Insurance within the Firm[*]

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Abstract

We evaluate the allocation of risk between firms and their employees using a long panel of matched employer-employee data. Unlike the previous literature, we focus on idiosyncratic shocks to the firm, the correct empirical counterpart of the theoretical notion of diversifiable risk. We allow for both transitory and permanent shocks to firm output and find that firms are willing to fully absorb transitory fluctuations but insure workers only partially against permanent shocks. Risk-sharing considerations can account for about 11 percent of the overall earnings variability, the remainder originating in idiosyncratic shocks to individual workers. Our welfare calculations indicate that firms are an important vehicle of insurance provision. Finally, we permit our insurance parameters to vary according to firm and worker characteristics. Such heterogeneity is not allowed for in simple versions of the insurance hypothesis, but has a sound economic justification in more general models with bankruptcy, informational asymmetries and differences in risk aversion between the parties involved.

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1 Introduction

The idea that an intrinsic component of the entrepreneur-worker relation is the allocation of risk has a long tradition in economics. It dates back at least to the work of Knight (1921), who ascribes the very existence of the …rm to its role as an insurance provider:

> "...the system under which the con…dent and venturesome assume the risk and insure the doubtful and timid by guaranteeing to the latter a speci…ed income in return for an assignment of the actual results...is the enterprise and wage system of industry. Its existence in the world is the direct result of the fact of uncertainty" (p.269-70).

This view underlies the theory of the …rm as an insurance device formalized in the implicit contract model of Borch (1962), Baily (1974) and Azariadis (1975): risk-neutral entrepreneurs provide full insurance to risk averse workers and insulate their salaries from adverse shocks to production. Modern …nance theory stresses the di¤erential access of …rms and workers to insurance markets rather than di¤erences in preferences: …rms (shareholders) can diversify idiosyncratic risk away and so act as risk-neutral agents in the relationship with workers, who have limited access to …nancial markets. In this view, the implicit contract framework is theoretically less justi…able for undiversi…able risk. Grossman and Hart (1981) maintain that if the …rm's marginal product is correlated with the income shareholders get from other …rms, then the …rm's pro…t will not be a diversi…able risk. Similarly, Romer (2000) argues that "because the …rm's owners can diversify away speci…c risk by holding a broad portfolio, the assumption that the …rm is risk-neutral is reasonable for …rm-speci…c shocks. For aggregate shocks, however, the assumption that the …rm is less risk-averse than the workers is harder to justify" (p. 434). This implies that the insurance role of the …rm pertains to idiosyncratic shocks.

The assumption that …rms can diversify idiosyncratic risk perfectly, and therefore o¤er full insurance to workers, might be an extreme one. First, empirical studies of households' portfolios show that investment in equities tends to be concentrated, particularly for private equity owners. For example, Moskowitz and Vissing-Jorgensen (2001) …nd that in the

United States households with private equity ownership invest on average almost two thirds of their private holdings in a single company in which they have an active management interest. In this case, the full-diversification assumption clearly fails and equity owners might want to shift part of the enterprise risk onto their employees. Second, Gamber (1988) shows that if the provision of insurance in an implicit contract model is constrained by the possibility of bankruptcy, then the optimal level of insurance depends on the persistence of the shocks: in particular, the firm might be less willing to insure permanent shocks than transitory ones, again placing the full insurance result in doubt.[1] Finally, the modern contract theory has emphasized the role of incentives as an obstacle to insurance provision by the firm. Principal-agent models stress the trade-off between insurance and incentives in determining the optimal compensation scheme: in the presence of moral hazard, the need to provide incentives to workers prescribes a link between their compensation and the firm's performance, in which case the full insurance solution may not obtain even if entrepreneurs could perfectly diversify idiosyncratic risk (Holmstrom and Milgrom, 1987).[2] In these models the optimal level of insurance will depend on specific characteristics of the contracting parties, such as informational asymmetry, differences in risk aversion , and other deviations from the simple benchmark.

In sum, the amount of wage insurance firms provide to their workers against insurable shocks is an empirical question. Is the typical real-world worker-compensation scheme close to the Azariadis-Baily model of full insurance or does it depart significantly from that benchmark? Previous empirical work on this issue has been based on aggregate data.[3] Given

[1] Baily (1980) notes that "the fundamental insurance a firm provides is that it will not use the temporary pressure that a recession creates to lower its workers' wages...A fundamental property of implicit contracts is that one party will default when the long-run gains from compliance fall below the losses. This property suggests that wages cannot remain permanently out of line with their long-run market valuation " (p. 128; italics in original).

[2] In moral hazard models the informational frictions are due to the fact that the principal cannot disentangle the effects of underlying exogenous shocks from workers' effort on output. Even in this more general context the relevant shocks are the indiosyncratic ones: aggregate shocks affecting all firms are easily told apart from effort, and should not enter the determination of compensation. This generalization therefore reinforces the importance of considering idiosyncratic shocks to the firm.

[3] Papers belonging to a first strand of literature regress individual wages on measures of aggregate profits. Blanchflower et al. (1996) and Estevão and Tevlin (2000) use industry-wide profits drawn from the NBER productivity database, while Christofides and Oswald (1992) use Canadian data. Other studies rely on firm-specific data. Abowd and Lemieux (1993) and Currie and McConnell (1992) use labor union contracts and regress collectively bargained wages on firm profits. Bargained wages, however, exclude bonuses and

the role of risk diversi…cation, the use of aggregate data is problematic for a convincing test of the theory. Aggregate shocks are common to all …rms and therefore undiversi…able, regardless of technology. Besides, …nding a positive correlation between measures of aggregate pro…tability and aggregate (or even disaggregated) wages may just be driven by the equilibrium response of wages to shift in the market demand for labor, and be completely unrelated to insurance considerations.

This paper casts the wage insurance test in the correct framework: within the …rm. We rely on linked employer-employee longitudinal data with enough information to compute measures of shocks both to the …rm and to its employees' compensation.[4] To obtain such information, we merge company-level data for a large sample of Italian …rms with social security data available for a random sample of their employees. Our data o¤er a unique opportunity to test risk allocation between …rms and workers because they allow us to isolate the idiosyncratic shocks that …rms and workers face, an essential step in any convincing test of the insurance hypothesis. We ask whether and how much workers' wages respond to shocks that are speci…c to the …rm and thus are, at least in principle, insurable. The empirical task we confront is how much risk shifting is realized (if any), and how this varies with the type of shock and with observable characteristics of the …rm or of its employees.

Like most European countries, Italy is characterized by a high degree of union coverage. While the hypothesis of wage insurance examined by Azariadis (1975) and others is cast in the context of a competitive labor market with incomplete credit and insurance markets, subsequent work has shown that implicit wage insurance not only arises in bargaining models with unions but also that unions may play an important role in making implicit contracts feasible. Riddell (1981) studies the case of Nash bargaining under uncertainty

other components of pay which constitute an important portion of wage variability. Nickell and Wadhwani (1991) and Hildreth and Oswald (1997) regress …rm-speci…c average wages on …rm pro…tability. While this is a step toward a more disaggregated approach, measures of …rm-level wages do not allow to fully control for individual workers' characteristics. For example, changes in wages might be related to changes in pro…ts because both re‡ect a change in the composition of the …rm's workforce, rather than risk sharing.

[4] The recent development of worker-…rm matched data has allowed researchers to better control for heterogeneity on both sides of the labor contract (Abowd et al., 1999). A …rst attempt to use matched data to test the relation between …rm performance and wage is by Bronars and Famulari (2001), who use a two-year survey of workers and …rms to look at the wage-pro…t relationship in the US. Due to the fact that their data set is limited in cross-sectional size and time length, however, they cannot properly distinguish idiosyncratic from market related risk.

theoretically and shows that as long as bargainers have di¤erent risk attitudes they will be better o¤ reaching an ex-ante agreement in the spirit of the implicit contract literature. Thus, wage insurance is not a peculiar characteristic of competitive markets. Malcomson (1983) considers a union-…rm bargaining model under uncertainty where the surplus arising from the production process is unknown ex-ante because of demand or technology shocks. As in Riddell (1981), state-contingent contracts that reallocate risk between a …rm and its employees are welfare-improving. But these contracts face an enforcement problem in that neither workers nor courts can observe the true state of the world (Grossman and Hart, 1981). Unions can mitigate this problem because they have the necessary power to enforce an (implicit) agreement: they provide workers with more accurate information about the true state of the world and can punish a cheating …rm by depriving of members' labor services (i.e., calling a strike). Unions can therefore enforce state-contingent contracts in an environment where courts cannot. The widely documented ine¢ ciency of the Italian judicial system (Guiso et al., 2000) conforms well with this description. Thus, unions should be seen not as an impediment to risk sharing contracts, but as a way to implement them.

Our empirical analysis takes into account two important generalizations of the basic implicit contract model discussed above. First, to test whether wages are more likely to be insured against transitory than permanent shocks to the …rm's performance, we allow wages to respond di¤erently to more and less persistent shocks to the …rm, consistently with Gamber (1988). Insofar as both types of shocks are present, ignoring the distinction may bias the results towards the full-wage-insurance hypothesis. To separate the response of wages to transitory and permanent shocks we propose a novel identi…cation strategy that can also be applied to analogous problems arising in di¤erent areas of research. Second, we test whether the degree of insurance varies systematically with …rms' and workers' characteristics, as implied by wage contracting models with …nancial market imperfections and moral hazard.

Our …ndings contradict the simple full-insurance paradigm. We …nd that while …rms fully insure workers against transitory shocks they o¤er only partial insurance with respect to enduring shocks. Nevertheless, a simple calculation of the social value of wage insurance shows that wages are remarkably well insulated against …rm shocks. Quantitatively, a permanent 10 percent change in …rm performance induces little less than a 1 percent permanent variation in earnings for those employed at the same …rm on a continuing basis. Moreover, the sensitivity of workers' wages to permanent shocks to the …rm varies systematically with …rm and worker attributes. In particular, it is negatively correlated with workers' risk aversion and overall …rm performance variability, while it increases with the probability of bankruptcy. These …ndings are consistent with the generalizations of the basic wage insurance model.

The rest of the paper proceeds as follows. In Section 2 we review the institutional aspects of wage determination in Italy. We show that, although a large component of workers' wages is determined through centralized bargaining, a signi…cant part is decided at the company level and thus potentially responsive to …rm idiosyncratic shocks. In Section 3 we review the insights of the wage insurance hypothesis. In Section 4 we characterize our empirical approach to the problem, considering a stochastic speci…cation for …rm performance and workers' earnings. We show that, in the spirit of the wage insurance hypothesis, a set of orthogonality conditions obtains that can be used to answer a number of empirically relevant questions. In particular, one can examine whether shocks to …rms' performance are passed on to wages, and to what extent this is a¤ected by whether the shock is transitory or permanent. This section concludes with a discussion of the identi…cation strategy. Section 5 describes the matched …rm-worker data set used in the empirical analysis, and Section 6 presents the estimation of the stochastic model of …rm performance and workers' earnings. The empirical risk-shifting results are presented and discussed in Section 7. At each step of our empirical analysis we present a variety of tests that check the robustness and validity of our assumptions. Section 8 concludes.

2 Institutional background

The Italian industrial relations system is characterized by a multi-tier bargaining process based on national-, industry- and company-level agreements. Collective contracts signed by the three major trade unions (CGIL, CISL and UIL) have erga omnes validity, e.g., they apply to all workers covered by the agreement independently of union membership. The relevant tiers for wage formation are at the industry- and company-level, with national-level bargaining dealing mainly with overall aspects of employment regulation, such as safety and employment protection rules. Contracts at the industry level are signed every three years and deal with the determination of the minimum wage for the various quali…cation levels. Additional components of the compensation package are determined at the company level. Bargaining at the company level involves the …rm's and the workers' representatives. The …rm can decide on some components of the compensation unilaterally; moreover, it can also sign with the unions a …rm-level contract, with provisions on both wage and non-wage aspects of the employment relation. Firm level contracts are not required and there is no provision on their duration.

The relevance of the …rm for wage determination has been evolving with industrial relations.[5] Our data span the 1982-94 period, which is characterized by a fairly high degree of decentralization of the bargaining process.[6] For this period, the wage bill can be decomposed into the following components:[7]

1. Minimi tabellari (contractual minimum), established at the industry level.

2. Indennità di contingenza (wage indexation provision), added to the contractual minimum according to the in‡ation rate.

[5] Until the early 1960s, company-level bargaining was not formally recognized. The economic boom of the 1960s encouraged the establishment of company-level bargaining, at that time mainly focusing on the topics of wages and productivity and essentially autonomous vis-à-vis industry-wide agreements. The 1970s were the period of maximum development of company-level agreements. Starting in the late 1970, the growing economic crisis and the assumption of responsibility by the unions for the unemployment problems led to a gradual reshaping of company-level bargaining. A major restructuring of the industrial relation system occurred in 1993, following the exchange rate devaluation and the severe recession ensuing. Given that our data cover up to 1994, the 1993 agreement plays essentially no role and the industrial relations regime we have to consider is the relatively autonomous one of pre-1993.

