1999) for a recent review of this literature.

[2]Oswald (1996) repeatedly mentions the potential for matched firm-worker panel data to resolve existing issues in the rent sharing literature.

[3]We conduct our analysis at the individual worker level, rather than at the firm level, since this allows us more flexibility in the sorts of heterogeneity for which we can control, most notably in terms of unobservable (to the econometrician) individual productive characteristics.

cycle effects and sector-specific shocks. All of these explanations imply that the source of the link between profits and earnings is not firm, but rather industry or year, specific. We next address the question of observable differences in individual characteristics in section 6. It is possible that certain firms are more profitable because they simply have more skilled workers, and these workers are only compensated for the market value of their extra skills. Conducting our analysis at the individual level, as opposed to the firm level, allows us to exploit variation in workforce composition over time, that may be related to profitability, in a much more detailed manner.[4]

After excluding the standard sources of omitted variable bias described above, we then address the question of the endogeneity of profits in section 7. We do this through instrumental variables estimation. Next, section 8 considers that possibility that different firms pay observably equivalent workers differently, but in some fixed manner. This would be the case, for example, if different firms employed different incentive compensation policies. This approach also allows us, in the case where rent sharing is present, to capture time-invariant unobserved heterogeneity in firm bargaining power. We finish in section 9 by considering the most general possible specification, which allows for the possibility that variations in observable and unobservable individual and firm heterogeneity that may be correlated with both profits and earnings are driving the observed relation. Section 10 summarizes our findings and concludes. Descriptive statistics for the data used are given in appendix table 1, and data construction is considered in the data appendix.

We find that we can reject almost all observationally equivalent explanations for rent sharing. However, the endogeneity of profits per worker in France seems to be a significant issue, in that the instrumental variables estimator becomes insignificant. This is mainly due to an increase in the standard error of the estimate (and a problem of weak instruments), however. The only observationally equivalent explanation that might be able to supplant rent sharing is that of unobserved fixed worker heterogeneity; although this too achieves the feat of reducing the significance of the coefficient on profits mainly through an increase in the standard error of the estimate, rather than a reduction in the point estimate towards zero in the most complete specification. The increase in the standard error is sufficient to make the coefficient on profits insignificant in France, and to reduce its significance below the 5 percent level in Norway (although it remains significant at the 10 percent level). Further examination of the relative impact of instrumentation versus individual fixed effets shows that instrumentation obscures the impact of unobserved individual heterogeneity in France, and thus worker sorting by unobserved individual characteristics may indeed be an important alternative explanation for the positive correlation between profits and wages.

[4]This could also be done by weighted estimation and using within-firm averages of observable individual characteristics, but such an approach would not be sufficient when we correct for unobserved individual heterogeneity below. As a result, we prefer using the same level of aggregation throughout.

2 Previous Research on Rent Sharing

The standard neoclassical microeconomic model of the labor market implies that, under perfect competition, firms equate the value of marginal product of labor (productivity) with its marginal cost (wages). However, it is still controversial whether other factors, such as industry or firm characteristics, influence wages above compensating differentials and omitted unobserved worker characteristics. If there exist persistent differentials in wages across industries or firms, this may be a sign that labor markets are non-competitive, and economic policies directed at the labor market should be adapted accordingly.

A growing literature, dating back to Slichter (1950) , appears to find persistent wage differences for observationally equivalent workers across different industries or different firms. The differences also appear to be stable over countries with quite different wage setting institutions in terms of degree of union coverage, degree of centralization of wage bargaining etc... In the spirit of Slichter (1950) , a series of papers in the late 1980s found strong inter-industry wage differentials for the US even when observed worker characteristics are included (Dickens and Katz (1987) , Krueger and Summers (1988) , Murphy and Topel (1987)). In line with rent-sharing theories, some of these studies also show that the wage differentials are correlated with the industries' profitability. The interpretation of this result, however, tends to vary by author.

One line of argument suggests that inter-firm or inter-industry wage differentials only reflect unobserved worker characteristics or unobserved job characteristics. This is a favored explanation, since much of the evidence works against both compensating differentials and omitted observed characteristics being the source of variations. The fact that some studies find uniformity of wage differentials across occupations is hard to reconcile with either explanation. The compensating differentials hypothesis also has difficulties in explaining differences in quit rates over industries, in that high wage industries tend to have lower quit rates (Katz and Summers 1989).

The issue of whether the hypothesis of unobserved worker characteristics may explain industry or firm differences is unsettled. One way to test this explanation is by testing whether or not wage changes for workers moving between sectors are of the same size as the average wage differentials between sectors. The quality of the worker is supposed to be held constant when he or she moves, and thus the unobserved heterogeneity explanation implies that industry differences should not be observed for movers. Both Krueger and Summers (1988) and especially Gibbons and Katz (1992) test this hypothesis carefully, and their results do not support that industry wage differences are due to unmeasured productivity differences between workers. In another setting Abowd, Kramarz and Margolis (1999) and Leonard and Audenrode (1996) show, however, that unobserved worker characteristics are indeed important in explaining wage differences.

Another literature, to which this paper is more closely related, focuses on which theoretical explanation can best be used to interpret the wage differentials. One recent strand of the literature focuses directly on the profit-pay correlation and uses firm level panel data (Abowd (1989) , Abowd and Lemieux (1990) , Abowd and Lemieux (1993) , Blanchflower, Oswald and Sanfey (1996) , Van Reenen (1996) , Hildreth and Oswald (1997)). The theoretical approach in these papers is a collective bargaining framework in which unionized workers bargain over a firm's rents. Although different papers in this literature prefer different models of collective bargaining,[5] all models produce similar wage equation where a correlation between wages and profits is predicted. Section 3.1 below provides an example of the sort of model considered in this literature.[6]

The above mentioned papers use panels of firm level data for the UK, US and Canada, and estimate versions of the wage equation with rents per worker included. Firm level panel data avoids aggregation problems induced by using industry level cross section data, and some of the panels used allow the authors to control for special market structures and technologies which, even under perfect competition, could lead to high profits and high wages. The results from this literature indicate strongly that more profitable firms pay higher wages when observed human capital characteristics are included as well as firm fixed effects. The puzzle still remains though, whether more profitable firms pay higher wages for identical workers, since omitted worker characteristics may account for the profit-pay effect if a positive correlation exists between wages and profits that occurs via the omission of unobserved worker characteristics.

[5]Labor demand, or right-to-manage, models (of which monopoly union models are an extreme form) involve bargining over wages alone (Oswald (1982) , Hildreth and Oswald (1997) etc...). Efficient contracting models (of which strongly efficient contracting is a special case), on the other hand, imply bargaining over both wages and employment (Abowd (1989) , Abowd and Lemieux (1990)). Manning (1987) suggests a third alternative, in which both wages and employment are bargained over, although the bargaining power of the parties is different in each negotiation. Hosken and Margolis (1997) test a structural form of the Manning (1987) model, which encompasses both labor demand and efficient contracting models in New York State public schools. They reject labor demand in 100% of their school districts, whereas efficient contracting fares slightly better.

[6]See Oswald (1982) , Oswald (1996) and Nickell (1998) for critical reviews of this literature, and especially Nickell (1998) for a critical assesment of the theoretical models.

3 The Basics

In this section, we lay out a simple model of bargaining over wages and employment[7] that leads to implications for the link between profits and wages, and we demonstrated empirically that such a simple correlation exists. We use this model simply as an expository device, and by no means attempt to estimate it in any structural form.

3.1 A Simple Bargaining Model

Consider a bargaining situation in which a firm with a profit function

$$\pi = pf\left(L, K\right) - wL - rK$$

negotiates with a group of workers, not necessarily a union, that attempts to maximize the expected gain of its members, defined as

$$U = \frac{L}{N}\left(w - \overline{w}\right).$$

In the equations given above, p is the market price for the firm's output, $f\left(L, K\right)$ is the production function, supposed concave, with L being employment and K being capital. w is the inside (bargained) wage, while $\overline{w}$ is the outside (opportunity, or market) wage. Finally, r is the rental rate of capital and N is the number of members in the group of workers that is conducting the bargaining.

Defining the bargaining power of the workers as γ, the firm and the workers engage in Nash bargaining over wages and employment, such that the negotiated wage and employment level solve

$$\max_{w,L} U^{\gamma}\pi^{1-\gamma} = \max_{w,L}\left(\frac{L}{N}\left(w - \overline{w}\right)\right)^{\gamma}\left(pf\left(L, K\right) - wL - rK\right)^{1-\gamma}.$$

The bargained employment level in this case equates the value of marginal product with the opportunity wage, i.e.

$$\widehat{L} \text{ solves } p\frac{\partial f}{\partial L} = \overline{w}$$

and the bargained wage w will be determined by

$$w = \widehat{w} = \overline{w} + \frac{\gamma}{1-\gamma}\frac{\widehat{\pi}}{\widehat{L}}.$$

[7] This model of bargaining is similar to those described in Farber (1986) , and is of the strongly efficient type (in the vocabulary of Brown and Ashenfelter (1986)). Abowd (1989) has shown that collective bargaining data from the U.S. is consistent with the implications of this type of bargaining model. Abowd and Kramarz (1993) tend to be skeptical about the relevance of efficient contracting in France.

