BANCA D'ITALIA

Temi di discussione

del Servizio Studi

Insurance within the Firm

by L. Guiso, L. Pistaferri and F. Schivardi

Number 414 - August 2001

The purpose of the Temi di discussione *series is to promote the circulation of working papers prepared within the Bank of Italy or presented in Bank seminars by outside economists with the aim of stimulating comments and suggestions.*

The views expressed in the articles are those of the authors and do not involve the responsibility of the Bank.

INSURANCE WITHIN THE FIRM

by Luigi Guiso*, Luigi Pistaferri** and Fabiano Schivardi***

Abstract

The full insurance hypothesis states that shocks to the firm's performance do not affect workers' compensation. In principal-agent models with moral hazard, firms trade off insurance and incentives to induce workers to supply the optimal level of effort. We use a long panel of matched employer-employee data to test the theoretical predictions of principal-agent models of wage determination in a general context where all types of workers, not only CEOs, are present. We allow for both transitory and permanent shocks to firm performance and find that firms are willing to fully absorb transitory fluctuations in productivity but insure workers only partially against permanent shocks. Risk-sharing considerations can account for about 10 percent of overall earnings variability, the remainder originating in idiosyncratic shocks. Finally, we show that the amount of insurance varies by type of worker and firm in ways that are consistent with principal-agent models but are hard to reconcile with competitive labor market models, with or without frictions.

JEL classification: C33, D21, J33, J41.

Keywords: insurance, incentive contracts, matched employer-employees data.

Contents

---

* Università di Sassari, Ente Luigi Einaudi and CEPR.

** Stanford University, SIEPR and CEPR.

*** Banca d'Italia, Research Department.

1. Introduction[1]

In the implicit contract literature firms act as insurers for their employees: risk-averse workers get insurance against income fluctuations in that wages are not linked to changes in productivity (see Rosen, 1985, for a survey).

Principal-agent models stress instead the role of incentives in determining the optimal compensation scheme, starting from the consideration that the interests of firm and workers rarely coincide. Workers enjoy leisure, and this may come at the expense of the firm's performance. If workers' actions in the workplace were perfectly observable it would be easy to devise ways of inducing them to produce the amount of effort agreed upon in a contract. But actions are only partially observable, or observability may require costly monitoring. A flat compensation scheme, while offering perfect insurance, removes any incentive for the worker to exert effort. One way to create proper incentives is to link compensation to the firm's performance. However, such risk-sharing has a cost, as it requires paying workers a premium that increases with their risk aversion. Thus, providing incentives curtails insurance.

The trade-off between incentives and insurance in the firm-worker relation has received a great deal of attention in the theoretical literature and is at the heart of modern contract theory. Much progress has been made in studying the design of incentive contracts under a variety of theoretical situations (see Gibbons, 1998, for a recent survey). Substantial progress has also been made in confronting some of the implications of the theory with the data, with two main strands: one attempting to measure whether incentives actually improve firm performance, the second aimed at testing the structure of the insurance-incentive model. Prendergast (1999) offers a thorough account of the empirical achievements.

Most of the empirical literature to date has been concerned with executives' compensation (see, for example, Jensen and Murphy, 1990; Margiotta and Miller, 2000).

[1] This paper is part of the EU-TMR research project "Specialisation versus diversification: the microeconomics of regional development and the spatial propagation of macroeconomic shocks in Europe". The Italian "Ministero dell'Università e della Ricerca Scientifica" and the "Taube Faculty Research Fund" at the Stanford Institute for Economic and Policy Research provided financial support. We thank Cris Huneeus, Marcel Jansen, Tullio Jappelli, Ed Vytlacil, Sevin Yeltekin, and seminar participants in Salerno, Turin, the Bank of Italy, the Society for Economi Dynamics 2001 meeting and the 2001 North-American conference of the Econometric Society for useful discussion and comments on a preliminary draft, and Antonio Alonzi, Antonietta Mundo and Salvatore Giovannuzzi of INPS for supplying us with the data on workers' compensation. The views expressed in this paper are our own and do not involve the Bank of Italy or the EU. We are responsible for all errors.

However, if one is interested in assessing the insurance role of the firm in the labor market, it is unlikely that CEOs provide a proper benchmark. First, CEOs and executives are in general a tiny fraction of the labor force. Second, they are a highly self-selected group of low risk averse (or perhaps even risk loving) workers who probably have only marginal interest in wage insurance. In contrast, little is known about the effects of incentive/insurance arrangements on lower-ranking workers' compensation. Yet incentives within the firm are not confined to managers. As is remarked by Lazear (1999), "there are many workers that a firm wants to be motivated". The use of bonuses and premiums related to the general conditions of the firm, possibly with different intensities for different types of workers, is common, as are more traditional payment mechanisms tied to individual output (such as piece rates).