[6] Iversen (1998) constructs an index of centralization of wage bargaining which combines a measure of union concentration with a measure of the prevalent level of bargaining. The index covers 15 OECD countries from 1973 to 1995. He divides his sample into three groups: centralized (Norway, Sweden, Denmark, Austria, Finland), intermediately centralized (Netherlands, Germany, Belgium, Japan, Switzerland), and decentralized (Italy, UK, France, USA, Canada), ranked in order of degree of centralization. The index ranges between 0.071 (USA and Canada) and 0.538 (Norway), with Italy having a value of 0.179. For comparison, the UK has a value of 0.177, France of 0.121, and Switzerland of 0.25.

[7] See Erickson and Ichino (1995) for further details on wage formation in Italy for the period covered by our data.

3. Superminimum (wage premia), decided at the company level. It adds to the contractual minimum and has a permanent character (in nominal terms). The superminimum has a …rm-level component and a worker-level component.

4. Premi di produzione (production premiums), determined at the …rm level. These are bonuses and other one-time payments, decided unilaterally by the …rm without formal negotiations with the trade unions. They have no permanent character.

5. Retribuzione variabile (variable compensation), determined by a …rm level contract. It can introduce a contingent component in the compensation package.

The relevance of …rm-speci…c e¤ects for wages which we study in this paper depends on the di¤usion of company level bargaining and on the quantitative importance of the …rm-speci…c components of the wage bill. In terms of di¤usion of …rm-level contracts, the yearly report of CESOS, an association of trade unions, indicates that approximately half of the workers were interested in the signing of a …rm-level contract each year in the period 1984-94.[8] The likelihood of a …rm-level contract increases with …rm size (Bellardi and Bordogna, 1997). As we shall see, in our sample large …rms are disproportionately represented, implying a greater relevance of …rm-level contracts than for the average …rm in the population.

Data that decompose the wage bill into its various components are not available for the economy as a whole. The dataset that has been most extensively used in the literature to address this issue uses wage formation data in the Metal products, Machinery and Equipment sector, assembled by Federmeccanica, the association of employers for that sector. While o¤ering a partial view, these data provide insights that are likely to extend to the economy as a whole, both because of the importance of this industry in the Italian economy and because the patterns observed here are induced by institutional characteristics fairly common across industries (Rossi and Sestito, 2000). Table 1 reports the decomposition of the average wage in the …ve components discussed above for the period 1984-1994 (ap-

[8] The 1994 Bank of Italy survey on manufacturing …rms with at least 50 employees …nds that more than 92% of workers were covered by a …rm-level contract in addition to the industry-wide wage and 40% had been interested in a contractual round in the year (information on previous years is not available).

proximately the same period covered by our data, which span the 1982-94 period). The table shows that the sector level component of the wage determination declines over time from 84% in 1984 to 77% in 1994. This means that between one-sixth and one-quarter of the wage bill is firm-specific, with a quantitatively important role for firms to influence the wages of their employees over and above industry-level bargaining.

Taken together, both the institutional setting and the description of the wage formation process indicates that an important component of the compensation of employees is determined at the level of the firm. This begs the question of how much insurance, if any, the firm provides to its employees vis-à-vis firm-specific shocks.

3 The wage insurance hypothesis

To provide a framework for the subsequent empirical analysis, consider the following stylized implicit contract model. The firm faces output uncertainty due to demand or technology shocks. There are S possible states of nature. The associated probabilities are $p(s)$ $(s = 1, 2, ..., S)$. The firm chooses a wage schedule $w(s)$ to maximize shareholders' utility, $V(.)$, solving:

$$\max \sum_{s=1} p(s) V(\pi(s))$$

subject to:

$$\sum_{s=1} p(s) U(w(s)) - U(b) \geq 0$$

and $\pi = y - w$, where π are profits, y (stochastic) value added, w wages, b unemployment benefits and $U(.)$ workers utility. Letting λ denote the Lagrange multiplier associated to the first constraint and denoting partial derivatives with a prime, the first order condition is:

$$-V'(\pi(s)) + \lambda U'(w(s)) = 0$$

for all s. Thus, for any two states of nature s_1 and s_2

$$\frac{U'(w(s_1))}{U'(w(s_2))} = \frac{V'(\pi(s_1))}{V'(\pi(s_2))} \tag{1}$$

Taking total differentials and using (1) yields:

$$\frac{d\ln w}{d\ln \pi} = \frac{\frac{V''(\pi)}{V'(\pi)}\pi}{\frac{U''(w)}{U'(w)}w} = \frac{r^{\pi}}{r^{w}} \tag{2}$$

Equation (2) shows that the elasticity of earnings with respect to profits depends on the ratio of the firm's (r^{π}) to the worker's risk aversion (r^{w}). It is plausible to assume that firms are not more risk averse then workers, i.e. $r^{\pi} \leq r^{w}$. Under this assumption, $0 \leq \frac{d\ln w}{d\ln \pi} \leq 1$. If the firm is risk neutral, $V''(.) = 0$ and a full insurance result obtains; when workers and firms are equally sensitive to risk, $r^{\pi} = r^{w}$ and shocks to the firm's output are transferred equally to wages and profits (we refer to this as the no-insurance case); in the intermediate cases, $0 < \frac{d\ln w}{d\ln \pi} < 1$ and partial insurance obtains.

As we explain below, our preferred measure of performance is value added $y = \pi + w$. After some algebra, the elasticity of wages to value added can be expressed as follows:

$$\frac{d\ln w}{d\ln y} = \frac{y}{w}\frac{\frac{d\ln w}{d\ln \pi}}{\frac{d\ln w}{d\ln \pi} + \frac{\pi}{w}}$$

Using equation (2), we obtain:

$$\frac{d\ln w}{d\ln y} = \frac{r^{\pi}}{\frac{w}{y}r^{\pi} + \left(1 - \frac{w}{y}\right) r^{w}}$$

where $\frac{w}{y}$ is the fraction of value added that goes to labor. It is immediate to show that $\frac{d\ln w}{d\ln y}$ lies in the same boundary as $\frac{d\ln w}{d\ln \pi}$.

Similar results obtain in a bargaining model under uncertainty, where a union and a single firm bargain over wages.[9] Since unions are potentially able to remove informational

[9] In the bargaining case, and with τ denoting the bargaining power of the union, the objective function is:

$$\max \left[\sum_{s=1}^{S} p(s)\,U(w(s)) - U(b)\right]^{\tau} \left[\sum_{s=1}^{S} p(s)\,V(\pi(s))\right]^{1-\tau}$$

subject to $\pi = y - w$. The first order condition is:

$$\frac{\tau}{1-\tau}\frac{U'(w(s))}{V'(\pi(s))} = \frac{\left[\sum_{s=1}^{S} p(s)\,U(w(s)) - U(b)\right]}{\left[\sum_{s=1}^{S} p(s)\,V(\pi(s))\right]}$$

asymmetries, they ease the implementation of long-term implicit contracts. Our goal in this paper is to demonstrate the potential importance of wage insurance against firm shocks, not to distinguish among alternative models that induce it.

We generalize this simple theoretical structure in the empirical analysis. First, we assume that shocks to value added can be transitory or permanent, and estimate the sensitivity of wages to shocks of different nature. Gamber (1988) shows that in the presence of implicit contracts the real wage responds more to permanent shocks than to transitory shocks if the firm is subject to bankruptcy constraints. Second, to account for unobserved heterogeneity in one's reservation wage, we allow for shocks to wages unrelated to firm performance. Finally, we add firm- and match-specific random unobserved components to the equations that describe firm profits and workers' wages, respectively.

4 The stochastic structure of firms' performance and workers' earnings

In this section we provide a statistical characterization of the evolution of firm performance and wages. We measure firm performance with the value added. We prefer value added to profits for two reasons. First, value added is the variable that is directly subject to stochastic fluctuations. Second, firms have discretionary power over the reporting of profits in balance sheets. This makes profits less reliable as an objective measure of firm performance.

4.1 Firm performance

We model firm performance according to the following process:

$$A(L;p)\, y_{jt} = Z'_{jt}\theta + f_j + \varepsilon_{jt} \tag{3}$$

i.e.,

$$\frac{U'(w(s_1))}{U'(w(s_2))} = \frac{V'(\pi(s_1))}{V'(\pi(s_2))}$$

for any two states s_1 and s_2. Taking differentials yields (2) as in the implicit contract model.

where j and t are subscripts for the j-th firm at time t, y_{jt} is a measure of observed firm performance (the logarithm of value added), $A(L;p)$ a lag polynomial of order $p \geq 0$ (i.e. $A(L;p)x_{j,t} = \sum_{\tau=0}^{p} \alpha_\tau x_{j,t-\tau}$, with $\alpha_0 \equiv 1$), Z_{jt} a vector of firm characteristics, f_j a firm fixed effect, and ε_{jt} a stochastic disturbance. We include lags of value added to capture predictable dynamics (e.g., pre-committed sales). The role of Z_{jt} is to control for non-idiosyncratic shocks (i.e., aggregate-, geographical- and industry-specific shocks). The dynamic structure of ε_{jt} is an empirical issue. As we show in Section 6, a careful data-consistent representation of the stochastic component of firm performance has the following structure:

$$\varepsilon_{jt} = \zeta_{jt} + v_{jt} \tag{4}$$

$$\zeta_{jt} = \zeta_{jt-1} + u_{jt} \tag{5}$$

Equation (4) decomposes the disturbance into a permanent component ζ_{jt} which follows the random walk process (5) with innovation u_{jt}, and a transitory component v_{jt} which is serially uncorrelated. Permanent shocks u_{jt} may capture non-mean-reverting unanticipated technological changes, changes in management or changes in the organizational structure of the firm, while mean-reverting transitory shocks are more likely to be associated with fluctuations in demand (price shocks). It is assumed that the firm can distinguish between transitory and permanent shocks.[10] While partly neglected in the theoretical literature, the distinction between persistent and transitory shocks is important from the point of view of the optimal wage contract. On the one hand, it may be optimal for a risk-neutral firm to insulate workers from transitory fluctuations in output; on the other hand, it is less obvious that the firm will be prepared to supply insurance against permanent shocks.

To simplify subsequent notation we assume covariance stationarity, so that $E\left(u_{jt}^2\right) = \sigma_u^2$, and $E\left(v_{jt}^2\right) = \sigma_v^2$ for all t.[11] We assume that the two shocks v_{jt} and u_{jt} are serially

[10] While this assumption is controversial, it rests on the idea that firms know the origin of the shock and can thus figure out whether it is short-lived or enduring. We also assume that workers have access to the same information, either directly or through unions (see Malcolmson, 1983). On the other hand if neither the firm nor the workers or only the firm but not the workers can discriminate between permanent and transitory shocks, wages will only vary with the total shock and the reaction to permanent and transitory shocks will be the same, an hypothesis we do test in our empirical analysis.

[11] This can be generalized to the case of covariance non-stationarity. All identification results and empirical

and mutually uncorrelated.[12]

4.2 Workers' earnings

Following (2), we assume that workers' pay can be described by the following equation:

$$w_{ijt} = X'_{ijt}\delta + \alpha y_{jt} + \psi_{ijt} \qquad (6)$$

where the subscript i stands for the i-th individual, and w_{ijt} is the logarithm of worker compensation.[13] X_{ijt} denotes a vector of systematic factors that affect individual i's compensation, which can vary across workers, firms and time. Among other things, X_{ijt} includes region, sector, occupation, and year dummies. These dummies remove any variation in wages that is due to industry- and national-level bargaining (including wage indexation). α is the elasticity of wages to firm shocks. Finally, ψ_{ijt} is the stochastic component of earnings unrelated to the firm's fortunes. These idiosyncratic shocks are meant to capture the unobservable component of one's outside wage (individual ability and fluctuations thereof), but also idiosyncratic changes in labor supply (child-raising, family labor supply effects, etc.), and perhaps measurement error. Based again on the evidence presented in Section 6, we model the idiosyncratic earnings shocks as the sum of a permanent random walk component $\eta_{ijt} = \eta_{ijt-1} + \xi_{ijt}$ and a serially uncorrelated transitory shock μ_{ijt}.[14] To save on notation, it is maintained that covariance stationarity holds, i.e. $E\left[\xi^2_{ijt}\right] = \sigma^2_\xi$, and $E\left(\mu^2_{ijt}\right) = \sigma^2_\mu$ for all t. The two shocks are serially and mutually uncorrelated.

Replacing (3) in (6) yields:

$$A(L;p)\,w_{ijt} = A(L;p)\,X'_{ijt}\delta + Z'_{jt}\theta + \alpha\varepsilon_{jt} + \upsilon_{ij} + A(L;p)\,\psi_{ijt} \qquad (7)$$

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findings are available on request.

[12] This structure (and subsequent identification strategy) can be generalized to the case where v_{jt} is serially correlated (for instance it follows an MA(q) process).

[13] We let earnings to depend on contemporaneous firm performance, i.e. assume that wages adjust immediately to changes in performance. In practice, wages might adjust with a lag (think of overtime or bonus decisions, which are usually taken at the end of the calendar year). Nevertheless, if adjustments are made at a frequency higher than the year (say, quarterly), annual data of the type used here will not detect deviations from the contemporaneous adjustment assumption.