This implies that inside wages will be a function of opportunity wages, quasi rents per worker (profits evaluated at the opportunity cost of labor) and relative bargaining power.

If we treat bargaining power as a constant in this simple model, this comes down to estimating the specification

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \varepsilon_{i,t}, \tag{1}$$

where $\pi_{J(i,t),t}$ now refers to profits per worker[8] in the firm J that employs the individual i at date t. β_1 is the relative bargaining power of the workers, such that the "absolute" bargaining power γ is equal to $\frac{\beta_1}{1+\beta_1}$. Much of the discussion that follows will be focused on determinants of the outside wage $\overline{w}$ (absorbed into the β_0 in this simplest specification) and of the relative bargaining power β_1. Nevertheless, as the above discussion makes clear, the observed profits per worker in the firm J will be endogenous to the determination of the inside wage w.

3.2 The Stylized Fact

Given the attention that rent sharing has received in the literature, and the demands for it often heard in the popular press, we present here the simple univariate regression of profits per worker on wages, which serves as the starting point for our analysis. The starting point is simple: do workers employed by firms with higher profits earn significantly more than workers employed by firms earning lower profits?

We specify earnings and profits in levels for three reasons. First, because of the presence of loss-making firms in our data, measuring profits per worker in logs would have necessitated discarding observations from poorly performing firms (and possibly introducing selection bias). Second, measuring profits per worker and wages in levels rather than logs means that our estimates can be interpreted directly as the share of each additional franc or kroner in profits per worker that gets translated into additional wages. Finally, the levels-levels specification is the most consistent with the theoretical models underlying the bargaining literature, as in section 3.1 above.

With a notable exceptions of Abowd and Lemieux (1993) and Abowd and Allain (1996) , the empirical rent sharing literature assumes homogeneous rent sharing by specifying identical bargaining power parameters for all firms. This may be an unrealistic assumption, so we estimate both homogeneous and heterogeneous bargaining power models. The homogeneous models assume that $\beta_{1j} = \beta_{1k} = \beta_1$ for all j, k, while the heterogeneous bargaining

[8] The theoretical model refers to quasi rents per worker, rather than profits per worker. This requires evaluating labor costs at their market value. Beginning in section 6, the included covariates allow us to capture this market value of labor. Nevertheless, for the sake of clarity with respect to the commonly held idea of "profit sharing", we prefer to specify the model in terms of profits per worker from the start.

power models specify a functional form for β_{1J}. In particular, we suppose that the relative bargaining power of the workers in firm j can be written as

$$\beta_{1j} = \left[\begin{array}{ccc} 1 & \triangle_{j,t} 1_{[\triangle_{j,t} \geq \triangle_{65}]} & \triangle_{j,t} 1_{[\triangle_{j,t} \leq \triangle_{35}]} \end{array} \right] \beta_1$$

where $\triangle_{j,t}$ represents the change in profits per workers in firm j between t and $t-1$, $\triangle_{65}$ is the 65th percentile in the distribution of such changes over the sample (3.218 1980KF for France and 39860,89 kroner for Norway) and $\triangle_{35}$ is the 35th percentile (-3.847 1980KF for France and -34053,43 kroner for Norway) in this distribution.[9] β_1, in this case, becomes a 3×1 vector of coefficients. This approach allows the bargaining power of the workers to vary as an asymmetric spline with increases and decreases in profits per worker of their employing firm. The specification captures both bargaining power asymmetry in firms with high performance and low performance as compared to the previous year, and introduces simple dynamics into the model. Finally, it should be noted that although we estimate the model on individual data, this model is equivalent to the weighted regression of profits on wages at the firm level, using firm employment as the weight.[10]

Model I in tables 1A and 1B present the results of estimating the benchmark equation with linear profits only and equation 1 with heterogeneous bargaining on our French and Norwegian data. Recall that we restrict our attention to full-time working men in the manufacturing sector, and that the dependent variable in these regressions is measured in levels, not logs. For France, our estimate of β_1 in the homogeneous case is 0.0134 (which implies an absolute bargaining power of 0.0132), whereas our results for Norway are slightly higher, at 0.0250 (and thus an implied absolute bargaining power of 0.0244). Both of these estimates are significantly different from zero. The estimates of the constant terms in the heterogenous case are also significant and slightly larger, being 0.0179 for France and 0.0315 for Norway. The spline terms are negative for exceptionally large increases in profits, and positive of exceptionally large decreases in profits, in both France and Norway (although the increase term in the spline is not significant in France). This suggests some smoothing of wages in response to fluctuation in profits. Two possible explanations that have been put forward by the literature that are consistent with this result are insurance[11] and slow adjustment.[12]

[Insert table 1af about here]

[9] Robustness tests using $\triangle_{60}$ and $\triangle_{55}$ in the place of $\triangle_{65}$, and $\triangle_{40}$ and $\triangle_{45}$ in the place of $\triangle_{35}$ yielded results that were both qualitatively and quantitatively very similar to those presented here.

[10] For lack of matched firm-worker data, the literature often addresses this problem using firm level data. We use individual level data for reasons that will become evident.

[11] The insurance explanation is based on the idea that although a risk neutral firm might share its rents with its risk averse employees, it also insures them against large fluctuations in wages in response to large fluctuations in profits. This would be reflected, in our case, by real wages that increase by slightly less than expected when there is a sharp increase in profits, and real wages that go down by slightly less than expected when there is a sharp fall in profits.

[12] The slow adjustment explanation suggests that the outcome of bargaining over wages depends on the

[Insert table 1an about here]

[Insert table 1bf about here]

[Insert table 1bn about here]

The interpretation of the sizes of these coefficients, which can be applied to all of the estimates that follow, is worth dwelling upon. Considering the homogeneous model, our estimates imply that when profits per worker in France increase by 1000 francs, wages increase by 13 francs, while in Norway when profits per worker increase by 1000 kroner, earnings increase by 25 kroner. In light of the sample statistics (found in table 1 of the appendix), these coefficient estimates correspond to elasticities of 0.002 for France and 0.011 for Norway, i.e. earnings seem (initially) much more responsive to profit fluctuations in Norway than in France.

A more informative measure, however, might be the average increase in wages due to shared rents in each country. This is calculated as the average contribution of profits to wages ($\overline{\pi}\beta_1$) divided by average wages net of this contribution ($\overline{w} - \overline{\pi}\beta_1$). The average increase in wages has the advantage over a simple elasticity in that it allows one to situate the size of the change induced by profits relative to average wages in the absence of this change. In the case of France, this comes to 0.24%, while rent sharing in Norway (on the basis of these preliminary estimates) induces a 1.13 percent increase over wages without rent sharing.

An alternative measure of the role of rent sharing in wage determination, highlighted by Oswald (1996) , is the share of variance in the distribution of wages due to rent sharing. Given that 95% of the mass of a symmetric distribution is within plus or minus 2 standard deviations of the mean, this amounts to first calculating the difference in the contribution of profits to wages between a firm earning the mean of profits per worker plus 2 standard deviations and a firm earning the mean minus 2 standard deviations, and dividing this amount by the difference in wages between a person earning the mean wage plus 2 standard deviations and a person earning the mean wage minus 2 standard deviations.[13] For France, we find that variability in profits per worker "explains" 2.56% of the variability in wages in France, while it explains 9.88% of the variability in wages in Norway. As noted in section 2,

information sets of both parties (notably the threat point of the firm, which is a function of its profits), and that the parties do not learn about profits as they occur, but rather with a lag. Since the implication is that current wages would be a function of past profits, if current profits are higher than past profits, workers should get a smaller share of current profits (corresponding to a fixed share of past profits), and correspondingly wages will seem to make up a larger share of current profits when profits fall.

[13]Given the linearity of our models, this calculation amounts to

$$\frac{\beta_1\left(\overline{\pi}+2\sigma_\pi\right)-\beta_1\left(\overline{\pi}-2\sigma_\pi\right)}{\left(\overline{w}+2\sigma_w\right)-\left(\overline{w}-2\sigma_w\right)}=\frac{4\beta_1\sigma_\pi}{4\sigma_w}=\beta_1\frac{\sigma_\pi}{\sigma_w}.$$

Oswald (1996) notes that previous studies on the United States have found values ranging from 24% to 70%, and analyses using data from the United Kingdom have found shares going from 4% to 25%.

It is clear from tables 1A and 1B that profits are positively significantly related to earnings in both countries, although the explanatory power of these models is tiny. We are only able to explain 0.07% of the variance in earnings by variation in profits for French men in manufacturing, while we explain slightly more (0.27%) of the variance in Norway. The heterogeneous bargaining power model does not add even a hundredth of a percentage point to the explanatory power of profits per worker in France, although it does improve the explanatory power on the Norwegian data by 6 one-hundredths of a percentage point.

Despite the fact that these models have low explanatory power, the significance of the profit coefficient is the point of interest. The significance of this estimate may be artificially inflated through an estimation of its standard error that is biased downwards. Alternatively, the coefficient estimate itself may be biased, be it through omitted variables, endogeneity or non-mean zero measurement error. We treat each of these possibilities below.

4 Statistical Issues

In this section, we address two statistical issues that could affect the estimated significance of our profits variable. First, we consider the possibility of group effects in the covariance matrix of the errors, which could bias the estimated standard errors towards zero. Second, we discuss the impact that measurement error would have on our estimates.

4.1 Group Effects

In an elegant exercise, Moulton (1990) demonstrated how an explanatory variable that was completely unrelated to the outcome of interest might be estimated to have a significant relation if the values of this variable were common across many individuals in the same group. Although his example referred to variables that were constant across all individuals in the same state, for our analysis we are concerned with group effects induced by the fact that all individuals in our data that are employed by the same firm in a given year will have the same value of the profits-per-worker measure. Equation 1 demonstrates how there is no across-individual variation in profits within the same firm in a given year, since the indices are of the form $J(i,t), t$.

In our data, we have (on average) 4.05 observations per firm per year in France, and 21.62 per firm per year in Norway,[14] although this measure hides a wide variety in firm sizes.

[14] Recall that our Norwegian data is drawn from registry information and contains all Norwegian establish-

The presence of such group effects could lead to least squares standard errors that are biased downwards, and thus overestimates of the significance of the profit coefficient.

Model II in tables 1A and 1B applies the correction for group effects as proposed by Huber (1967) and White (1980) .[15] Since this correction only changes the estimated covariance matrix for the estimated coefficients, the parameters and R-squared do not change. We note a clear increase of approximately a factor 6 in France and a factor 15 in Norway for the homogeneous model, while the standard error increases more than fourfold for the constant term in the heterogeneous model in France (the increase is still 15-fold in Norway), but this is not sufficient to reduce the estimated t-statistic on the level profits term below 1.96 in absolute value in any of the four cases. It should be noted, however, that the decreasing spline terms are now significant in neither France nor Norway, and the increasing spline term becomes marginally significant in France. With these considerations in mind, we continue to apply the group effects correction to the estimated covariance matrices in all of models we estimate below, with the exception of those that account for unobserved firm heterogeneity since these models explicitly model the group effect.[16]

4.2 Measurement Error

Given that the coefficient on profits per worker remains significant in a least squares model that corrects for group effects at the firm level, it is useful to consider the role that measurement error might have on our estimates. It is reasonable to expect that our profits per worker variable might be measured with error, since it is derived from accounting statements that are forwarded to the government for tax reasons (among others), and the tax codes allow firms to spread losses out over time to reduce the tax burden in high-profit years. Even in the absence of explicit permission, firms can often shift accounting receivables and payables across fiscal year boundaries to smooth reported profits. This sort of measurement error is not a simple, additive and uncorrelated measurement error, but rather it is mean-reverting,

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ments in manufacturing, with all individuals at each establishment (aggregated up to the enterprise level), whereas our French data is a 1/25 sample of individuals, for which we have kept the observations corresponding to manufacturing and aggregated to the enterprise (firm) level. Correcting for sampling rates implies that French enterprises have an (employment weighted) average employment of just over 100 individuals per year.

[15] Our analysis was carried out in Stata, and thus the Over, Jolliffe and Foster (1996) procedure for standard errors in the presence of group effects where the grouping variable is known is preprogrammed in the "cluster" option. See StataCorp (1999) , p. 178-179 for the explicit formulae.

[16] It is worth noting a technical point that renders the error structure of the fixed effects models slightly different than that with the group effects correction mentioned here. In this section, we consider a group to be a unique firm-year unit. In the fixed effects models, the unit considered is just the firm . There exist year indicators in the fixed effects models, but we do not explictly model the interaction of the fixed effets with the year indicators below in the manner that we do here.

in the sense that years with particularly high profits will have declared profits underreporting the true value, and years with high losses will have declared profits (losses) overreporting true profits (or, equivalently, underreporting true losses).[17]

The standard implication of measurement error, i.e., additive i.i.d. white-noise type measurement error, is an attenuation bias in the estimated coefficient, and under standard conditions of uncorrelated measurement error over time, the attenuation bias is increased when panel data techniques such the within and first-difference estimators are used (Griliches and Hausman 1986). In the following we will, however, argue that the attenuation bias for profits due to the use of a fixed effect estimator is not necessarily increased as compared to the cross section attenuation bias, and that accounting for mean reverting (negatively correlated) measurement error changes the bias formulae even more.

Suppose that (imperfectly) measured profits, $\pi_{j,t}$, can be written as

$$\pi_{j,t} = \pi^*_{j,t} + \nu_{j,t}$$

where $\pi^*_{j,t} \sim N\left(0, \sigma^2_{\pi^*_{j,t}}\right)$ is the true profits and $\nu \sim N\left(0, \sigma^2_\nu\right)$ is the measurement error. Assuming $\text{cov}\left(\pi^*_{j,t}, \nu_{j,t}\right) = 0$, the standard text book result is that the proportional bias in estimating β using OLS is

$$p\lim \hat{\beta} = \beta \frac{\sigma^2_{\pi^*}}{\sigma^2_{\pi^*} + \sigma^2_v}$$

where $\sigma^2_{\pi^*}$ and σ^2_v are the variances of the true profits and the measurement error, respectively. This implies that the estimate of the effect of profits will be attenuated in general. Now, introducing correlation between the true profits and the measurement error, i.e., $cov(\pi^*_{j,t}, \upsilon_{j,t}) \neq 0$, the proportional bias has the following expression

$$p\lim \hat{\beta} = \beta \frac{\sigma^2_{\pi^*} - \sigma_{\pi^* v}}{\sigma_{\pi^*} + \sigma^2_v} \tag{2}$$

where $\sigma_{\pi^* v}$ is the covariance between true profits and the measurement error. The result shows that with a negative relationship between the measurement error and the true profits, which we argue is the case when considering accounting profits, the bias can be smaller than in the classical case.

Both of the previous results carry over to the case of panel data, but now we also have to take into consideration the changes in true profits and measurement error and thus possible

[17] To avoid confusion, it should be noted that this expected mean-reversion in the measurement error term does not imply autocorrelation in reported profits per se, although it might tend to exaggerate the impression of the presence of a unit root in a given firm's profits per worker time series. The precise effect would also depend on the autoregressive properties of the firm's employment time series.

autocorrelation in true profits and the error. If the first-difference fixed effect estimator is used it can be shown that (Griliches and Hausman 1986)

$$p\lim \hat{\beta}_{Panel} = \beta \frac{\left[\left(1-\rho_{\pi^*}\right)/\left(1-\rho_{\upsilon}\right)\right]\sigma^2_{\pi^*}}{\left[\left(1-\rho_{\pi^*}\right)/\left(1-\rho_{\upsilon}\right)\right]\sigma^2_{\pi^*}+\sigma^2_{\upsilon}} \tag{3}$$

Whether the inconsistency or attenuation bias due to (white noise) measurement error is larger for the panel estimator than the OLS estimator depends on the relative size of ρ_{π^*} and ρ_{υ},where $\rho_{\pi^*} = cov\left(\pi^*_{j,t},\pi^*_{j,t-1}\right)/var\left(\pi^*_{j,t}\right)$ and $\rho_{\upsilon} = cov\left(\upsilon_{j,t},\upsilon_{j,t-1}\right)/var\left(\upsilon_{j,t}\right)$. Usually the argument is that $\rho_{\pi^*}>\rho_{\upsilon}$, since a correlation in the true values are expected, while measurement error is supposed to be relatively independent over time (see, for example, Altonji (1986) for the case of wage changes). The reason why the attenuation bias increases due to measurement error is that differencing increases the noise to signal ratio. On the other hand, if $\rho_{\pi^*}<\rho_{\upsilon}$, the conclusion is opposite; the attenuation is lower using a panel estimator than OLS (Griliches and Hausman 1986). Furthermore, including also a negative correlation between $\pi^*_{j,t}$ and $\nu_{j,t}$ implies that the reliability of the results from the panel estimator is increased as shown in the previous equation (Bound and Krueger 1991).

As a result, the estimated effect of profits or rents on earnings will generally be a lower bound for the true effect when one does not correct for measurement error in profits. Furthermore, it is not obvious that the attenuation bias increases when we use panel data estimators to control for firm or worker fixed effects in our analyses in sections 8 and 9 below. Hence, if the estimated effect of profits is weakened when a fixed effect estimator is used to control for unobserved worker or firm heterogeneity, it is likely due to the fixed effects and not due to an exacerbated measurement error problem.

It should be noted that we have no alternative sources of information on "true" profits per worker that we might use to quantify the measurement error problem, and we have no explicit prior beliefs concerning the sizes (or even the relative sizes) of the various components of expressions (2) or (3) that might guide us in a possible correction for this bias. We do, however, have instruments (see section 7 below). Instrumenting a variable measured with error will reduce the variability in the instrumented variable in the model (equation 1), which will tend to offset the additional variability in the variable introduced by the measurement error, and thereby reduce the bias.