[14] In this context, an idiosyncratic shock to earnings is a purely individual innovation, i.e. it is not shared by co-workers within the firm.

where $\tilde{\theta} = \phi\theta$, and ψ_{ij} denotes a firm-worker match-specific unobserved component. Taking first differences of (3) and (7) to eliminate firm- and match-specific random components, and using the stochastic structure outlined above, we obtain:

$$A(L;p)\,\Delta y_{jt} = \Delta Z_{jt}^{0}\theta + u_{jt} + \Delta v_{jt} \tag{8}$$

$$A(L;p)\,\Delta w_{ijt} = A(L;p)\,\Delta X_{ijt}^{0}\delta + \Delta Z_{jt}^{0}\tilde{\theta} + \phi\,(u_{jt} + \Delta v_{jt}) + \vartheta_{ijt} \tag{9}$$

where $\vartheta_{ijt} = A(L;p)\left(\xi_{i;j;t} + \Delta\nu_{i;j;t}\right)$. For the purpose of this paper it is more convenient to define firm performance and earnings growth after adjusting for observable firm and worker characteristics, i.e.:

$$\Delta\varepsilon_{jt} = u_{jt} + \Delta v_{jt} \tag{10}$$

$$\Delta\omega_{ijt} = \phi u_{jt} + \phi\,\Delta v_{jt} + \vartheta_{ijt} \tag{11}$$

where $\Delta\varepsilon_{jt} \equiv A(L;p)\,\Delta y_{jt} - \Delta Z_{jt}^{0}\theta$, and $\Delta\omega_{ijt} \equiv A(L;p)\left(\Delta w_{ijt} - \Delta X_{ijt}^{0}\delta\right) - \Delta Z_{jt}^{0}\tilde{\theta}$. Equations (10) and (11) isolate exogenous shocks to firm performance and wages. In the empirical analysis we replace unexplained growth rates $\Delta\varepsilon_{jt}$ and $\Delta\omega_{ijt}$ with their consistent estimates, obtained as the residuals of the regressions (8) and (9).[15] Note that since we are using panel data, identification of shocks requires a long cross-sectional dimension, not a long time series.

The serial correlation properties of $\Delta\varepsilon_{jt}$ and $\Delta\omega_{ijt}$ are well defined. They follow MA(1) and MA(p+1) processes, respectively. Thus autocovariances of $\Delta\varepsilon_{jt}$ at the second order are all zero; and those of $\Delta\omega_{ijt}$ are zero at order p + 2 and higher. The restrictions on the variance-covariance matrix of $\Delta\varepsilon_{jt}$ are standard:

$$E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right) = \begin{cases} \sigma_u^2 + 2\sigma_v^2 & \text{for } \tau = 0 \\ -\sigma_v^2 & \text{for } |\tau| = 1 \\ 0 & \text{for } |\tau| > 1 \end{cases} \tag{12}$$

[15] A technical requirement for inference to be valid when working with residuals rather than with true disturbances is that fourth moments of both $\Delta\varepsilon_{jt}$ and $\Delta\omega_{ijt}$ exist and are constant across individuals (MaCurdy, 1982).

This simple structure has the obvious advantage that one can identify the variance of the transitory shock and that of the permanent shock to firm performance using only information on the variance and the first-order autocovariances of $\Delta\varepsilon_{jt}$. From equation (12) one can immediately recover σ_u^2 and σ_v^2. Measurement error makes this identification strategy no longer operational; however, as we show later, given the administrative nature of our data, it is reasonable to assume that measurement error is negligible both at firm and at worker level. It is straightforward to show that the presence of classical measurement error in firm data increases the estimate of σ_v^2 but has no effect on that of σ_u^2.[16]

In equation (11) it is implicitly assumed that wages respond equally to transitory and permanent shocks to the firm's performance, i.e. that the Λ coefficient is the same for the two shock components u_{jt} and Δv_{jt}. As seen before, straightforward generalizations of the basic model (Gamber, 1988) indicate that this may not be the case. We thus test whether the sensitivity of wages varies with the temporary or permanent nature of the firm shock. Let α and β denote respectively the different response of wages to permanent and transitory shocks. We can distinguish various insurance regimes depending on the values of α and β. The contemporaneous covariance between shocks to performance and shocks to wage growth has the following structure:

$$E\left(\Delta\varepsilon_{jt}\Delta\omega_{ijt}\right) = \begin{cases} 0 & \text{full insurance} \\ \Lambda\left(\sigma_u^2 + 2\sigma_v^2\right) & \text{homogeneous partial insurance} \\ \alpha\sigma_u^2 + 2\beta\sigma_v^2 & \text{heterogeneous partial insurance} \\ \alpha\sigma_u^2 & \text{transitory full insurance} \\ 2\beta\sigma_v^2 & \text{permanent full insurance} \end{cases} \tag{13}$$

where we have assumed that $E\left(\vartheta_{ijt}u_{j\tau}\right) = E\left(\vartheta_{ijt}v_{j\tau}\right) = 0$ for all t, τ. For simplicity, we have also assumed covariance stationarity. If workers are fully insured against fluctuations in the performance of the firm, the contemporaneous covariance between shocks to performance improvement and shocks to wage growth is zero and full insurance obtains against

[16] If a classical measurement error $r_{jt} \sim i.i.d.\left(0, \sigma_r^2\right)$ is present, it is straightforward to show that $-E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right) = \sigma_v^2 + \sigma_r^2$ for $|\tau| = 1$, but $E\left[\Delta\varepsilon_{jt}\left(\Delta\varepsilon_{jt-1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt+1}\right)\right]$ still identifies the variance of the permanent shock σ_u^2.

shocks of any nature. On the other hand, if workers share part of the fluctuations, without distinguishing between short-lived and durable shocks, equation (13) equals $\delta\left(\sigma_u^2 + 2\sigma_v^2\right)$, where $\delta = \alpha = \beta$. We call this case "homogeneous partial insurance". Three other cases of interest may arise. The optimal contract may result in a different reaction to shocks of different nature. For instance, workers may bear a substantial portion of the firm's permanent shocks but a limited portion of transitory shocks: in this case, which we call heterogeneous partial insurance, the contemporaneous covariance equals $\alpha\sigma_u^2 + 2\beta\sigma_v^2$. Two special cases occur when workers bear only transitory shocks but are insulated from permanent shocks ("permanent full insurance", characterized by $E\left(\Delta\varepsilon_{jt}\Delta\omega_{ijt}\right) = 2\beta\sigma_v^2$) or bear permanent shocks but are insured against transitory shocks ("transitory full insurance", and $E\left(\Delta\varepsilon_{jt}\Delta\omega_{ijt}\right) = \alpha\sigma_u^2$).

An interesting extension would be to allow for asymmetric response of wages to positive and negative transitory and permanent shocks. For instance, one could imagine workers to be insured against downside risk but still take part of the upside movements in value added. Clearly, models that allow for such asymmetric effects are not identifiable. The reason is that we do not separately observe transitory and permanent shocks; it is only (a consistent estimate of) their convolution ($\Delta\varepsilon_{jt}$) that is observable.

4.3 Identification strategy

Without further restrictions, from equation (13) we cannot separately identify α and β, nor can we gauge whether $\alpha = \beta = \delta$. To see how identification of the relevant parameters is achieved, start from the general case where $\alpha \neq \beta$ in (11):

$$\Delta\omega_{ijt} = \alpha u_{jt} + \beta\Delta v_{jt} + \xi_{ijt} \tag{14}$$

Subtract $\beta\Delta\varepsilon_{jt}$ from both sides to obtain:

$$\Delta\omega_{ijt} - \beta\Delta\varepsilon_{jt} = (\alpha - \beta)u_{jt} + \xi_{ijt} \tag{15}$$

Multiply both sides by $\Delta\varepsilon_{jt-1}$ and $\Delta\varepsilon_{jt+1}$, respectively, and take expectations to yield the two orthogonality conditions:

$$E[\Delta\varepsilon_{jt-1}(\Delta\omega_{ijt} - \beta\Delta\varepsilon_{jt})] = 0 \tag{16}$$

$$E[\Delta\varepsilon_{jt+1}(\Delta\omega_{ijt} - \beta\Delta\varepsilon_{jt})] = 0 \tag{17}$$

Intuitively, equations (16) and (17) tell us that once one filters the unexplained component of earnings growth $\Delta\omega_{ijt}$ by the unexplained component of value added growth $\Delta\varepsilon_{jt}$ (weighted by a factor β, the extent of transitory insurance), what is left is uncorrelated with the past and future unexplained component of value added growth. In an OLS regression of $\Delta\omega_{ijt}$ on $\Delta\varepsilon_{jt}$ the latter is obviously endogenous because correlated with the right hand side of equation (15) via u_{jt}.[17] However, the first lag and lead of $\Delta\varepsilon_{jt}$ will be valid instruments, because correlated with $\Delta\varepsilon_{jt}$ (via the transitory component) and uncorrelated with the error term. Equations (16)-(17) can be used to identify the first parameter of interest β with one overidentification restriction. This can be tested with standard methods.

Identification of α proceeds along similar lines. Start from (14), subtract $\alpha\Delta\varepsilon_{jt}$ on both sides and multiply them by the term $(\Delta\varepsilon_{jt-1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt+1})$. Taking expectations it yields the orthogonality condition:

$$E[(\Delta\varepsilon_{jt-1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt+1})(\Delta\omega_{ijt} - \alpha\Delta\varepsilon_{jt})] = 0 \tag{18}$$

Equation (18) identifies the second parameter of interest α. Similarly to the moment conditions (16) and (17), the intuition for this is that after filtering the unexplained component of earnings growth $\Delta\omega_{ijt}$ by the unexplained component of value added growth $\Delta\varepsilon_{jt}$ (weighted by a factor α, the extent of permanent insurance), what is left is uncorrelated with an MA(2) term centered in $\Delta\varepsilon_{jt}$ with unity coefficients.[18] Thus one can use

[17] It is worth noting that OLS estimation provides unbiased and consistent estimation if $\alpha = \beta = \ast$. Thus an exogeneity (Hausman) test for $\Delta\varepsilon_{jt}$ can implicitly be used to check whether $\alpha = \beta = \ast$.

[18] To see why this is so, consider equation (14) and rewrite it as:

$$\Delta\omega_{ijt} = \alpha\Delta\varepsilon_{jt} + [(\beta - \alpha)\Delta v_{jt} + \vartheta_{ijt}]$$

In an OLS regression of $\Delta\omega_{ijt}$ on $\Delta\varepsilon_{jt}$ the latter is endogenous because correlated with the error term (the term in square brackets) via Δv_{jt}. However, the variable $(\Delta\varepsilon_{jt+1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt-1})$ is a valid instrument, because correlated with $\Delta\varepsilon_{jt}$ (via the permanent component u_{jt}) and uncorrelated with the error term, as

$(\Delta\varepsilon_{jt-1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt+1})$ as an instrument. By identical logic, any other MA term that contains $(\Delta\varepsilon_{jt-1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt+1})$ is a valid instrument. For instance, $\sum_{\tau=-q}^{1} \Delta\varepsilon_{jt+\tau}$ (for any $q \geq 2$) is a valid instrument as well. It follows that the model can be tested via these additional overidentifying restrictions. In the empirical analysis, we use a set of three instruments (corresponding to $q = \{1; 2; 3\}$). This gives us two overidentifying restrictions.

Note that in (18) and (16)-(17) different instruments identify different parameters, and that instruments that are valid in one equation are not valid in the other. Moreover, the moment conditions derived above are valid regardless of the covariance stationarity hypothesis, which provides a convenient level of generality.

Our identifying assumption is that measurement error is negligible given the administrative nature of our data. What if we relax this assumption? The reader can verify that the presence of a classical measurement error in the unexplained growth of value added (i.e., the fact that the true value obeys the relation: $\Delta\varepsilon^{*}_{jt} = \Delta\varepsilon_{jt} + \Delta r_{jt}$) implies that the IV estimate of β is biased toward zero while that of α is unaffected. If the true β is zero, however, there is no bias. The problem is one of invalid instruments (in equations 16 and 17 $\Delta\varepsilon^{*}_{jt+1}$ and $\Delta\varepsilon^{*}_{jt-1}$ will be correlated with the variable to instrument, $\Delta\varepsilon^{*}_{jt}$, but also with the measurement error Δr_{jt}). An indirect way to check measurement error bias is thus to check whether overidentifying restrictions are rejected in our model.

In our view, the identification strategy proposed in this paper can be usefully applied to analogous problems confronted in other areas of research. For instance, in intertemporal consumption choice models of the type considered by Blundell and Preston (1998), innovations in consumption (the equivalent of $\Delta\omega_{ijt}$ above) are directly related to the stochastic process of income. The popular income process involving permanent random walk plus transitory serially independent component implies that the consumption innovation adjusts fully to permanent income shocks, but only to the annuity value of transitory shocks. With longitudinal data on consumption and income it is possible to identify the different response of consumption to permanent and transitory income shocks using a slightly modified version of our strategy.

$(\Delta\varepsilon_{jt+1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt-1}) = (u_{jt+1} + u_{jt} + u_{jt-1}) + (v_{jt+1} - v_{jt-2})$ as can be checked after some algebra.