5 Inter-Industry Earnings Differentials, Business Cycles and Sectoral Shocks

One important alternative explanation to that of rent sharing is that of inter-industry wage differentials. As noted above in section 2, certain sectors tend to pay "above market" wages, even after controlling for most observable individual characteristics. It is completely possible that these sectors also happen to be the most profitable sectors, and thus the above market wages are simply the result of competition on the demand side of their labor market. Firms in profitable sectors can pay more than those in less profitable sectors, although this premium is not a division of the particular firm's rents per se. In the context of section 3.1, this amounts to considering that demand-side competition bids up $\overline{w}$.

In a more general formulation, the economy as whole, or just a sector of the economy, may be subject to demand shocks that increase revenues. In order to satisfy the additional demand, firms will typically resort (at least in the short-run) to overtime usage before adding workers. As there is a premium in overtime pay, this induces a spurious correlation between high wages and high profits that is independent of a firm's decision to share rents.[18]

Econometrically, equation 1 needs to be modified to account for this possibility. We define a vector of time and sector indicators, and their interactions, as

$$\zeta_{i,t} = \begin{bmatrix} 1_{\{t=t_{\min}\}} \\ \vdots \\ 1_{\{t=t_{\max}\}} \\ 1_{\{S(J(i,t))=Sect_{\min}\}} \\ \vdots \\ 1_{\{S(J(i,t))=Sect_{\max}\}} \\ 1_{\{t=t_{\min}\}}1_{\{S(J(i,t))=Sect_{\min}\}} \\ \vdots \\ 1_{\{t=t_{\min}\}}1_{\{S(J(i,t))=Sect_{\max}\}} \\ \vdots \\ 1_{\{t=t_{\max}\}}1_{\{S(J(i,t))=Sect_{\min}\}} \\ \vdots \\ 1_{\{t=t_{\max}\}}1_{\{S(J(i,t))=Sect_{\max}\}} \end{bmatrix}',$$

where $t_{\min}$ is the earliest date in our data, $t_{\max}$ is the latest date in our data, $Sect_{\min}$ is the first sector in our data,[19] $Sect_{\max}$ is the last sector in our data,[20] $S(\cdot)$ is a function that maps firm identifiers to sector identifiers, and $1_{\{\cdot\}}$ is an indicator function that takes on the value 1 when the subscripted element is true, and 0 otherwise. With this notation, we can write

[18] This is one of many possible explanations that would induce a positive correlation between pay and profits and that would be captured by a set of industry, year and industry-year indicator variables. Others include increased demand-side competition in (possibly specialized) labor markets, general equilibrium effects, changing payroll taxes, etc...

[19] Using the ISIC revision 3 codes at the 2 digit level, this is the manufacture of food products and beverages (ISIC 15) sector.

[20] Using the ISIC revision 3 codes at the 2 digit level, this is the recycling (ISIC 37) sector.

the specification estimated in model III in table 1 as

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \zeta_{i,t}\beta_2 + \varepsilon_{i,t}, \tag{4}$$

Our results suggest that sectoral shocks do indeed influence the relation between profits and earnings in Norway, although the impact is not very large. The coefficient on profits per worker drops from 1.3% to 1.2% of profits going to workers in France, and from 2.5% to 2.2% in Norway in the homogeneous specifications, although in the heterogeneous bargaining power specification the share going to workers in France actually increases from 1.8% to 1.9% when year and sector controls are added, whereas the share of profits per worker going to employees in Norway drops in this specification as well (from 3.1% to 2.8%).The fact that the movements in the coefficients are different in the heterogeneous specifications may be related to the fact that the spline terms were already capturing, in part, the effect of shocks in profits on wages.

This difference in levels between the French and Norwegian results can be at least partially explained by the fact that over our sample period (1987-1995, 1990 and 1991 excluded) in France, sector-level collective agreements had become increasingly irrelevant to wage determination,[21] whereas sector-level collective agreements still play an important role in wage determination in Norway. Further evidence of this comes from the dramatic improvement in the model's explanatory power when sectoral shocks are added, with over 7 percent of the variance in earnings being explained simply by profits and sector shocks in the heterogeneous bargaining power model in Norway (relative to 1/3 of a percent without these shocks). The relative improvement in explanatory power is even larger in France (7/100 of a percent to 2.8%), although the model remains fairly poor when measured by the R-squared statistic.

6 Observable Individual Heterogeneity

Of course, it is practically heresy among labor economists to estimate what amounts to an earnings equation without considering observable individual characteristics.[22] It is quite likely that the profit coefficient is subject to some omitted variable bias, due to the absence of variables (such as those related to the human capital stock of the work force) that would affect simultaneously the firm's profits (by improving productivity) and wages (since these characteristics have value on the labor market). In terms of equation (1), this implies a

[21]See Margolis (1993) for details.

[22]It should be noted that analyses of rent-sharing that are based on firm-side data do precisely this, however, since they typically lack information on the characteristics of the individuals employed by the different firms.

correlation between the profits per worker term and the disturbance of the model that could bias estimates of β_1.

Formally, this suggests a specicifation for our earnings model along the lines of

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \zeta_{i,t}\beta_2 + X_{i,t}\beta_3 + \varepsilon_{i,t}, \tag{5}$$

with $X_{i,t}$ being a set of observable individual characteristics.[23] If equation 5 provides the true specification, the least squares estimator for β_1 will be $\widehat{\beta}_1 = \left[\pi' M_{[\zeta,X]}\pi\right]^{-1} \pi' M_{[\zeta,X]} w$.[24] Least squares estimates of β_1 derived from equation 4 will be of the form $\widetilde{\beta}_1 = \left[\pi' M_{[\zeta]}\pi\right]^{-1} \pi' M_{[\zeta]} w$. If $E(\pi X) \neq 0$, then $\widetilde{\beta}_1 \neq \widehat{\beta}_1$. This will typically be the case if workers with more human capital tend to be more productive and the labor market is not perfectly competitive (so that there is a gap between the wage and the value of marginal product).[25]

Column IV of table 1 presents the results of estimating specification 5. In both France and Norway, the added regressors enter into the model in a highly significant manner, and the explanatory power of both models in both countries increases by a factor of roughly 2.7. That said, the estimated coefficients decrease only slightly, and a Hausman (1978) test does not reject the identity of the two coefficients on profits per worker in models III and IV in either specification. In other words, the significance of profits in the determination of earnings is not simply due to an omitted variable bias.[26]

7 Endogeneity of Profits

To this point, we have failed to explain away the significance of the profits per worker variable in an earnings equation on men in manufacturing in France and Norway by simple least squares-based techniques and by least squares accounting for group effects. One other possible explanation for its significance, however, is that profits are endogenous to wages. When wages increase (all else held equal), profits fall since wages represent the cost of the labor input to production.[27] Of course, the implied correlation here is negative, which

[23] In our case, we consider the following individual observable characteristics: number of years of education and its square, job seniority and its square, and potential labor market experience, its square, cube and fourth power. Note that the only time-invariant individual characteristics here are the number of years of education and its square.

[24] This is a simple application of the Frisch-Waugh(-Lovell) theorem (see Greene (1997) for details). In this notation, $M_{[\cdot]}$ is the matrix that projects onto the space orthogonal to the vector in the subscript.

[25] Abowd et al. (1999) find that this is indeed the case for France between 1976 and 1989, using a specification similar to the one we will use in section 9 below.

[26] Although certain coefficients in model IV in tables 1a and 1b have unexpected signs, it is worth recalling that we estimate our models in levels, and not in logs. This makes our results not directly comparable to those in the majority of the earnings determination literature.

[27] It should be noted that some strands of the literature, notably the contributions of John Abowd and coauthors (Abowd and Allain (1996) , Abowd and Lemieux (1993)), focus explicitly on the share of the

suggests that the endogeneity of profits may actually be biasing our previous estimates downward, and thus the true relation between profits and earnings may be even stronger than presented in the models in table 1.

In order to account for the possibility of endogeneity, we reestimate equation 5 using instrumental variables. Our instrument set includes all of the variables contained in the ζ and X vectors, plus the set of variables in the Z (orthogonalizing) vector described in section 9 below. Furthermore, as additional exclusion restrictions, we include sales per worker[28] and operating subsidies per worker,[29] both contemporaneous and once lagged, as instruments for profits per worker.[30]

Instrumental variables estimation has the additional advantage that, if a white noise measurement error is present in profits, the effect of this measurement error bias (which would reduce the estimated coefficient, see section 4.2) on the estimated coefficient will be eliminated. This is another reason, beyond simple endogeneity, to expect the instrumental variables method to yield higher point estimates than least squares.

Column V of table 2 provides the results of this model, maintaining the correction to the covariance matrix due to the presence of group effects. As is typical for instrumental variables estimation (relative to least squares estimation) almost all of the standard errors increase, most particularly those on profits per worker (as expected). In the case of France, the standard error on profits per worker is 23 times larger in the IV estimates with group effects than in the same specification but under least squares with group effects. Furthermore, the estimated R-squared in France becomes negative. This is possible (in theory and, apparently, in practice) since the error sum of squares is calculated using the observed value of the endogenous variable while the coefficients are estimated with the predicted value of the endogenous variable. Thus the sum of squares being minimized by the least squares technique

quasi-rent per worker (rents when all factors are priced at their opportunity costs) that goes to workers, rather than the share of profits per worker. This measure is not subject to the straightforward source of endogeneity mentioned here, although Abowd and Lemieux (1993) make a convincing case for quasi-rents per worker being endogenous to wage determination nevertheless.

[28]One can think of sales per worker, conditional on the sector being correctly identified, representing a measure of the firm's market power.

[29]Goolsbee (1998) has raised some doubts about the validity of operating subsidies, when in the form of R&D subsidies, as an instrument for profits when considering the earnings of scientists and engineers. We are unable to distinguish the type of operating subsidy in our data, although (by this stage) we control for time-varying industry-specific shocks, and in the next two steps we control for firm and individual fixed effects.

[30]The results of the first step instrumenting equations are given in appendix table 2. In both countries our first-step instrumenting equations pass an F-test for the joint significance of the instruments at the 0.01% level. Testing the validity of the instruments by including them in the main equation and testing for significance suggests that lagged subsidies, at least, is a valid instrument for France (being insignificant at the 10% level), while contemporaneous and lagged sales, and contemporaneous subsidies are valid exclusion restrictions for Norway (by the same criterion).

in the second step is not the sum of squares used in the R-squared calculation. Such an occurrence is typically a sign of weak instruments, and the low explanatory power of our (large) set of covariates in the first step equation (only 0.0972) suggests that this is likely to be the case.[31]

[Insert table 2f about here]

[Insert table 2n about here]

However, as suggested above, the endogeneity of profits was clearly biasing our estimated coefficients on profits downwards in France. The coefficient on profits per worker is 15 times larger in model V than in model IV for France, although it is slightly smaller in Norway, when one moves from least squares estimation to instrumental variables estimation. Despite the fact that the increase in the point estimate is large, it is not large enough relative to the increase in the standard error to maintain significance for profits per worker in France at the 10% level (although it is significant at the 11% level). Profits per worker remains significant in Norway at the 5% level.

8 Fixed Firm-Specific Unobserved Heterogeneity

Another possibility is that the correlation between earnings and profits is due to some aspect of work organization or of the production process that is specific to the firm. Production processes involving dangerous work that would require paying compensating differentials but that also can provide the firm with higher ex-post profits, or an organization of work that substitutes efficiency wages (in a Shapiro and Stiglitz (1984) , equilibrium unemployment type model) for supervision are two sorts of unobservable, firm specific characteristics that could induce a correlation between observed earnings and profits that would not fit in the rent sharing framework, since the earnings premium does not fluctuate with idiosyncratic fluctuations in profits. Once again, the implication is that there could be a correlation between the profits per worker term and the disturbance of equation (1) that would bias estimates of β_1.

Estimating a model that allows for such unobserved characteristics to be correlated with the firm's profits requires panel data in which at least one person is observed in the firm for several years (not necessarily the same person). Over and above the fact that we use our matched worker-firm panel data to merge in profits to each individual's earnings observations, this is another constraint to which only data of this sort can respond. However, if one does

[31]Our instrument set was chosen so as to include only variables that were available for both countries, and in samples with the largest coverage of the manufacturing sector. Additional instruments become available (in France) when choosing narrower estimation samples, but this comes at the cost of reducing our ability to identify the firm and worker effects in section 9 below (see Abowd et al. (1999)for details on identification in the section 9 model).

have a panel in the firm dimension with time variation in profits, wages and observables, one can estimate the model in differences from firm means.

Formally, we define the vector of firm indicators $F_{i,t}$, where $F_{i,t}$ is of dimension 1 x J (where j is the number of different firms in the sample, 47511 for France and 8148 for Norway) and a typical element of $F_{i,t}$, call it $F_{i,t}^{j}$, is an indicator variable $F_{i,t}^{j} = 1_{\{J(i,t)=j\}}$. Thus we write the model we estimate as

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \zeta_{i,t}\beta_2 + X_{i,t}\beta_3 + F_{i,t}\beta_4 + \varepsilon_{i,t}. \tag{6}$$

Given that

$$\begin{aligned} \widehat{\beta}_1 &= \left[\pi' M_{[\zeta,X,F]}\pi\right]^{-1}\pi' M_{[\zeta,X,F]}w \\ &= \left[\pi' M_{[F]}' M_{[\zeta,X]} M_{[F]}\pi\right]^{-1}\pi' M_{[F]}' M_{[\zeta,X]} M_{[F]}w \end{aligned},$$

this is equivalent to the regression described by equation 5, where all of the variables $w_{i,t}$, $\pi_{J(i,t),t}$, $\zeta_{i,t}$, and $X_{i,t}$ are replaced by their values differenced from within firm means of the firm $j = J(i,t)$ that the individual is in at date i, t.

Column VI of table 2 presents these results. The explanatory power of the model jumps dramatically in France to the point where we can explain roughly 25 percent of the variance in log real full year equivalent earnings by our set of explanatory variables and the firm effects in both countries, even considering that we are still instrumenting profits per worker and its square (which reduces the explanatory power of the model). In this framework, the precision of practically all of the estimated parameters increases, and the (instrumented) profits per worker becomes significant again in France.

However, the most notable effect of controlling for unobserved fixed firm heterogeneity is the decrease in the profit coefficient to roughly one third of the value found in the instrumental variables specification in France,[32] although to a level that is roughly 3 times as large as in the most complete least squares specification estimated (model IV). The point estimate in Norway decreases even more, by 44% relative to the instrumental variables estimator. This suggests that there is indeed unobserved heterogeneity in firm compensation practices that induces a spurious positive correlation between profits and earnings, although even in the absence of such a spurious correlation, idiosyncratic fluctuations in firm profits seem to be transmitted through to worker earnings.

9 Worker and Firm Unobserved Heterogeneity

A final possible explanation for the significant relation between profits and earnings that has yet to be specifically addressed is the role of unobserved individual characteristics. Even if

[32]The decrease in the profit coefficient for Norway is much less substantial, which is reasonable given that the coefficient was estimated with a much higher precision in model V.

the link between profits and earnings is not an artefact of differences across firms over time in human capital as measured by education, job seniority and potential experience (if it were, the coefficient on profits in model IV would no longer have been significant), it is possible that this effect is due to variations in the composition of the workforce along unobserved dimensions.

For example, a firm may decide to increase the average quality of the inflow of workers (for a given number of years of education and experience) during years in which demand for its product is abnormally high, with the idea that each of these workers will be more productive than workers of lower quality, which will help the firm meet the additional demand and not suffer the negative reputational effects that being an unreliable supplier might have. This workers, although observationally (to the econometrician) equivalent to lower quality workers, will tend to earn more on the labor market. Furthermore, the increase in average worker quality (and thus wages) will occur at the same time as firms are able to increase their margins (due to the high demand shock), and thus at a time when they will be making higher profits. The result will be an observed positive relation between earnings and profits, even conditional on all of the other variables included up until here, that is due exclusively to changes in work force composition, and not to rent sharing.

Estimation of such a model is complicated. The full least squares solution, even after differencing within individuals, would still lead to a model to be estimated by least squares of dimension well over 47000 in France and over 12000 in Norway. Abowd et al. (1999) provide an econometric technique that allows us to recover (under certain conditional orthogonality assumptions) unbiased estimators of all of the parameters of interest in our model, and in particular the vector of coefficients β_1. If we define a vector of individual indicator variables, $D_{i,t}$, in a manner similar to the definition of $F_{i,t}$ in section 8 above, we can write the model of interest as

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \zeta_{i,t}\beta_2 + X_{i,t}\beta_3 + F_{i,t}\beta_4 + D_{i,t}\beta_5 + \varepsilon_{i,t}. \tag{7}$$

The idea is to find a set of variables Z such that $E\left[\pi M_{[Z]}F\right] = 0$, $E\left[\zeta M_{[Z]}F\right] = 0$, $E\left[XM_{[Z]}F\right] = 0$ and $E\left[DM_{[Z]}F\right] = 0$.[33] Under these assumptions, we can estimate the much simpler model

$$w_{i,t} = \beta_0 + \pi_{J(i,t),t}\beta_1 + \zeta_{i,t}\beta_2 + X_{i,t}\beta_3 + D_{i,t}\beta_5 + Z_{i,t}\lambda + \varepsilon_{i,t}. \tag{8}$$

by differences from within-individual means to recover consistent estimates of β_1 and β_2 and of the time-varying components of β_3. Additional steps are needed to recover β_4 and to

[33] For the purposes of this paper, we defined the the Z vector as the the set of interactions of education and average (over all observations for the individual) experience with the set of 22 industry indicators, average firm size, average sales and average operating subsidies (all over the set of years of data available for the employer). Formally, this is the kronecker product of the 2-dimensional vector of fixed individual characteristics and the 25-dimensional vector of fixed firm characteristics.

separately distinguish β_5 from the time-invariant component of β_3. Finally, a correction (to correctly measure the error sum of squares and to account for the error degrees of freedom) is also needed in order to recover an estimate of the covariance matrix of the estimated parameters.[34]

Column VII of table 2 provides the results of estimating this most complicated, and most encompassing specification. In this specification, all time-invariant sources of spurious correlation on the firm and worker side are eliminated, and a large share of the observable time-varying characteristics are controlled for as well. Furthermore, we are still instrumenting profits, to control for possible endogeneity. If there remains any significant relation between profits and earnings, it can only be due to idiosyncratic fluctuations in profits that are directly converted into earnings.