The foregoing is a discussion of the identification of the two insurance parameters α and β. To close the circle on identification, we need to identify the variances of the shock to value added growth and the variances of the idiosyncratic component of earnings growth. As far as the former are concerned, we will use the fact that (in the more general case of covariance non-stationarity) the period t variances are identified by the expressions:

$$E\left(u_{jt}^2\right) = E\left[\Delta\varepsilon_{jt}\left(\Delta\varepsilon_{jt+1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt-1}\right)\right] \tag{19}$$

$$E\left(v_{jt}^2\right) = -E\left(\varepsilon_{jt+1}\varepsilon_{jt}\right) \tag{20}$$

and use minimum distance (Chamberlain, 1984) to obtain the estimates of the parameters of interest. We do this by choosing the parameters that minimize the distance between the actual moments and the moments predicted by the restrictions above.

From the autocovariance function of workers' earnings, we can recover the variance of the transitory and permanent idiosyncratic shocks to wages. In the simple case where $p = 0$ in $A(L;p)$, with heterogeneous partial insurance and covariance stationarity, one obtains:

$$E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau}\right) = \begin{cases} \alpha^2\sigma_u^2 + 2\beta^2\sigma_v^2 + 2\sigma_1^2 + \sigma_\xi^2 & \text{if } \tau = 0 \\ -\beta^2\sigma_v^2 - \sigma_1^2 & \text{if } |\tau| = 1 \\ 0 & \text{if } |\tau| > 1 \end{cases} \tag{21}$$

Conditioning on the estimated values of α, β, σ_u^2 and σ_v^2, the remaining two variances can be identified. A slightly more complicated expression can be derived for arbitrary values of p. Again, minimum distance estimation is used to identify the variances. For more technical details see Appendix B; for a more thorough discussion of covariance estimation see Chamberlain (1984).

More complicated version of the risk shifting hypothesis maintain that the insurance parameters depend on a variety of employer and employees characteristics, such as their risk aversion, the presence of bankruptcy risk, etc. To address this issue, we assume that the insurance parameters are functions of a set of observable individual and firm characteristics, namely:

$$\alpha_{ij} = \alpha_0 + \sum_{r=1}^{R} \alpha_r D_{rj} + \sum_{r=R+1}^{R'} \alpha_r D_{rij} + e_{ij} \qquad (22)$$

$$\beta_{ij} = \beta_0 + \sum_{r=1}^{R} \beta_r D_{rj} + \sum_{r=R+1}^{R'} \beta_r D_{rij} + g_{ij} \qquad (23)$$

where it is assumed that $E\left(e_{ij}u^k_{j\tau}\right) = E\left(e_{ij}v^k_{j\tau}\right) = E\left(g_{ij}u^k_{j\tau}\right) = E\left(g_{ij}v^k_{j\tau}\right) = 0$ for all $\tau; k$. Identification of the parameters β_r and α_r ($r = 0; 1; \ldots; R'$) proceeds as in equations (16)-(17), and (18), respectively, with the matrix of covariates being given by $\Delta\varepsilon_{jt}$ and the interactions of this with the exogenous variables D_{rj} (denoting firm characteristics) and D_{rij} (denoting worker-firm characteristics). Identification is achieved using as instruments the original instruments and the interactions with the variables D_{rj} and D_{rij}.

5 The data

We rely on two administrative data sets, one for firms and one for workers. Data for firms are obtained from Centrale dei Bilanci (Company Accounts Data Service, or CAD for brevity), while those for workers are supplied by Istituto Nazionale della Previdenza Sociale (National Institute for Social Security, or INPS). Since for each worker we can identify the firm he/she works for, we combine the two data sets and use them in a matched employer-employee framework.[19] There is a burgeoning empirical literature on the use of matched employer-employee data sets (see Hamermesh, 1999, for an account).

The CAD data span from 1982 to 1994, i.e. a period that comprises two complete business cycles, with detailed information on a large number of balance sheet items together with a full description of firm characteristics (location, year of foundation, sector of operation, ownership structure), plus other variables of economic interest usually not included in balance sheets, such as employment and flow of funds. Balance sheets are collected for

[19] The INPS data set has been used by Casavola et al. (1999) to describe the determinants of pay in the Italian labor markets and by Galizzi and Lang (1998) to test whether quitting patterns depend on outside employment opportunities. The CAD data set has been used by Guiso and Schivardi (1999) to explore the impact of information spillovers on firms' behavior. To our knowledge, the two data sets have not been used jointly.

approximately 30,000 …rms per year by Centrale dei Bilanci, an organization established in the early 1980s jointly by the Bank of Italy, the Italian Banking Association, and a pool of leading banks to gather and share information on borrowers. Since the banks rely heavily on it in granting and pricing loans to …rms, the data are subject to extensive quality controls by a pool of professionals, ensuring that measurement error should be negligible.

INPS provides us with data for the entire population of workers registered with the social security system whose birthday falls on one of two randomly chosen days of the year. Data are available on a continuous basis from 1974 to 1994. We use the data after 1981 for consistency with the timing of the CAD data. The INPS lacks information on self-employment and on public employment (public …rms are also absent in the CAD). As we describe in Appendix A, the INPS data set derives from forms …lled out by the employer that are roughly comparable to those collected by the Internal Revenue Service in the US.[20] Misreporting is prosecuted.

Given that the INPS data set includes a …scal identi…er for the employer which is also present in the CAD data set, linking the employer's records to the employees is relatively straightforward. As in other countries where social security data are available, the Italian INPS data contain some detailed information on worker compensation but information on demographics is scant.

Table 2 reports various descriptive statistics for the …rms (Panel A) and workers (Panel B) present in our sample. From an initial sample of 177,654 …rm/year observations, we end up with a sample of 122,860 corresponding to 17,272 …rms, excluding …rms with intermittent participation (40,225 observations) and those with missing values on the variables used in the empirical analysis (14,569 observations). Since the panel is unbalanced, the …rms in this sample appear from a minimum of one to a maximum of 13 years.

The sample ranges from very small …rms to …rms with almost 180,000 employees, with an average of 194 and a median of 56. As expected, most of the …rms are in the North (74 percent). As for the distribution by industry, …rms in the chemical, metal production

[20] While the US administrative data are usually provided on a grouped basis, INPS has truly individual records. Moreover, in the US earnings records are censored at the top of the tax bracket, while the Italian data set is not subject to top-coding.

and machinery sectors account for more than 40 percent of the final sample. Firms in more traditional productions (textile, food, paper) account for almost 25 percent. Construction and retail trade take another 25 percent. The remaining 10 percent is scattered in the service sectors, which, with a high share of self-employment and small firms, are under-represented in the CAD data set.

Panel B reports sample characteristics for the workers in the 1982-1994 INPS sample. We start with an initial sample of 267,539 worker/year observations (including multiple observations per year for the same worker due to within-firm position change, multiple employers, employer change, etc.) and end up with 130,785.[21] Sample selection was made with the explicit aim of retaining workers with stable employment and tenure patterns. First we excluded those younger than 18 or older than 65 (2,652 observations), circumventing the problem of modelling human capital accumulation and retirement decisions. To avoid dealing with wage changes that are due to job termination (quits or layoffs) or unstable employment patterns, we excluded workers with part-time employment, those who change position during the year, and those with multiple jobs (81,117 observations). For similar reasons, we dropped individuals who worked for less than 12 months (43,750 observations). In this way we isolate the on-the-job aspect of the wage insurance contract, leaving the consideration of changes in the occupational status to future work. Moreover, we keep only individuals with non-zero recorded earnings in all years (105 observations lost) and eliminate some outliers (503 observations).[22] Finally, we eliminate those with missing values on the variables used in the empirical analysis (8,627 observations). Since these selections — particularly those that exclude job- and firm-movers — can potentially affect our results, in Section 7.2 we check the robustness of our findings by retaining these observations.

Our measure of earnings covers remuneration for regular and overtime pay plus non-wage compensation. We compute net earnings using the Italian tax code for the various years and deflate them using the CPI.[23] For workers with intermittent participation we

[21] Additional observations are lost (for both firms and workers) in the empirical analysis given the dynamic nature of most of our estimators.

[22] An observation is classified as an outlier if (a) real earnings are below 500 euro, (2) real earnings are below 3,000 euro and the change in log real earnings is < -2, or (c) real earnings exceed 50,000 euro and the change in log real earnings is > 2.

[23] Results are very similar if we use gross income as a measure of earnings.

treat two strings of successive observations separated-in-time as if they pertained to two different individuals.

Workers in the resulting sample are on average 41 years old in 1991; production workers account for 64 percent of the sample, 35 percent are clericals and about 2 percent managers. Males are 74 percent of our sample and those living in the South 15 percent. Finally, net earnings in 1991 are roughly 13,000 euro on average, with a median of 12,000 euro.[24]

6 Estimation of the stochastic structure of firm performance and workers' earnings

6.1 Firm performance

As a measure of the idiosyncratic shock to firm performance (ε) we use unexplained variation in the logarithm of value added at 1991 prices (deflated by the CPI). The firm's value added corresponds to the volume of the contractible output that remains once intermediate inputs have been remunerated (i.e., the sum of pre-tax profits, wages and perks).

To identify shocks to firm performance we proceed along the lines of Section 3. The first step is to filter out the predictable component. To this end we consider equation (3):

$$A(L;p)\,y_{jt} = Z'_{jt}\theta + f_j + \varepsilon_{jt}$$

where y_{jt} is the log of real value added for firm j at time t. We set $p = 1$. Taking the first difference of the data eliminates the firm fixed effect and yields:

$$\Delta y_{jt} = \rho\Delta y_{jt-1} + \Delta Z'_{jt}\theta + \Delta\varepsilon_{jt} \tag{24}$$

Included in ΔZ_{jt} is a full set of time dummies, sector dummies and location dummies.[25]

[24] These descriptive statistics can be compared to a representative sample of the Italian population of private sector workers drawn from the 1991 Bank of Italy SHIW. We find that demographic characteristics in the SHIW are very similar to those in the INPS (in particular, the proportion of males, production workers, clericals and managers) and average age is the same in the two samples.

[25] Sector dummies are for agriculture and fishery; mining; food and tobacco products; textile and leather products; paper, wood products and publishing; chemicals and petroleum; primary and fabricated metal

Since OLS estimates are inconsistent, we use IV and instrument Δy_{jt-1} using y_{jt-2} and y_{t-3} (Anderson and Hsiao, 1982). The residual from (24) constitutes our measure of output on which the payment to the worker should be made contingent if workers are to share part of the firm risk.[26]

The results of the IV regression are reported in Table 3. Our estimate of ρ is 0.27 with a standard error of 0.02. Region, year and sector dummies are jointly statistically significant. We use the residual of the IV regression above to construct a consistent estimate of $\Delta\varepsilon_{jt}$. A close examination of the estimated autocovariances $E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})$, reported in Panel A of Table 3 pooling over all years, reveals the absence of any large or statistically significant correlation at lags greater than one. This can be tested more formally using the zero restriction test proposed by Abowd and Card (1989) (see Panel B). This is a test that all the autocovariances greater than a given lag are jointly zero. We find that the null that $E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}) = 0$ is overwhelmingly rejected for $\tau \geq 1$ (p-value <0.0001), but not for $\tau \geq 2$ (a p-value of 49 percent) or higher.

The autocovariances of $\Delta\varepsilon_{jt}$ can be analyzed to gain knowledge about the dynamic structure of $\Delta\varepsilon_{jt}$. This is an important step in our paper, because different dynamic structures will imply different characterizations of the optimal wage contract.

The simplest dynamic structure is one where $\Delta\varepsilon_{jt}$ contains only a mean-reverting component, such as some generic ARMA$(p; q)$ process. Since autocovariances decline quite dramatically from order 0 to order 1 and, especially, from order 1 to order 2, and since they are statistically undistinguishable from zero at orders greater than 2, the presence of an AR component of any order can be safely ruled out. With an AR component, in fact, one should observe autocovariances to decline gently over time at a rate defined by the autoregressive parameter(s); and autocovariances would not become zero as rapidly as they do.

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products; machinery and electric/electronics; energy, gas and water; constructions; retail and wholesale trade and hotels; transport and telecommunications; credit, insurance and business services; and other private services (agriculture and fishery is the excluded category). Location dummies are for North, Center and South (Center is the excluded category). By including these dummies we explicitly remove shocks to firm performance pertaining to a given location and a given industrial sector at a point in time, i.e., non-idiosyncratic shocks.

[26] We run the value added regression on a sample of firms with non-missing values for the variables of interest (i.e., value added, year, sector and location), irrespective of whether there are workers to match them with. This ensures that the results for the value added specification are not peculiar to large firms, which are obviously over-represented in the subset of firms with matched workers.

If $\Delta \varepsilon_{jt}$ contains just an $MA(q)$ component, then autocovariances at lags $q+1$ or higher should all be zero. The autocovariances reported in Table 4 appear to be statistically zero for $\tau \geq 1$, implying $\Delta \varepsilon_{jt} \sim MA(1)$. Since an $MA(1)$ in first differences implies an $MA(0)$ in levels, we can conclude that in the absence of additional variance components $\Delta \varepsilon_{jt} = \Delta v_{jt}$, with v_{jt} being serially uncorrelated, or a transitory shock in the parlance of our model.