This additional correction for changes in the unobserved characteristics of work force composition seems to be sufficient to reduce the profits per worker term to insignificance in France, and to reduce its significance level below 5% in Norway (although it remains significant at the 10% level). The point estimates of the coefficients on profits fall slightly, while the standard error increases. Another sign of the importance of unobserved individual characteristics for the explanation of earnings is the increase of the explanatory power of the model by 69% that occurs in both countries when individual fixed effects are added.

One can not, therefore, reject the hypothesis of an absence of rent sharing in France, although this failure to reject is driven largely by the decreased precision in our estimated coefficient. On the other hand, this remains the most likely explanation for Norway. Although the estimated coefficient on profits again drops by 14%, it remains significantly related to the log of real full year equivalent gross earnings. The explanatory power of the model is indeed increased by the addition of unobserved individual heterogeneity in both countries, such that we can explain roughly 40 percent of the variance in earnings in France and 50% in Norway.[35]

[34]See Abowd et al. (1999) for details.

[35]It should be noted here that our estimates of the model described in 7 indicate a lower explanatory power than the model estimated by Abowd et al. (1999) . Their equivalent of the column VII estimation explained 77.20% of the variance in log real full year equivalent compensation costs, while we explain 39.30% of the variance in log real full year equivalent gross earnings. There are several possible sources of this difference. First, we consider the period 1987-1995 (1990 and 1991 excluded), while they consider the period 1976-1986 (1981 and 1983 excluded). As noted above, sector-level collective bargaining gradually became obsolete over the 1980s, and these collective agreements could have been very important determinants of earnings. Second, they consider men and women (albeit seperately), while we consider only men. This is because of the third difference, that we consider only manufacturing, while they consider the entire economy; there are too few repeat observations on individuals in the female data when restricting one's attention to manufacturing. This restriction was necessary since manufacturing is the only sector of the economy for which Norwegian data was available.

10 Conclusion and Interpretation

In this paper, we have put the rent sharing theory through a series of indirect tests that are designed to allow us to rule out observationally equivalent explanations of the positive correlation between profits per worker and wages. We have shown that the positive correlation is not due to econometric errors based on group effects in residuals or measurement error, it is not due to omitted variable bias in observable individual productive characteristics, it is not due to inter-industry wage differentials, the business cycle or sectoral shocks, it is not due to endogeneity of profits (in Norway) and it is not due to time-invariant unobserved heterogeneity in production processes or work organization at the firm level. We do find, however, that (conditional on all of the previous explanations) changes in the unobservable characteristics of the work force tend to reduce the degrees of freedom available for estimating the covariance matrix, which increases the estimated standard error by enough to render the coefficient on profits per worker insignificant in France. In Norway, however, the significant relation holds even in the most complete specification. Although we treat rent sharing as a residual explanation for the significance of the coefficient, it still survives in Norway at the 10 percent level.

Why might we expect this to be the case? As noted above, collective bargaining played a much more important role in earnings determination in Norway over our sample period than it did in France. Furthermore, collective bargaining in Norway is a multi-step process, with sector-wide framework agreements being settled first, and then separate agreements being negotiated at successively finer levels of disaggregation, even down to the plant level. In a collective bargaining agreement, earnings are fixed as a function of characteristics of the job held, and not necessarily of the individual (beyond the obvious correlation between qualifications and occupations). The agreement may, however, include explicit profit sharing provisions. The 44 percent reduction in the coefficient on profits in Norway that occurred when we added firm fixed effects suggests that this is an economically important phenomenon. Given these negotiated wages, Norwegian firms use their flexibility in hiring and firing to try to adapt their workforce composition to include workers that will generate profits. The relatively important role of wage drift above the collectively bargained levels (Holden 1989) suggests that some of this unobserved quality is being compensated, and hence the importance of the reduction insignificance due to individual fixed effects. Nevertheless, although the significance of the relation is reduced in Norway, it is not eliminated.

In France, another story may be relevant. Despite strict rules (by North American standards) concerning layoffs (Lefebvre 1996), many French workers are able to switch employers without much intervening unemployment (Margolis forthcoming). Our results in section 9 suggest that what may be happening is that the workers with unobserved (to the econometrician) characteristics that are worth the most on the labor market are sorted into firms that become profitable. The causality running from unobserved worker characteristics to wages would explain the role of the instrumental variables in rendering the coefficient initially non-significant, while the most complete model allows us to directly capture this heterogeneity, thereby rendering the coefficient on profits insignificant.

It should be noted that that the coefficient on profits becomes insignificant in table 2 through an increase in the standard error, and not through a decrease in the point estimate of the effect of profits on wages. However, given the weakness of our instruments, we expect the standard errors of our estimates to be very large. Furthermore, the most complete specification consumes over a third of the degrees of freedom available with the individual effects alone in France, so the increased standard errors may just be a problem of insufficient degrees of freedom (although ther are still 285316 degrees of freedom left to identify the variance of the disturbance). As a further check on the role of instrumenting versus fixed effects, we reran model VII with observed, instead of instrumented profits. In this case, the standard error of the profit coefficient was of a similar order of magnitude to the model VI regressions (1.852E-03 in model VII without instruments versus 1.036E-03 in model VI). However, the point estimate was two orders of magintude smaller than the model VI estimates (1.475E-04 in model VII without instruments versus 6.326E-02 in model VI). This suggests that there may indeed be a "truly" insignificant coefficient on profits per worker that comes as a result of including individual fixed effects, in the sense that it is neither economically nor statistically significant, although such a conclusion must be treated carefully as it depends on a model that does not instrument the (clearly endogenous) profits per worker variable.

In light of all of this evidence, the answer to our initial question deserves a mixed response. Most alternative explanations for the positive correlation between profits and wages fall by the wayside, although the role of sorting of individuals by their unobserved characteristics may merit further attention. In our case, the reduction in precision of our estimates that such a procedure implies means that we can no longer reject the hypothesis of zero rent sharing in France, although this hypothesis remains valid at the 10 percent level in Norway.

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A Appendix - Data description

A.1 France

The data used in this paper come from several administrative sources available at INSEE, the French National Institute for Statistics and Economic Studies. The two main data sources are the DADS (Déclarations annuelles des données sociales, or Annual social data declarations) and the BIC-BRN files of the SUSE (Système unifié des statistiques d'entreprises, or Unified enterprise statistics system). The first source consists of private and state-owned employer filings with the government for the purpose of determining retirement pensions and other social services. This file contains one record for each unique employee-establishment-enterprise-year combination, with standardized identifiers for each element. The second source is derived from firms' filings with the tax authorities. This contains one record for each unique enterprise-year combination, using identifiers common to the DADS data.

For a detailed description of the treatment of the DADS data, see Abowd et al. (1999) and Roux (2000). These data were merged with the BIC-BRN files, which cover all enterprises with at least 100,000 francs worth of assets. After harmonizing sector codes and aggregating the DADS data to the individual-enterprise-year level, the data were merged by enterprise-year identifiers. Given the hole in the DADS data in 1990 (a census year in France), our merged sample consists of an unbalanced panel of linked firm-worker data from 1986 through 1989 and from 1991 through 1995. Given that we use lags in certain firm variables as instruments, this further reduces the analysis sample by eliminating the 1986 and 1991 data. We then restricted our attention to manufacturing (ISIC3 codes beginning with numbers between 15 and 37, inclusive) and full-time working men aged 16-55 in the set of observations which could be matched in both samples.

Our earnings variable is real (1980) gross, full year equivalent labor earnings and our profits variable is operating income (excédent brut d'exploitation). Our education and school leaving age information comes from a preliminary merge and interpolation (see Abowd et al. (1999)) of the DADS data with data from the EDP (Echantillon démographique permanent, or Permanent demographic sample), a 1/10 sample of individuals in the DADS. Experience is calculated as age minus school leaving age and education as school leaving age minus 6. Seniority is determined from pre-sample information (going back to 1976), and is imputed for left-censored (at 1976) employment spells; see Abowd et al. (1999) for details.

A.2 Norway

For Norway, we use the linked employer-employee data set based on different administrative register files from Statistics Norway, supplemented with firm level information on economic performance for plants from the annual census for manufacturing, see Halvorsen, Jenssen and Foyn (1991) . The sample period covers 1986 to 1994.

In these administrative registers, individuals are characterized by their personal identity code and firms with an identification code. This enables us to match persons to firms and to combine information on education, age, tenure etc. with employer characteristics at the individual level. In the second quarter each year every worker is matched to the individual's main employer. The start date of this match is provided by the main employer, as is the stop date if it finishes within the year. Our database contains yearly information for all employed individuals over the age of 16 and all plants in Norway. The employers are defined at the firm level by an identification code dependent on geographical location and independent of ownership conditions. We restrict our attention to firms with an average size of at least five employees, since plant or firm specific information is not available for plants below five employees. When merging in the data from the census file for the manufacturing sector for the econometric analysis, the match by plant numbers is about 90 percent. For the econometric analysis only male workers with full-time jobs (30 hours or more per week) were included.

The real (1990) earnings was derived using the annual plant wage payment to the worker (including salaries and wages in cash and kind). Profits per worker is defined as revenues minus costs of materials and costs of labor divided by the number of workers. Education level is based on the normal duration of the education and includes only completed (and highest attained) education, and all formal education courses exceeding 300 hours are registered. The consumer price index was used to derive the wage variable in 1990 values. Tenure is defined as the number of years worked for each employer. Plant size is defined as the total number of workers at the plant level including also part time workers. As an exclusion criterion, we used an hourly wage rate, and excluded an hourly wage rate below 30 kroner per hour and above 500 kroner per hour, since these are obviously either below or above possible wage rates.

[Insert appendix table 1 about here]

[Insert appendix table 2 about here]

Table 1A: Results from OLS Estimates, Without Profit Spline				
Model	I	II	III	IV
France				
Profits / Worker	1.339E-02 *	1.339E-02 *	1.237E-02 *	1.159E-02 *
	(7.355E-04)	*(4.766E-03)*	*(4.795E-03)*	*(4.803E-03)*
Years of Schooling				-7.938E+00 *
				(4.820E-01)
(Years of Schooling^2)/100				5.204E+01 *
				(2.244E+00)
Job Seniority				-8.536E-01 *
				(7.559E-02)
(Job Seniority^2)/100				2.221E+00 *
				(2.796E-01)
Potential Labor Market Experience				3.576E+00 *
				(1.744E-01)
(Potential Labor Market Experience^2)/100				-1.017E+01 *
				(2.189E+00)
(Potential Labor Market Experience^3)/1000				2.979E+00 *
				(9.876E-01)
(Potential Labor Market Experience^4)/10000				-4.454E-01 *
				(1.423E-01)
R-Squared	0.0007	0.0007	0.0278	0.0737
Number of Obs.	504586	504586	504586	504586
Number of Groups (Enterprise-Years)	-	124448	124448	124448

Table 1A: Results from OLS Estimates, Without Profit Spline (continued)

Model	I	II	III	IV
Norway				
Profits / Worker	2.499E-02 *	2.499E-02 *	2.239E-02 *	2.146E-02 *
	(4.713E-04)	*(7.364E-03)*	*(6.436E-03)*	*(5.922E-03)*
Years of Schooling				-2.115E+04 *
				(1.47E+03)
(Years of Schooling^2)/100				1.700E+05 *
				(6.66E+03)
Job Seniority				-2.547E+03 *
				(3.09E+02)
(Job Seniority^2)/100				6.873E+03 *
				(8.53E+02)
Potential Labor Market Experience				6.574E+03 *
				(4.91E+02)
(Potential Labor Market Experience^2)/100				-8.382E+03 *
				(2.96E+03)
(Potential Labor Market Experience^3)/1000				-2.148E+03 *
				(7.511E+02)
(Potential Labor Market Experience^4)/10000				3.569E+02 *
				(6.640E+01)
R-Squared	0.0027	0.0027	0.0758	0.2138
Number of Obs.	1041039	1041039	1041039	1041039
Number of Groups (Enterprise-Years)	-	48148	48148	48148
Group Effects	No	Yes	Yes	Yes
Year Effects	No	No	Yes	Yes
Sector Effects	No	No	Yes	Yes
Year*Sector Effects	No	No	Yes	Yes

Sources: DADS and SUSE Data for France, Industrial Statistics and Register Data for Norway.

Notes: Standard errors are in parentheses. Group effects estimations use the correction for common variance components within groups proposed by Greenwald (1983) and Moulton (1990). * Refers to a significant coefficient at the 95% confidence level, ° refers to a significant coefficient at the 90% confidence level. Both the French and Norwegian data cover all 23 manufacturing sectors (ISIC revision 3 codes). The French data cover 7 years (1987-1989 and 1992-1995) and the Norwegian data cover 8 years (1988-1995). Profit change splines are interacted with current profits.

Table 1B: Results from OLS Estimates, With Profit Spline				
Model	I	II	III	IV
France				
Profits / Worker	1.787E-02 *	1.787E-02 *	1.938E-02 *	1.927E-02 *
	(1.225E-03)	*(5.326E-03)*	*(4.552E-03)*	*(4.359E-03)*
$(\Delta(\text{Profits/Worker})-\Delta_{65})*1(\Delta>\Delta_{65})$	-1.780E-06 *	-1.780E-06 °	-2.780E-06 *	-3.080E-06 *
	(3.300E-07)	*(1.020E-06)*	*(8.690E-07)*	*(8.560E-07)*
$(\Delta(\text{Profits/Worker})-\Delta_{35})*1(\Delta<\Delta_{35})$	1.140E-07	1.140E-07	1.640E-07	1.600E-07
	(9.060E-08)	*(6.370E-07)*	*(6.120E-07)*	*(6.050E-07)*
Years of Schooling				-7.931E+00 *
				(4.818E-01)
(Years of Schooling^2)/100				5.202E+01 *
				(2.243E+00)
Job Seniority				-8.553E-01 *
				(7.554E-02)
(Job Seniority^2)/100				2.219E+00 *
				(2.795E-01)
Potential Labor Market Experience				3.576E+00 *
				(1.745E-01)
(Potential Labor Market Experience^2)/100				-1.015E+01 *
				(2.189E+00)
(Potential Labor Market Experience^3)/1000				2.975E+00 *
				(9.877E-01)
(Potential Labor Market Experience^4)/10000				-4.453E-01 *
				(1.423E-01)
R-Squared	0.0007	0.0007	0.0279	0.0739
Number of Obs.	504586	504586	504586	504586
Number of Groups (Enterprise-Years)	-	124448	124448	124448

Table 1B: Results from OLS Estimates, With Profit Spline (continued)

Model	I	II	III	IV
Norway				
Profits / Worker	3.148E-02 *	3.148E-02 *	2.819E-02 *	2.035E-02 *
	(5.378E-04)	*(8.258E-03)*	*(7.003E-03)*	*(5.888E-03)*
$(\Delta(\text{Profits/Worker})-\Delta_{65})*1(\Delta>\Delta_{65})$	-1.110E-09 *	-1.110E-09 *	1.630E-08 *	-6.480E-10 *
	(5.740E-11)	*(4.450E-10)*	*(4.260E-09)*	*(3.160E-10)*
$(\Delta(\text{Profits/Worker})-\Delta_{35})*1(\Delta<\Delta_{35})$	2.200E-08 *	2.200E-08	-9.720E-10 *	1.270E-08
	(1.170E-09)	*(1.837E+03)*	*(3.840E-10)*	*(3.730E-09)*
Years of Schooling				-2.112E+04 *
				(1.48E+03)
(Years of Schooling^2)/100				1.698E+05 *
				(6.69E+03)
Job Seniority				-2.551E+03 *
				(3.09E+02)
(Job Seniority^2)/100				6.883E+03 *
				(8.52E+02)
Potential Labor Market Experience				6.574E+03 *
				(4.91E+02)
(Potential Labor Market Experience^2)/100				-8.382E+03 *
				(2.96E+03)
(Potential Labor Market Experience^3)/1000				-2.148E+03 *
				(7.511E+02)
(Potential Labor Market Experience^4)/10000				3.569E+02 *
				(6.640E-03)
R-Squared	0.0033	0.0033	0.0762	0.2140
Number of Obs.	1041039	1041039	1041039	1041039
Number of Groups (Enterprise-Years)	-	48148	48148	48148
Group Effects	No	Yes	Yes	Yes
Year Effects	No	No	Yes	Yes
Sector Effects	No	No	Yes	Yes
Year*Sector Effects	No	No	Yes	Yes

Sources: DADS and SUSE Data for France, Industrial Statistics and Register Data for Norway.

Notes: Standard errors are in parentheses. Group effects estimations use the correction for common variance components within groups proposed by Greenwald (1983) and Moulton (1990). * Refers to a significant coefficient at the 95% confidence level, ° refers to a significant coefficient at the 90% confidence level. Both the French and Norwegian data cover all 23 manufacturing sectors (ISIC revision 3 codes). The French data cover 7 years (1987-1989 and 1992-1995) and the Norwegian data cover 8 years (1988-1995). Profit change splines are interacted with current profits.