Most processes, especially at the micro level, contain both mean-reverting and non mean-reverting components. While our first-differencing procedure eliminates any firm-specific fixed effect in the levels, we might still have a random growth component, i.e. $\Delta \varepsilon_{jt} = g_j + \Delta v_{jt}$. The implication of this process is that autocovariance at distant lags are non-zero due to the presence of the random growth component, i.e., $E(\Delta \varepsilon_{jt} \Delta \varepsilon_{jt-\tau}) = \sigma_g^2$ for $\tau \geq 2$. Once more, the fact that autocovariances at order 2 or higher are all close to zero militates against this specification. An alternative characterization of the non mean-reverting component is a random walk process in the levels of the form $\varepsilon_{jt} = v_{jt} + \zeta_{jt}$, with $\zeta_{jt} = \zeta_{jt-1} + u_{jt}$ with u_{jt} being an i.i.d. serially uncorrelated process (a permanent shock in the parlance of our model). In first differences, $\Delta \varepsilon_{jt} = \Delta v_{jt} + u_{jt}$. Assume for the time being that v and u are uncorrelated at all lags and leads. This process would be consistent with the autocovariance structure of $\Delta \varepsilon_{jt}$ in Table 3 because it implies $E(\Delta \varepsilon_{jt} \Delta \varepsilon_{jt-\tau}) \neq 0$ for $\tau = 0$ and $|\tau| = 1$ and $E(\Delta \varepsilon_{jt} \Delta \varepsilon_{jt-\tau}) = 0$ for $|\tau| \geq 2$. Obviously, since $\Delta \varepsilon_{jt} \sim MA(1)$ even in the absence of a random walk in the levels, this is not enough to conclude that there is a random walk component. However, it is possible to distinguish between the null $\Delta \varepsilon_{jt} = \Delta v_{jt}$ and the alternative $\Delta \varepsilon_{jt} = \Delta v_{jt} + u_{jt}$ by noting that under the null $E(\Delta \varepsilon_{jt} (\Delta \varepsilon_{jt-1} + \Delta \varepsilon_{jt} + \Delta \varepsilon_{jt-1})) = 0$, while under the alternative $E(\Delta \varepsilon_{jt} (\Delta \varepsilon_{jt-1} + \Delta \varepsilon_{jt} + \Delta \varepsilon_{jt-1})) = \sigma_u^2$. If we perform this test using the residuals $\Delta \varepsilon_{jt}$ the null is rejected quite overwhelmingly (a p-value of less than 0.01 percent), so we conclude in favor of the existence of a non mean-reverting component random walk in addition to a mean-reverting, serially uncorrelated component. This is the structure that we used above to derive the restrictions implied by different insurance regimes. Overall, the results of this section suggest that the random walk plus serially uncorrelated transitory shock specification is the only reasonable representation of the stochastic component of value added data (conditioning on lagged growth, sector, year and location dummies).[27]

Two extra sources of correlations may arise from: (a) correlation between permanent shocks in two subsequent periods (for example because technological shocks come in waves), and (b) correlation between transitory and permanent shocks. As for (a), let $E(u_{jt}u_{jt-\tau}) = \sigma_{uu_{-1}}$ for $\tau = 1$ and zero otherwise. Then we can prove that this extra source of correlation does not affect our identification strategy for α (the sensitivity of wages to permanent firm shocks), but leads to bias in the identification of β (the sensitivity of wages to transitory shocks) due to the presence of invalid instruments. We can use the test of the overidentifying restrictions to assess whether this is an important issue. As for (b), let $E(u_{js}v_{j\tau}) = \sigma_{uv}$ for $s = \tau$ and zero otherwise. In this case both the identification of α and that of β fail due to the use of invalid instruments. Once more, the test of overidentifying restrictions will signal rejection of the null if this extra source of correlation is present.

6.2 Workers' earnings

For workers' earnings we consider a logarithmic specification of the process (6), in which the first difference of log annual net real earnings is regressed on its first lag and on a set of observable individual attributes: a quadratic in age, education (here proxied by a set of occupation dummies), gender, region dummies, sector dummies and time dummies. As noted above, nominal gross earnings are first transformed into nominal earnings net of taxes and social security contributions (using the rules coded in the Italian tax system at each point in time), and then deflated by the CPI to 1991 prices. We use the available data for all workers rather than just those in the matched sub-sample. First differencing the data removes any individual- and match-fixed component. Recall from Section 4 that under our hypothesis, the length of the AR process for firm performance carries over to the length of the AR process for workers' earnings (with the same coefficient if the autoregressive component of wages is exclusively driven by risk shifting considerations). Moreover, under

[27] Covariances tend to decay rapidly even when estimated on a year-by-year basis. This exercise, however, reveals that a distinctive feature of the data is covariance non-stationarity, in particular around the strong recessionary episode of 1993. This recession was particularly anomalous because it was characterized by a sharp devaluation and a major tax increase. The former was advantageous only for exporters, while the latter bore on all firms. Before 1992, however, stationarity appears to be a reasonable characterization of the data. The full matrix of estimated autocovariances is available on request.

the same hypothesis, $\Delta\omega_{ijt}$ is an $MA(2)$ process if $p = 1$, as we have assumed in the previous section. We estimate the earnings growth equation by IV using w_{ijt-3} and w_{ijt-4} as instruments for Δw_{ijt-1}.

The results are reported in Table 5. The autoregressive coefficient is 0.28 (with a standard error of 0.05), almost the same as in the firm regression case. The growth of earnings is more rapid for males, managers, and increases with age at a declining rate. Year dummies are jointly significant, while the joint significance of region dummies is borderline. Sector indicators are jointly insignificant. These dummies remove any variation in wages that is due to industry-level bargaining or economy-wide inflation through wage indexation.

We use the residual of this regression to construct a consistent estimate of $\Delta\omega_{ijt}$. We calculate the autocovariances of the latter pooling over all years and report the results in Table 6.

A thorough examination of the estimated autocovariances of the unanticipated component of the rate of growth of earnings (Panel A) reveals that covariances are very small at lags greater than one. On average, the autocovariance of order zero is 0.014 and that of order one -0.005.[28] In Panel B we test formally the hypothesis that the autocovariances of order τ and higher are jointly zero for various values of τ. We strongly reject the hypothesis that $\Delta\omega_{ijt}$ is serially uncorrelated. The hypothesis that all the autocovariance of $\Delta\omega_{ijt}$ of order 2 and higher are zero has a borderline p-value of 12 percent. The hypothesis that all the autocovariance of $\Delta\omega_{ijt}$ of order 3 and higher are zero has instead a larger p-value of 19 percent.

As in the firms' case, these autocovariances can be analyzed to gain knowledge about the dynamic structure of $\Delta\omega_{ijt}$. Briefly, our conclusions are similar to those arrived at in the firms' case (all the details are available on request). There is evidence for $\Delta\omega_{ijt}$ containing both a permanent random walk component in the levels and a mean-reverting component which is either serially uncorrelated or MA(1) with a small MA coefficient (to accommodate the low autocovariance of order 2). Note that the transitory and the permanent component

[28] These are much lower than the estimates for the US using the PSID (Meghir and Pistaferri, 2001) and those for Italy using the SHIW (Pistaferri, 2001), perhaps reflecting the fact that measurement error is less of a problem in this data set.

of $\Delta \omega_{ijt}$ could be related to the corresponding firm components (through risk sharing), purely individual-specific (the case of full insurance), or both. These issues are addressed in the next Section, where we estimate the insurance parameters α and β following the procedure described in Section 4.3 and focusing on the matched employer-employee data set.

7 Shocks and insurance: the estimates

The matched data set includes 45,446 individual/year observations for 9,203 workers and 4,691 firms with contemporaneous observations on both $\Delta \omega_{ijt}$ and $\Delta \varepsilon_{jt}$. The number of overall observations declines because in the data-matching process we lose three types of observations: workers in the INPS sample whose firm is not in the CAD sample (for instance, firms that have not applied for credit in the period of interest), firms in the CAD dataset that have no worker included in the INPS sample (recall that the former includes only those born in two randomly selected days of the year), and observations lost due to the dynamics of our estimation procedure. The final dataset is an unbalanced matched panel of firms and workers. The mean number of annual matches (i.e., the number of workers) per firm is 1.5, with a minimum of 1 and a maximum of 63 per year.[29] Table 7 reports characteristics for the firms and the workers in the matched data set. As expected, the major difference with respect to the full sample is average firm size, which is significantly greater. The median number of employees is 107 in the matched sample, compared with 56 in the full sample. Naturally, larger firms have a greater likelihood of being matched with at least one of the workers in the INPS sample. Nevertheless, other characteristics (such as location, industry, and workers' demographic characteristics), are fairly similar in the two samples.

We apply the identification strategy outlined in Section 4.3 to estimate the parameters of interest, α (the sensitivity of earnings to permanent shocks to value added) and β (the sensitivity to transitory shocks). In both cases, our estimating equation consists of an IV regression of $\Delta \omega_{ijt}$ onto $\Delta \varepsilon_{jt}$.

[29] The low number of matches per firm implies that, while potentially interesting, an analysis of the cross-covariance of wage shocks within the firm (i.e., using co-workers' information) cannot be pursued.

As explained in Section 4.3, parameter β is identified using $\Delta\varepsilon_{jt-1}$ and $\Delta\varepsilon_{jt+1}$ as instruments, while parameter α is identified using $\sum_{\tau=-q}^{1}\Delta\varepsilon_{jt+\tau}$ $(q = \{1; 2; 3\})$ as instruments. The overidentifying restrictions are tested with a standard J-statistic (generalized Sargan test). Under the null hypothesis that the model is correctly specified, J is asymptotically distributed χ^2 with as many degrees of freedom as overidentifying restrictions and is robust to heteroskedasticity of unknown form. Low values of J (high p-values of the test) will signal that the model is correctly specified. The power of the instruments in the reduced-form regressions is checked by looking at the p-value of the F-test on the excluded instruments. Finally, as explained in Section 4.3, an exogeneity (Hausman) test for $\Delta\varepsilon_{jt}$ (Davidson and MacKinnon, 1993) is an implicit test for $\alpha = \beta = \Lambda$.

We also comment on the estimates of the variances of transitory and permanent shocks to value added (v_{jt} and u_{jt}, respectively) and of idiosyncratic transitory and permanent shocks to earnings (μ_{ijt} and ξ_{ijt}, respectively). In both cases, we use minimum distance estimation and for simplicity impose covariance stationarity. The resulting estimates can be viewed as unconditional averages of the underlying (changing) variances. But it is possible to allow for non-stationarity and still identify the parameters of interest (for brevity, these are not reported here; they are available on request). Finally, we construct an estimate of the ratio $\frac{\sqrt{\alpha^2\sigma_u^2+\beta^2\sigma_v^2}}{\sqrt{E[(\Delta\omega_{i,j,t})^2]}}$, which informs us on how much wage variability is due to workers sharing the firm's fortunes. This turns out to be a useful way to summarize the evidence.

7.1 Main results

Table 8 shows the results of our exercise. The first result is that workers' wages do reflect shocks to the firm's value added: $E(\Delta\omega_{ijt}\Delta\varepsilon_{jt}) = 0.0019$ with a standard error of 0.0002 (Panel B). Moreover, there is a substantial difference in impact between permanent and transitory shocks: wages do respond to permanent shocks but the hypothesis of full insurance with respect to transitory shocks cannot be rejected (Panel A). The estimated value of β (which measures the sensitivity of workers' earnings to transitory shocks) is economically small (point estimate 0.004) and not statistically different from zero (a standard error of similar magnitude). The estimated value of α (responsiveness to permanent shocks) is

0.0821, more than 20 times higher, with a small standard error of 0.0128.[30] Joint consideration of the point estimates and of the standard errors of the two parameters suggests that $\alpha \neq \beta$. This is confirmed by the result of the exogeneity test conducted on $\Delta\varepsilon_{jt}$. The test statistic displays a p-value below 0.1 percent, which rejects the null $\alpha = \beta = \ast$. More precisely, while we cannot reject the hypothesis of "full transitory" insurance, "full permanent" insurance can be ruled out. The J-test of overidentifying restriction has a p-value well above 10 percent in both cases, which signals that the models are not misspecified. This result also implies that our IV estimates are not subject to measurement error bias (see the discussion in Section 4.3), and to bias arising from either correlation between subsequent permanent shocks or between transitory and permanent shocks (see the discussion in Section 6.1). Instruments' power is not a concern, as is shown by the low p-value of the F-test in the reduced form regressions (this is a test that the excluded instruments are jointly insignificant).

Table 8, Panel B, also reports the estimated value of the relevant moments of the shocks to output and wages. One can notice that they are rather close to those estimated for the full sample (see Tables 4 and 6).[31] Consistently with the estimates of α and β, as seen before we find that the estimate of $E(\Delta\omega_{ijt}\Delta\varepsilon_{jt})$ is positive and statistically significant, while that of $(\Delta\omega_{ijt}\Delta\varepsilon_{jt-1})$ is minuscule and statistically insignificant.

To allow for an evaluation of the amount of insurance involved, we use equally weighted minimum distance methods (EWMD) to estimate the variances of idiosyncratic shocks to value added and (conditioning on these and the estimated insurance parameters α and β) the variances of idiosyncratic shocks to earnings.[32] The estimate of the variance of the permanent shock to value added, σ_u^2, is 0.0247 (with a standard error of 0.0032), while the

[30] There may be some concern due to the fact that we regress wage shocks against firm shocks, which are common to all individuals working in the same firm. Moulton (1986) shows that the effect of common group errors is to produce artificially low standard errors in such regressions. We corrected standard errors assuming that errors are not independent within firm and find that the correction has no dramatic effects: α is estimated with a slightly higher standard error of 0.0159, but the associated p-value is still below 0.01 percent.

[31] For example, $E(\Delta\omega_{ijt}\Delta\omega_{ijt})$ is 0.0136 in the matched sample and 0.0143 in the full sample; $E(\Delta\omega_{ijt}\Delta\omega_{ijt-1})$ is -0.0052 in the matched sample and -0.0053 in the full sample.

[32] An alternative would be to use a generalized least squares procedure (optimal minimum distance, or OMD). Our choice is dictated by the evidence presented in Altonji and Segal (1996), who show that EWMD dominates OMD even for moderately large sample sizes.

estimated variance of the transitory shock, σ_v^2, is 0.0326 (with a standard error of 0.0043). These are both sizeable and imply standard deviations of 16 and 18 percent, respectively.

Next, we estimate the parameters of the idiosyncratic part of the earnings process, i.e., after filtering the variability that is due to the amount of insurance/incentives provided by the firm. Following the discussion in Section 4, we assume that this idiosyncratic part of the earnings process can be written as:

$$\Delta\hat{\omega}_{ijt} \equiv \Delta\omega_{ijt} - \alpha u_{jt} - \beta\Delta v_{jt} = \xi_{ijt} + \theta\xi_{ijt-1} + \Delta\varepsilon_{ijt} + \theta\Delta\varepsilon_{ijt-1} \tag{25}$$

i.e., $\Delta\hat{\omega}_{ijt}$ follows a composite MA(2) process. In part, the coefficient θ will reflect the legacy of the autoregressive process of the value added (see equations 7 and 9); in part, however, it will be related to an idiosyncratic moving average component in earnings. The EWMD-estimated variances of idiosyncratic shocks to wages are smaller than the firm counterpart: σ_ξ^2, the variance of permanent shocks, is 0.0065 (standard error 0.0026), while σ_ε^2 is 0.0029 (s.e. 0.0018). The MA coefficient θ in the stochastic process of earnings is negative (-0.18) but measured imprecisely.

To summarize, our findings imply that a 10 percent permanent change in firm performance induces little less than a 1 percent permanent variation in earnings for those employed at the same firm on a continuing basis. To get a sense of the economic significance of this effect consider the median firm (value added of 3.62 million euro, 107 employees, paying an average salary of 12,502 euro per year). Evaluated at the sample median, a permanent decrease in value added of 362,000 euro (10 percent) - equivalent to a 3,383 euro drop in value added per (initial) worker - would permanently lower the earnings of the continuing workers by 103 euro.

The variability in compensation induced by risk shifting depends on $\alpha\sigma_u$ (ignoring the reaction to transitory shocks, which is virtually zero, both economically and statistically), as can be seen from equation (21). Since the overall standard deviation of the shocks to wage growth is 0.1166, one can infer that little more than 11 percent of the total earnings variability can be explained by firm-specific risk (see the last row of Table 8), while the remaining component is related to worker-specific shocks.[33]

As a rough gauge of the social value of the firm as an insurance provider, suppose no devices (other than the firm) were available to smooth consumption in the face of wage risk. Adapting to our case the simple characterization suggested by Lucas (1987) to evaluate the welfare costs of recessions, suppose workers maximize $U(c_t) = \frac{c_t^{1-r^w}}{1-r^w}$ and receive stochastic earnings $A\omega_t$, where r^w is the degree of workers relative risk aversion, A is a non-stochastic component and $\ln \omega_t \sim N(0; \Sigma)$, with $\Sigma = \alpha^2\sigma_u^2 + \beta^2\sigma_v^2$. Wage risk depends on α and β, the insurance parameters, and on σ_u^2 and σ_v^2, the variances of firm shocks.

Lucas assumes that the approximate consumption solution is $c_t = (1+\lambda)(1+g)^t e^{-\frac{\Sigma}{2}}\omega_t$ where g is the growth rate of consumption per period and λ a scaling parameter. The proportional increase in consumption required to leave the consumer indifferent between the full insurance consumption path ($\alpha = \beta = 0$) and the partial insurance path ($\alpha > 0$ and $\beta > 0$) is $r^w\frac{\Sigma}{2}$ (Clark et al., 1994). This provides a first, reasonable approximation to the cost of risk sharing. Assuming $r^w = 5$ and $\sigma_u^2 + \sigma_v^2 = 0.0573$ (values consistent with our empirical findings), the proportional increase in consumption that would leave the consumer indifferent between full and no insurance ($\alpha = \beta = 1$) is $r^w\frac{\sigma_u^2+\sigma_v^2}{2} = 0.14$. According to our estimates the consumer is fully insured against transitory shocks ($\beta = 0$) and partially insured against permanent shocks ($\alpha = 0.082$). Assuming a variance of the permanent shock to firm performance $\sigma_u^2 = 0.025$, workers would be indifferent between the insurance the firm currently provides and no insurance if their consumption were raised by little less than 14% ($r^w\frac{\sigma_u^2+\sigma_v^2}{2} - r^w\frac{\alpha^2\sigma_u^2}{2} = 0.14 - 0.00042 = 0.13958$). Starting with an average consumption of 13,929 euro (see Table 7, Panel B), and no wage insurance, a consumer would be willing to sacrifice as much as 1,944 euro in order to obtain the insurance firms provide according to our estimates, and about 6 extra euro to obtain full insurance. In this respect, the firm proves to be a formidable provider of insurance for individuals.

[33] Using French longitudinal matched data, Abowd et al. (1999) conclude that most of the unexplained variation in earnings levels is due to individual, rather than firm effects. We complement their evidence showing that most of the unexplained variation in earnings growth is also due to worker-specific effects.

7.2 Sensitivity analysis

Recall from Section 5 that our sample selection eliminates those who change position within the …rm and those who have multiple employers over the period of observation (including those with multiple jobs in a given year). This was justi…ed by our attempt to focus on workers with stable employment and tenure patterns. Moreover, we use an unbalanced panel of …rms and workers. Since these selections could be problematic, in Table 9 we check that our results are not driven by sample selection.

In column (1) we report the estimates obtained with the baseline sample. In column (2) we include those who change position within a …rm due to promotion, demotion or related events. The sensitivity to permanent shocks is slightly lower, but statistically indistinguishable from that for the baseline sample. The sensitivity to transitory shocks is unchanged.

In column (3) we additionally include those with multiple employers over the sample period. This accounts for people that quit or are laid o¤ and move to a di¤erent employer. Since our objective is to study annual earnings growth within a …rm (not across …rms), we treat employment spells in di¤erent …rms as di¤erent individual-…rm pairs. We correct the standard errors of our estimates by assuming that the error of the wage equation is correlated across observations pertaining to the same individual. The estimates for this sample are, once again, quite similar to those reported for the baseline sample.

In column (4) we focus on a balanced panel of …rms and workers (continuously observed from 1982 to 1994). This checks that the …rms that are in the sample for fewer years are not di¤erent than those that are observed for the entire sample period. The results are the same as in our baseline sample, although the tests of overidentifying restrictions have somewhat borderline p-values.

We conclude that our sample selection is not responsible for the results, in particular the result of low sensitivity of wages to …rm shocks.

7.3 Insurance and …rm-worker characteristics

Once the full insurance hypothesis is rejected, one can ask whether the level of insurance varies systematically with workers' and …rms' characteristics. In this respect, the implicit contract model only o¤ers limited guidance. The simple framework of Section 3 shows that the parameter linking wages to performance decreases with workers' risk aversion and increases with the …rm's (equation 2). Moreover, the model proposed by Gamber (1988), which allows for bankruptcy, indicates that wage insurance decreases with the probability of default, because a higher correlation of wages to performance reduces the probability of being at the corner of a binding bankruptcy constraint. Additional theoretical restrictions stem from extending the wage insurance model. Consider the principal-agent model (Holmstrom and Milgrom, 1987), which constitutes a generalization of the implicit contract model. This class of models analyzes the trade-o¤ between insurance and incentives to exert e¤ort in the presence of moral hazard, and, by bringing additional characteristics of both the principal (the …rm) and the agent (the worker), o¤ers interesting comparative statics predictions. In particular, we consider two implications of the theory that are likely to extend to our framework. First, the principal-agent model predicts that the compensation scheme should be made more dependent on performance the more sensitive the performance to e¤ort. We should therefore expect that the …rm o¤ers less insurance to workers whose e¤ort is more relevant for performance, such as managers. Second, the theory predicts that the link between wages and performance should be stronger the more precise the signal the principal obtains regarding the agent's e¤ort. When the underlying performance of the …rm is noisy, and the signal less precise, we should expect more insurance.[34]

The above discussion suggests four comparative statics exercises, two regarding workers (risk aversion and occupational status) and two the …rm (probability of default and noisiness of performance). As far as worker characteristics are concerned, the data report the occupational status of each worker, so that we can create a dummy for managers. Ideally, we would like to have a measure of risk aversion for each worker in the CAD data set. This

[34] If there is sorting of less risk averse …rms into more volatile sectors, then noise in performance can be interpreted as an indirect measure of …rm risk aversion. Consistently with (2), …rms with more volatile output should provide more insurance.

variable is not available. To classify individuals by risk aversion, we use outside information on a measure of relative risk aversion obtained from the Bank of Italy's 1995 Survey of Household Income and Wealth (SHIW), which collects data on income, consumption and wealth and several demographic variables for a representative sample of about 8,000 Italian households. The 1995 wave of the survey elicits attitudes towards risk. The household head is offered a hypothetical lottery and asked to report the highest price he would be willing to pay to participate, from which a measure of the Arrow-Pratt index of relative risk aversion can be obtained (Guiso and Paiella, 2001). Then, through an imputation procedure based on characteristics observable in both samples (a cubic in age, net real earnings, dummies for firm size, industry, region of residence, occupational status and gender), we obtain a measure of risk aversion of workers in the INPS data set. The details of the construction of the risk aversion indicator are in appendix A.3. The resulting measure is very reasonable and conforms to prior expectations: average relative risk aversion is 5.03 and the median 4.86. We construct an indicator for high risk aversion (an imputed coefficient above the cross-sectional median). Using an indicator dummy should reduce misclassification error due to the imputation procedure.[35]

As for firm characteristics, we use the frequency of defaults at provincial level as a proxy for bankruptcy risk[36] and the historical performance variability as an inverse indicator of the precision of the signal. The variability is measured by the standard deviation of log real value added over the period observed (from a minimum of 5 to a maximum of 13 years).

We modify our IV estimation strategy, allowing the sensitivity coefficients α and β to depend on observable worker and firm characteristics. Thus we estimate by IV the following equation:

$$\Delta \omega_{ijt} = \alpha_{ij} u_{jt} + \beta_{ij} \Delta v_{jt} + \varepsilon_{ijt} \tag{26}$$

[35] Direct use of the imputed risk aversion variable in levels or logs gives qualitatively similar results, although somewhat less precisely measured.

[36] We have two measures of the frequency of default for each province (the Italian territory is divided into 95 provinces) and for each year. The first measure is the number of firms that defaulted on a bank loan divided by the total number of borrowers at the beginning of the period. The second is the value of defaulted bank loans divided by the value of outstanding loans at the beginning of the period. We summarize the information contained in these variables by using factor analysis and extracting the first principal component for each province-year pair (see Greene, 1997, p. 424-27). Qualitatively similar results are obtained using either variable separately or a simple average of them.

where α_{ij} and β_{ij} are defined by (22) and (23), respectively. This amounts to including interactions of such variables with value added growth shocks, $\Delta\varepsilon_{j,t}$, and augmenting the set of instruments by the interactions of the original instruments with the relevant worker and firm characteristics.[37]

Table 10 reports the results. Column (1) shows the effect of worker and firm characteristics on the sensitivity of wages to permanent shocks and column (2) the effect on transitory shocks. To check the power of the instruments excluded in the reduced-form regressions, we report the partial R^2 measure suggested by Shea (1997) in the context of multivariate models with multiple endogenous variables. Standard errors are corrected for province clustering.

We first comment on the results reported in column (1). The indicator for high risk aversion is associated with a statistically significant lower sensitivity of wages to permanent shocks to performance (i.e., more insurance and a lower value of α). Overall, there is a quite sizable sensitivity differential due to risk aversion (-0.06). In the same direction, managers have less insurance than other workers, but standard errors are high and prevent reliable inference, arguably due to the small number of observations on such workers (a little more that 1 per cent of the sample).