Table 2: Results from IV and Panel Estimates

Model	V	VI	VII
France			
Profits / Worker	1.765E-01	6.326E-02 *	6.104E-02
	(1.102E-01)	*(1.036E-03)*	*(1.665E-01)*
Years of Schooling	-7.309E+00 *	-6.898E+00 *	-1.179E-01
	(5.907E-01)	*(3.937E-01)*	*(1.037E-01)*
(Years of Schooling^2)/100	4.909E+01 *	4.716E+01 *	2.218E+01 *
	(2.765E+00)	*(1.682E+00)*	*(8.032E-01)*
Job Seniority	-9.709E-01 *	-1.240E+00 *	7.383E-01 *
	(1.113E-01)	*(5.905E-02)*	*(2.009E-01)*
(Job Seniority^2)/100	2.296E+00 *	2.677E+00 *	-2.885E+00 *
	(3.066E-01)	*(2.143E-01)*	*(8.730E-01)*
Potential Labor Market Experience	3.601E+00 *	3.216E+00 *	2.718E+00 *
	(1.836E-01)	*(2.154E-01)*	*(4.689E-01)*
(Potential Labor Market Experience^2)/100	-9.594E+00 *	-5.248E+00 *	-3.105E+00
	(2.311E+00)	*(2.166E+00)*	*(5.281E+00)*
(Potential Labor Market Experience^3)/1000	2.672E+00 *	9.659E-01	4.243E-01
	(1.039E+00)	*(8.339E-01)*	*(2.309E+00)*
(Potential Labor Market Experience^4)/10000	-4.062E-01 *	-1.814E-01 °	-5.452E-02
	(1.486E-01)	*(1.073E-01)*	*(3.245E-01)*
R-Squared	-0.0200	0.2325	0.3930
Number of Obs.	504586	504586	504586
Number of Groups (Enterprises or Enterprise-Years)	124448	47511	47511
Number of Individuals	-	-	171603

Table 2: Results from IV and Panel Estimates (continued)

Model	V	VI	VII
Norway			
Profits / Worker	2.714E-02 *	1.530E-02 *	1.316E-02 °
	(7.113E-03)	*(6.360E-03)*	*(7.574E-03)*
Years of Schooling	-2.109E+04 *	1.460E-07 *	1.680E-08 *
	(6.844E+02)	*(9.430E-09)*	*(3.740E-09)*
(Years of Schooling^2)/100	1.695E+05 *	-3.160E-07	-3.750E-05
	(3.131E+03)	*(4.330E-06)*	*(2.040E-06)*
Job Seniority	-2.549E+03 *	1.617E+04 *	6.634E-02 *
	(1.84E+02)	*(3.137E+02)*	*(9.909E-04)*
(Job Seniority^2)/100	6.835E+03 *	-1.388E+05 *	-1.333E+06 *
	(5.47E+02)	*(2.459E+04)*	*(5.94E+04)*
Potential Labor Market Experience	6.516E+03 *	4.315E+01 *	5.448E+01 *
	(2.93E+02)	*(6.426E+00)*	*(1.82E+01)*
(Potential Labor Market Experience^2)/100	-8.056E+03 *	4.466E+05 *	5.862E+05 *
	(1.83E+03)	*(4.231E+04)*	*(9.20E+04)*
(Potential Labor Market Experience^3)/1000	-2.219E+03 *	2.066E+04	-2.663E+05 *
	(4.71E+02)	*(2.817E+04)*	*(4.93E+04)*
(Potential Labor Market Experience^4)/10000	3.623E+02 *	5.232E+02 *	-2.301E+02 *
	(4.211E+01)	*(6.937E+01)*	*(1.130E+02)*
R-Squared	0.2146	0.2912	0.4920
Number of Obs.	1041039	1041039	1041039
Number of Groups (Enterprises or Enterprise-Years)	48148	12345	12345
Number of Individuals	-	-	122153
Group Effects	Yes	-	-
Year Effects	Yes	Yes	Yes
Sector Effects	Yes	Yes	Yes
Year*Sector Effects	Yes	Yes	Yes

Sources: DADS and SUSE Data for France, Industrial Statistics and Register Data for Norway.

Notes: Standard errors are in parentheses. Group effects estimations use the correction for common variance components within groups proposed by Greenwald (1983) and Moulton (1990). * Refers to a significant coefficient at the 95% confidence level, ° refers to a significant coefficient at the 90% confidence level. Both the French and Norwegian data cover all 23 manufacturing sectors (ISIC revision 3 codes). The French data cover 7 years (1987-1989 and 1992-1995) and the Norwegian data cover 8 years (1988-1995). Profit change splines are interacted with current profits.

Appendix Table 1: Descriptive Statistics

	Mean	Std. Dev.
France		
Real Full-Year Equivalent Earnings (1980 KF)	77.6712	80.6337
Profits / Worker (1980 KF)	13.8029	154.2769
$(\Delta(\text{Profits/Worker})-\Delta_{65})*1(\Delta>\Delta_{65})$	13.3546	124.3388
$(\Delta(\text{Profits/Worker})-\Delta_{35})*1(\Delta<\Delta_{35})$	-9.1053	62.1448
Sales / Worker (1980 KF)	562.0781	821.8943
Operating Subsidies / Worker (1980 KF)	1.7744	26.7658
Years of Schooling	12.4420	1.7670
(Years of Schooling^2)/100	1.5793	0.4114
Job Seniority	7.1818	7.9261
(Job Seniority^2)/100	1.1440	2.0153
Potential Labor Market Experience	18.2438	9.8677
(Potential Labor Market Experience^2)/100	4.3021	3.8295
(Potential Labor Market Experience^3)/1000	11.5008	13.5948
(Potential Labor Market Experience^4)/10000	33.1727	47.9684
Number of Obs.	504586	
Norway		
Real Full-Year Equivalent Earnings (1990 NOK)	2.579E+05	1.381E+05
Profits / Worker (1990 NOK)	1.187E+05	2.412E+05
$(\Delta(\text{Profits/Worker})-\Delta_{65})*1(\Delta>\Delta_{65})$	1.351E+11	7.281E+12
$(\Delta(\text{Profits/Worker})-\Delta_{35})*1(\Delta<\Delta_{35})$	-1.324E+10	4.61E+11
Sales / Worker (1990 NOK)	1.140E+06	1.140E+06
Operating Subsidies / Worker (1990 NOK)	21.2497	137.0129
Years of Schooling	10.8700	2.3700
(Years of Schooling^2)/100	1.1739	0.5035
Job Seniority	7.8900	5.6900
(Job Seniority^2)/100	0.8960	1.5793
Potential Labor Market Experience	22.9000	13.3000
(Potential Labor Market Experience^2)/100	6.9460	6.7046
(Potential Labor Market Experience^3)/1000	24.3796	31.5069
(Potential Labor Market Experience^4)/10000	93.1992	147.9623
Number of Obs.	1041039	

Sources: DADS and SUSE Data for France, Industrial Statistics and Register Data for Norway.

Notes: Profit change splines are interacted with current profits in the estimation.

Appendix Table 2
First Stage IV Estimating Equations

	France		Norway	
Sales / Worker	4.181E-02	*	9.717E-02	*
	(5.871E-04)		*(1.562E-04)*	
Lag(Sales / Worker)	-2.469E-03	*	-5.192E-03	*
	(6.609E-04)		*(1.605E-04)*	
Operating Subsidies / Worker	-1.105E-01	*	-2.709E-01	*
	(1.721E-02)		*(1.315E-02)*	
Lag(Operating Subsidies / Worker)	-6.600E-02	*	-6.115E-01	*
	(1.629E-02)		*(1.245E-02)*	
Years of Schooling	-2.791E+00	*	-2.776E+03	*
	(7.452E-01)		*(4.936E+02)*	
(Years of Schooling^2)/100	1.251E+01	*	1.709E+04	*
	(3.008E+00)		*(2.148E+03)*	
Job Seniority	7.482E-01	*	1.507E+03	*
	(9.776E-02)		*(6.751E+01)*	
(Job Seniority^2)/100	-1.161E+00	*	-4.474E+03	*
	(3.728E-01)		*(2.273E+00)*	
Potential Labor Market Experience	1.756E-01		2.911E+02	
	(4.111E-01)		*(3.133E+02)*	
(Potential Labor Market Experience^2)/100	-7.227E-01		-6.410E+03	*
	(3.706E+00)		*(2.205E+03)*	
(Potential Labor Market Experience^3)/1000	1.411E+00		2.033E+03	*
	(1.445E+00)		*(6.055E+02)*	
(Potential Labor Market Experience^4)/10000	-2.136E-01		-1.955E+02	*
	(1.879E-01)		*(5.644E+01)*	
R-Squared	0.0972		0.5110	
Number of Obs.	504586		1041039	

Sources: DADS and SUSE Data for France, Industrial Statistics and Register Data for Norway.
Notes: Standard errors are in parentheses. All models contain year effects, sector effects, the interaction of year and sector effects, and the set of Z variables described in section 8.
* Refers to a significant coefficient at the 95% confidence level, ° refers to a significant coefficient at the 90% confidence level. Both the French and Norwegian data cover all 23 manufacturing sectors (ISIC revision 3 codes). The French data cover 7 years, the Norwegian data cover 8 years.