In terms of firms, we find that a higher probability of default is associated with less insurance, in line with the theoretical prediction. Moreover, consistently with the predictions of the basic agency model, firms with higher variability in performance provide more insurance: the coefficient is negative (-0.067) and statistically significant. We interpret this as evidence that incentive schemes are less effective the noisier the relation between effort and performance, supporting one of the fundamental implications of the principal-agent theory.

To get a sense of the results contained in Table 10, consider a highly risk-averse production worker employed in a firm with a historical performance variability of 22 percent (the median variability of value added in the matched sample) located in a province with median

[37] Our estimates of α_{ij} and β_{ij} are not affected by the relationship between $\Delta\omega_{ijt}$ and the firm/worker characteristics D_{rj} and D_{rij} in (22) and (23) that may happen to exist in the cross-section. Including main effects has virtually no effect on the estimates of α_{ij} and β_{ij} (results available on request). We also experimented including additional interactions for firm (firm size and a dummy indicator for co-op's) and worker (age, gender, region), but none of these were statistically significant.

bankruptcy rates. For this worker, the sensitivity to …rm permanent shocks is 0.08. For an individual with the same characteristics ¡ except risk aversion ¡ the coe¢cient becomes as high as 0.14. For employees of a …rm with the same characteristics ¡ but a larger standard deviation of, say, 50 percent ¡ the coe¢cient declines to 0.06. For employees of a …rm with similar characteristics ¡ but located in a province at the 75th percentile of the distribution of bankruptcy rates ¡ the coe¢cient is as high as 0.1. In line with the predictions of extended versions of the implicit contract model, changes in worker and …rm characteristics may thus impart a wide range of variability in ®.

Note …nally that the p-value of the J-test does not point to misspeci…cation of the model (51 percent), and that in all cases the power of the instruments (as measured by the partial R^2) is high enough to allow identi…cation of the relevant parameters and to dismiss the possibility of …nite sample bias.

In column (2) we repeat the estimation exercise for the sensitivity of earnings to transitory shocks. In accordance with the results reported in Table 9, neither worker nor …rm characteristics appear to be statistically signi…cant. This implies that insurance of transitory shocks to value added is pervasive, even after conditioning on workers' and …rms' characteristics.

8 Conclusions

We o¤er empirical evidence on the extent of wage insurance within the …rm, based on a matched employer-employee data set for Italy that spans the years from the early 1980s to the mid-1990s. We …nd that while full insurance is provided against temporary shocks, enduring disturbances to output are only partially insured. In addition, the sensitivity of workers' wages to permanent shocks to the …rm varies systematically with …rm and worker attributes. In particular, it is negatively correlated with the worker' risk aversion and the overall variability of …rm performance, and positively with the probability of bankruptcy.

These …ndings are su¢ciently robust for us to draw a few conclusions. First, all workers at least partly share the fortunes of their company, to an extent that depends on their relationship with the …rm and their preferences (risk-averse workers self-select into more secure firms). Second, insurance coverage depends on the nature of the shocks to the firm: it is complete when temporary but only partial when permanent. This obviously helps a firm's adjustment when shocks hit. Ignoring the distinction between short- and long-run shocks and imposing a common coefficient biases the estimate of the insurance parameter towards full insurance if transitory shocks are more likely to be insured than permanent ones (a solid conclusion of our empirical analysis).

While we reject the full insurance paradigm, we find that transitory shocks are not transmitted to wages. Furthermore, compared to the no-insurance benchmark, which implies that wages vary in the same proportion as the firm's value added (see Section 3), little more of one-tenth of the wage variability is due to workers sharing the firm's (permanent) risk. The average standard deviation of wage growth shocks is about 12 percent while that of shocks to value added growth is 35 percent; if temporary shocks were transferred to workers in the same proportion as permanent shocks, the variability of earnings would increase by as much as 20 percent. This, along with a simple calculation of the social value of wage insurance (see Section 7.1), suggests that wages are remarkably well insulated from firm-specific shocks.

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A Appendix: The data

A.1 The INPS data set

The Italian National Institute for Social Security (Istituto Nazionale della Previdenza Sociale) requires …rms to …le a yearly report (form O1M) for each worker on the payroll. The data are used to estimate the amount of withholding tax the employer has to pay on behalf of the employees, and to INPS as contributions towards health insurance and pension funds.

The database covers the universe of employees in the private sector (thus excluding the self-employed, public employees, and o¤-the-books work). Our data set is a sub-sample of the universe, based on workers born on two particular days of the year; the data are available on a continuous basis for the 1974-1994 period. The form reports information on annual earnings and on the number of weeks worked. Information on hours worked are not available. Earnings are divided into two components: normal and occasional. Occasional earnings includes sums drawn from the wage supplementation fund for laid-o¤ or short-time workers, seniority and loyalty premia, one-time bonuses, moving expenses and business travel refunds, the monetary value of goods in kind, and allowances for lost tips and commissions. On average, occasional earnings are less than 10 percent of the total. Our measure of gross income is the sum of the two components.

The data set also has information on job categories, albeit workers with a rough breakdown: apprentices, production workers, clericals and managers. Information on education is not available. From the worker's social security number it is possible to retrieve the gender, the year of birth (and therefore age), and place of birth. Finally, the data set also contains the employer tax code, which allows us to match information on the worker with that for the …rm.

A.2 The CAD data set

Firm data are drawn from the archives of the Italian Company Accounts Data Service, which collects balance sheet information and other items on over 30,000 Italian …rms. The data, available since 1982 and up to 1996, are gathered by Centrale dei Bilanci, an organization established in the early 1980s jointly by the Bank of Italy, the Italian Banking Association (ABI), and a pool of leading banks to build up and share information on borrowers. Besides reporting balance sheet items, the database contains detailed information on …rm demographics (year of foundation, location, type of organization, ownership status, structure of control, group membership, etc.), on employment, and on ‡ow of funds. Balance sheets are reclassi…ed to reduce dependence on the accounting conventions. Balance sheets for the banks' major clients (de…ned according to the level of borrowing) are collected by the banks. The focus on the level of borrowing skews the sample towards larger …rms. Furthermore, because most of the leading banks are in the northern part of the country, the sample has more …rms headquartered in the North than in the South. Finally, since banks mainly deal with …rms that are creditworthy, …rms in default are not in the data set, so that the sample is also tilted towards better than average quality borrowers. Despite these biases, comparison between sample and population moments (not reported) suggests that the CAD is not too far from being representative of the whole population (with the exception of the over-representation of …rms larger than 1,000 employees).

A.3 Constructing workers' risk aversion

To classify individuals by risk aversion, we use outside information on a measure of relative risk aversion obtained from the Bank of Italy's 1995 Survey of Household Income and Wealth (SHIW).

The survey collects data on income, consumption and wealth and several demographic variables for a representative sample of about 8,000 Italian households. The 1995 wave of the survey elicits attitudes towards risk. The household head is offered a hypothetical lottery and asked to report the highest price he would be willing to pay to participate. Specifically, respondents are asked the following question:

> "We would like to ask you a hypothetical question that we would like you to answer as if the situation were a real one. You are offered the opportunity of acquiring a security permitting you, with the same probability of 1/2, either to gain 10 million lire or to lose all the capital invested. What is the most that you are prepared to pay for this security?".

Ten million lire corresponds to about 5,000 Euro (little more than $5,000). Interviews are conducted personally at home by professional interviewers, who to help respondents understand the question show an illustrative card and are ready to provide explanations. The respondent can answer in one of following three ways: a) declare the maximum amount he or she is willing to pay to participate; b) don't know; c) unwilling to answer. Following Guiso and Paiella (2001), we use the answers to obtain a measure of the Arrow-Pratt index of relative risk aversion for each consumer. Let Z_i be the maximum amount consumer i is willing to pay to enter the lottery; c_i the endowment, x equals 10 million lire, and U_i is the utility function. The maximum price for participating in the lottery is then defined by:

$$EU_i(c_i) = \frac{1}{2}U_i(c_i + x) + \frac{1}{2}U_i(c_i - Z_i) \tag{27}$$

Next, we construct a SHIW sample that is comparable to the INPS sample (people aged 18 to 65, neither self-employed nor working in the public sector), and run a regression of the coefficient of relative risk aversion on attributes that are observed in both data sets: a cubic in age, net real earnings, dummies for firm size, industry, region of residence, occupational status and gender. The R^2 of the regression is 0.2. We retrieve the estimated coefficients and use them to impute the relative risk aversion of all the workers present in the INPS/CAD matched data set. The resulting measure is very reasonable and conforms to prior expectations: average relative risk aversion is 5.03 and the median 4.86. The index ranges from 1.79 to 20.64. Our SHIW sample includes 1,919 workers with valid answers to the risk aversion question. The sample distribution of the degree of relative risk aversion is right-skewed with a median of 5.35; its value ranges from 0.005 to 36.26 but 90 percent of the cross-sectional distribution is comprised between 1.5 and 12.6.

B Appendix: Covariance estimation

For each firm in the sample we obtain a consistent estimate of $\Delta\varepsilon_{jt}$ as the residual from the IV regression (8). For an unbalanced sample of firms observed for at most T periods, define the vector:

$$\Delta\varepsilon_j = \begin{pmatrix} \Delta\varepsilon_{j1} \\ \Delta\varepsilon_{j2} \\ \dots \\ \Delta\varepsilon_{jT} \end{pmatrix}$$

If the $\Delta\varepsilon_{jt}$ observation is missing, it is replaced by zero. Conformably with $\Delta\varepsilon_j$, define with d_j a vector of 0-1 dummy variables. The dummy is 0 if the observation for $\Delta\varepsilon_{jt}$ is missing, 1 otherwise.

This notation allows us to handle the problem of unbalanced panel data in a simple manner. All the autocovariances of the type $E(\Delta\varepsilon_{js}\Delta\varepsilon_{jt})$ are consistently estimated by the sample analogs collected in the following autocovariance matrix:

$$C = \sum_{j=1}^{F} \Delta\varepsilon_j \Delta\varepsilon_j' \oslash \sum_{j=1}^{F} d_j d_j'$$

where F is the number of firms present in the data set and $\oslash$ denotes an elementwise division.

Define with m the vector of all the distinct elements of C, i.e. $m = \mathrm{vech}(C)$. Since C is a symmetric matrix, the number of distinct elements in it is $\frac{T(T+1)}{2}$. Conformably with m, define $m_j = \mathrm{vech}\left[\Delta\varepsilon_j \Delta\varepsilon_j' \oslash d_j d_j'\right]$. The standard errors of the $\frac{T(T+1)}{2}$ autocovariances can be retrieved by the variance-covariance matrix of C, i.e.:

$$V = \sum_{j=1}^{F} \left[(m_j - m)(m_j - m)' \odot \left(d_j d_j'\right)\right] \oslash DD'$$

where $D = \mathrm{vech}\left(\sum_{j=1}^{M} d_j d_j'\right)$ and $\odot$ denotes an elementwise product. The standard errors of the estimated moments are simply the square roots of the elements in the main diagonal of V.

In Panel B of Table 8 we estimate models for m:

$$m = f(\theta) + \upsilon$$

where υ captures sampling variability and θ is the vector of parameters we are interested in. We solve the problem of estimating θ by minimizing:

$$\min_{\theta} (m - f(\theta))' A (m - f(\theta))$$

where A is a weighting matrix. Optimal minimum distance (OMD) imposes $A = V^{-1}$, while equally weighted minimum distance (EWMD) imposes $A = I$. For inference purposes we require the computation of standard errors. Chamberlain (1984) shows that these can be obtained (for the EWMD case) as:

$$\mathrm{var}\left(\hat{\theta}\right) = (G'G)^{-1} G'VG (G'G)^{-1}$$

where $G = \left.\frac{\partial f(\theta)}{\partial \theta}\right|_{\theta = \hat{\theta}}$ is the Jacobian matrix evaluated at the estimated parameters $\hat{\theta}$.

The test of zero restrictions reported in Tables 4 and 6 is described in Abowd and Card (1989). The test statistic has the form $m_2' V_{22}^{-1} m_2$, where m_2 is the sub-vector of m (the vector of estimated moments) assumed to be zero under the null and V_{22} the associated covariance matrix. The statistic is distributed χ^2 with degrees of freedom equal to the dimension of m_2.

The strategy we follow in the workers' case is similar and therefore omitted.

Table 1
Wage bill decomposition for the Machinery industry, 1984-94

Each entry represents the percentage contribution to the wage determination. The …rst column is the contractual minimum, the second the indexation component, the third the sum of the two, which constitutes the industry-wide component of the wage. The remaining columns represent the …rm level components: the superminimum is the wage premium above the contractual minimum, the production premiums are bonuses and other one-time payments, the variable compensation is determined by …rm-level contracts, the last column in the sum of the …rm level components. The data source is Federmeccanica (the association of employers for the Metal products, Machinery and Equipment sector).

Year	Contr. Min. [1]	Indexation [2]	Industry Share [1]+[2]	Supermin. [4]	Production premium [5]	Variable component [6]	Firm share [4]+[5]+[6]
1984	32.4	51.9	84.3	13.0	2.7	0.0	15.7
1985	32.1	51.2	83.3	14.2	2.5	0.0	16.7
1986	30.1	51.3	81.4	15.5	3.1	0.0	18.6
1987	31.2	50.0	81.2	15.8	3.0	0.0	18.8
1988	30.4	47.5	77.9	17.3	3.1	1.7	22.1
1989	29.5	47.8	77.3	16.8	3.4	2.5	22.7
1990	28.1	48.3	76.4	17.9	3.3	2.4	23.6
1991	30.3	48.0	78.3	16.6	3.0	2.1	21.7
1992	31.0	46.8	77.8	17.2	3.0	2.0	22.2
1993	32.7	45.2	77.9	17.3	3.0	1.8	22.1
1994	32.5	44.6	77.1	18.2	2.9	1.8	22.9

Table 2
Firms' and workers' characteristics in the full sample

Panel A reports summary statistics for the …rms in our data set; panel B shows descriptive statistics for the sample of workers. All statistics refer to 1991.

Panel A: Firms' characteristics

	Mean	Median	Stand. dev.
Value added (million euro)	8.27	1.91	123.02
Employees	194	56	2,281
South	0.0917	0	0.2887
Center	0.1643	0	0.3706
North	0.7439	1	0.4365
$\frac{\text{Number of …rms defaulting}}{\text{Number of …rms borrowing}}$ (province level)	0.0228	0.0219	0.0091
$\frac{\text{Value of defaulted loans}}{\text{Value of outstanding loans}}$ (province level)	0.0301	0.0190	0.0285
Agriculture and Fishery	0.0027	0	0.0519
Mining	0.0062	0	0.0783
Food and tobacco products	0.0470	0	0.2116
Textiles and leather products	0.1211	0	0.3262
Paper, wood products and publishing	0.0933	0	0.2909
Chemicals and petroleum	0.1297	0	0.3360
Primary and fabricated metal products	0.1086	0	0.3111
Machinery and electrical/electronic	0.1920	0	0.3939
Energy, gas and water	0.0028	0	0.0529
Construction	0.0778	0	0.2679
Retail and wholesale trade, hotels	0.1573	0	0.3641
Transport and telecommunications	0.0257	0	0.1582
Credit, insurance and business services	0.0183	0	0.1341
Other private services	0.0176	0	0.1314

Panel B: Workers' characteristics

	Mean	Median	Stand. dev.
Earnings (euro)	13,363	12,086	6,004
Age	40.83	41	9.69
Male	0.7454	1	0.4357
Blue Collars	0.6275	1	0.4835
Clericals	0.3556	0	0.4787
Managers	0.0169	0	0.1289
South	0.1446	0	0.3517
Center	0.1837	0	0.3872
North	0.6717	1	0.4696

Table 3
The value added regression

This table reports the results of an IV regression for value added growth at time t. Instruments include the log of value added dated $t-2$ and $t-3$. Values in round brackets are standard error. For region, year and sector dummies F statistics are reported; values in square brackets are p-values.

Variable	Estimate
Value added growth at $t-1$	0.2672 (0.0220)
Region dummies	8.13 [0.0003]
Year dummies	30.54 [< 0.0001]
Sector dummies	5.31 [< 0.0001]
Number of observations	77,770

Table 4
The autocovariance structure of shocks to value added

Panel A reports the estimates and the corresponding standard errors of the autocovariances at various orders of the residual of value added growth in first differences, i.e., estimates of $E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})$. The data are pooled over all years. Panel B tests the null hypothesis that all the autocovariances $E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})$ are jointly insignificant for different values of τ.

Panel A

Order	All years	Order	All years
0	0.1196 (0.0031)	5	0.0007 (0.0008)
1	−0.0428 (0.0021)	6	0.0002 (0.0010)
2	0.0007 (0.0008)	7	−0.0026 (0.0012)
3	−0.0002 (0.0008)	8	0.0008 (0.0011)
4	0.0007 (0.0009)	9	−0.0040 (0.0023)

Panel B

$E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})=0$	$\forall\tau\geq1$	$\forall\tau\geq2$	$\forall\tau\geq3$	$\forall\tau\geq4$
χ^2 [p-value] (d.o.f.)	292.11 [< 0.0001] (45)	35.49 [0.4927] (36)	27.76 [0.4774] (28)	20.93 [0.4630] (21)

Table 5
The earnings equation

This table reports the results of an IV regression for earnings growth. Instruments include the log of earnings dated $t-3$ and $t-4$. Values in round brackets are asymptotic standard errors. For region, year and sector dummies F statistics are reported; values in square brackets are p-values.

Variable	Estimate
Earnings growth at $t-1$	0.2788
	(0.0524)
Male	0.0069
	(0.0012)
Age	0.0010
	(0.0005)
$\text{Age}^2/100$	-0.0014
	(0.0006)
Productions	-0.0228
	(0.0041)
Clericals	-0.0153
	(0.0039)
Region dummies	2.29
	[0.1010]
Year dummies	249.69
	[<0.0001]
Sector dummies	1.12
	[0.3320]
Number of observations	58,293

Table 6
The autocovariance structure of shocks to earnings

The table reports the estimates and the corresponding standard errors of the autocovariances of the unexplained component of real earnings growth, i.e., estimates of $E(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau})$. Data are pooled over all years. Panel B tests the null hypothesis that all the autocovariances $E(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau})$ are jointly insignificant for different values of τ.

Panel A

Order	All years	Order	All years
0	0.0143	5	-0.0001
	(0.0003)		(0.0001)
1	-0.0053	6	0.0001
	(0.0002)		(0.0001)
2	0.0001	7	-0.0001
	(0.0001)		(0.0001)
3	0.0002	8	0.0002
	(0.0001)		(0.0002)
4	-0.0002	9	0.0004
	(0.0001)		(0.0002)

Panel B

$E(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau})=0$	$\forall\tau\geq 1$	$\forall\tau\geq 2$	$\forall\tau\geq 3$	$\forall\tau\geq 4$
χ^2	665.34	46.18	34.46	20.79
[p-value]	[<0.0001]	[0.1192]	[0.1860]	[0.4717]
(d.o.f.)	(45)	(36)	(28)	(21)

Table 7
Firms' and workers' characteristics in the matched sample

Panel A reports summary statistics for the matched …rms in our data set; panel B shows descriptive statistics for the sample of matched workers. All statistics refer to 1991.

Panel A: Firm characteristics

	Mean	Median	Stand. dev.
Value added (million euro)	17.09	3.62	158.89
Employees	375	107	2,034
South	0.0995	0	0.2994
Center	0.1630	0	0.3694
North	0.7376	1	0.4400
$\frac{\text{Number of …rms defaulting}}{\text{Number of …rms borrowing}}$ (province level)	0.0231	0.0223	0.0098
$\frac{\text{Value of defaulted loans}}{\text{Value of outstanding loans}}$ (province level)	0.0298	0.0188	0.0273
Agriculture and Fishery	0.0016	0	0.0402
Mining	0.0057	0	0.0750
Food and tobacco products	0.0550	0	0.2280
Textiles and leather products	0.1302	0	0.3366
Paper, wood products and publishing	0.0902	0	0.2865
Chemicals and petroleum	0.1589	0	0.3657
Primary and fabricated metal products	0.1132	0	0.3169
Machinery and electrical/electronic	0.2240	0	0.4170
Energy, gas and water	0.0049	0	0.0695
Construction	0.0457	0	0.2089
Retail and wholesale trade, hotels	0.1132	0	0.3169
Transport and telecommunications	0.0230	0	0.1501
Credit, insurance and business services	0.0162	0	0.1262
Other private services	0.0182	0	0.1337

Panel B: Workers' characteristics

	Mean	Median	Stand. dev.
Earnings (euro)	13,929	12,502	6,189
Age	42.33	43	9.01
Male	0.7657	1	0.4236
Productions	0.6362	1	0.4812
Clericals	0.3452	0	0.4755
Managers	0.0185	0	0.1352
South	0.1113	0	0.3146
Center	0.1845	0	0.3879
North	0.7042	1	0.4565

Table 8
The sensitivity of earnings to value added shocks

The first row in Panel A reports the IV estimate of the sensitivity of wages to value added shocks (α for permanent shocks and β for transitory shocks). See Section 7.1 for more details on the instruments used in each regression. J-test is the test of overidentifying restrictions. F-test is the test of joint insignificance of excluded instruments. The Exogeneity test tests the null that in an OLS regression of $\Delta\omega_{ijt}$ on $\Delta\varepsilon_{jt}$, the latter is exogenous. $E(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau})$ is an estimate of the autocovariance of wage shocks of order τ; $E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})$ an estimate of the autocovariance of value added shocks of order τ; $E(\Delta\omega_{ijt}\Delta\varepsilon_{jt-\tau})$ an estimate of the cross-covariance of wage and value added shocks of order τ. σ^2_u, σ^2_v, σ^2_ξ and σ^2_ε are EWMD estimates of the variances of value added permanent shocks, value added transitory shocks, wage permanent shocks and wage transitory shocks, respectively. ϱ is an estimate of the MA coefficient of earnings. Asymptotic standard errors are reported in parenthesis. The Ratio is calculated as: $\sqrt{\frac{\alpha\sigma_u}{E[(\Delta\omega_{ijt})^2]}}$ and measures the amount of earnings variability attributable to value added shocks.

Panel A

	Permanent shock	Transitory shock
Sensitivity	0.0821	0.0037
	(0.0128)	(0.0049)
Number of obs.	24,956	34,931
J-test (p-value)	0.3463	0.4826
F-test (p-value)	<0.0001	<0.0001
Exogeneity test	20.03	
[p − value]	[<0.0001]	

Panel B

Moment	Estimate	Parameter	Estimate
$E(\Delta\omega_{ijt}\Delta\omega_{ijt})$	0.0136	σ^2_u	0.0247
	(0.0004)		(0.0032)
$E(\Delta\omega_{ijt}\Delta\omega_{ijt-1})$	−0.0052	σ^2_v	0.0326
	(0.0003)		(0.0043)
$E(\Delta\omega_{ijt}\Delta\varepsilon_{jt})$	0.0019		
	(0.0002)		
$E(\Delta\omega_{ijt}\Delta\varepsilon_{jt-1})$	−0.0001	σ^2_ξ	0.0065
	(0.0001)		(0.0026)
$E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt})$	0.0826	σ^2_ε	0.0029
	(0.0047)		(0.0018)
$E(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-1})$	−0.0302	ϱ	−0.1800
	(0.0036)		(0.1478)
		Ratio	0.1106

Table 9
The sensitivity of earnings to value added shocks: Accounting for sample selection

The table reports the IV estimate of the sensitivity of wages to value added shocks (Panel A is for permanent shocks and Panel B for transitory shocks). See Section 7.1 for more details on the instruments used in each regression. J-test is the test of overidentifying restrictions. F-test is the test of joint insigni…cance of excluded instruments.

Panel A: Sensitivity to permanent shocks

	Baseline sample (1)	Including those who change position (2)	Including those who leave the …rm (3)	Balanced sample (4)
Sensitivity	0:0821	0:0665	0:0659	0:0888
	(0:0128)	(0:0126)	(0:0120)	(0:0215)
Number of obs.	24,956	28,380	31,975	7,297
J-test (p-value)	0.3463	0.2878	0.2881	0.0703
F-test (p-value)	<0.0001	<0.0001	<0.0001	<0.0001

Panel B: Sensitivity to transitory shocks

	Baseline sample (1)	Including those who change position (2)	Including those who leave the …rm (3)	Balanced sample (4)
Sensitivity	0:0037	0:0036	0:0045	0:0016
	(0:0049)	(0:0048)	(0:0043)	(0:0109)
Number of obs.	34,931	39,961	45,675	9,390
J-test (p-value)	0.4826	0.4910	0.3128	0.0681
F-test (p-value)	<0.0001	<0.0001	<0.0001	<0.0001

Table 10
The sensitivity of earnings to value added shock:
Accounting for parameter heterogeneity

Asymptotic standard errors corrected for province clustering are reported in parenthesis; the partial R^2 for the reduced-form regression is reported in square brackets (see Shea, 1997). See Section 7.3 for more details on the instruments used in each regression. J-test is the test of overidentifying restrictions.

	Sensitivity to permanent shocks (1)	Sensitivity to transitory shocks (2)
¢"$_{j;t}$	0:1546	0:0107
	(0:0241)	(0:0100)
	[0:0679]	[0:2321]
¢"$_{j;t}$¤High risk aversion	¡ 0:0584	¡ 0:0109
	(0:0256)	(0:0108)
	[0:0637]	[0:2230]
¢"$_{j;t}$¤Manager	0:0552	¡ 0:0091
	(0:0950)	(0:0328)
	[0:0666]	[0:2986]
¢"$_{j;t}$¤s.d.[ln (V A$_{jt}$)]	¡ 0:0670	¡ 0:0034
	(0:0333)	(0:0063)
	[0:0412]	[0:3922]
¢"$_{jt}$¤Bankruptcy index	0:0484	0:0108
	(0:0234)	(0:0062)
	[0:0330]	[0:2331]
Number of obs.	24,956	34,931
J-test (p-value)	0.5062	0.0065