In this paper we study the role of the firm as an insurance provider. We test the insurance-incentive trade-off in a context where various types of workers and firms interact, relying on matched firm-employee data available over a long period of time for Italy.

We start from the observation that if incentive/insurance considerations are important in shaping the relationship between workers and firms, then they offer testable implications for the observed compensation scheme. First, the model predicts that firm performance and workers' compensation move in parallel. Second, and more interestingly, it predicts that the amount of insurance varies with firms and workers characteristics in ways that are typical of agency models and are not shared by other schemes of workers compensation. The wealth of information of our data set allows us to measure most of these characteristics (e.g., workers' risk aversion or firm performance variability), and to verify if their effects on the compensation scheme are consistent with the predictions of the model. As we shall see, our results lend support to the agency model as a better representation of the compensation scheme than alternative models, such as the standard spot model with or without frictions and the implicit contract model.

Our study contributes to the empirical literature in several respects. First, since we base our analysis on a representative sample rather than a specific type of workers and firms, we can draw more general conclusions on the relative importance of insurance and incentives. Second, we can study how insurance coverage varies with types of firm and worker; this allows us to test some direct implications of the basic incentive/insurance model and help discriminate between the agency model and the competitive model with or without frictions. Third, since our data cover a span of years, we can study whether insurance provision is sensitive to the temporary or permanent nature of shocks. This issue has received little attention in the empirical as well as in the theoretical literature. However, it is plausible that the extent to which shocks are passed on to wages depends on their degree of persistence.[2] For instance, it may be that only transitory shocks to output are absorbed by the firm, while permanent shocks are shared, at least partially, with workers. Insofar as both types of shock are present, ignoring the distinction may bias the results towards insurance or incentives, depending on the relative importance of the transitory and permanent components. Fourth, our methodology allows for wage shocks that are unrelated to firm performance. Thus, we can compute how much of the observed earnings variability can be traced to workers sharing firms' risk and how much to idiosyncratic shocks to wages; while the former can potentially be insured within the firm, the latter are unlikely to be insurable. Finally, we propose a novel identification strategy that can also be applied to analogous problems arising in different areas of research.

The rest of the paper proceeds as follows. In Section 2 we review the insights of the standard principal-agent model. In Section 3 we characterize our empirical approach to the problem, considering a stochastic specification for firm performance and workers' earnings. We show that, in the spirit of the principal-agent model, if worker compensation is related to firm performance, 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 firms' performance are passed on to wages, and to what extent this is affected by whether the shock is transitory or permanent. Section 4 discusses how identification of the parameters of interest can be achieved. Section 5 describes the matched firm-worker data set used in the empirical analysis, and Section 6 presents the estimation of the stochastic model of firm performance and workers' earnings. The main empirical results are presented in Section 7, where we focus on the estimates of the incentive-insurance trade-off for the total sample and then examine the implications of the agency model for the sensitivity of wages to performance. Section 8 discusses the results and contrasts our findings with the implications of perfectly competitive models with and without frictions. Section 9 concludes and traces directions for further research.

---

[2] One paper we are aware of that addresses this point theoretically is Govindaraj and Ramakhrishnan (2000). They show that the sensitivity of the agent's payment to firm performance increases with the persistence of the latter (see their Proposition 2).

2. The standard principal-agent model

Labor market theories have different predictions concerning the relationship between the variability of earnings and that of firm output. In the competitive model with infinitely elastic labor supply curve, price-taking firms choose employment to equate the marginal product of labor to the market wage. According to implicit contract models, risk-neutral firms insure risk-averse workers against fluctuations in levels of activity. As a consequence, in both models the wage paid is orthogonal to firm-specific shocks.[3] In principal-agent models, by contrast, moral hazard considerations lead firms to link wages to performance. Under a series of assumptions, the sensitivity of the workers' pay to performance depends on many factors, including the noise in measured performance, the marginal cost of effort, the elasticity of performance to effort and risk aversion (Holmstrom and Milgrom, 1987). The differing implications that these models have for the compensation scheme allow us to design an empirical test to discriminate among them.

As noted, empirical studies so far have focused on the consistency of specific compensation contracts - such as those of CEOs - with the predictions of incentive models. Overall, the findings broadly support the predictions of the theory: for the groups under study, compensation responds to performance, implying that less than full insurance is offered (see Rosen, 1992, for a comprehensive survey of empirical studies of CEOs' compensation). Our goal is to assess whether incentive models of wage determination help predict the structure of compensation for an array of workers that goes beyond the limited groups that have attracted interest thus far. In addition we want to distinguish between transitory and permanent fluctuations in firm performance. As far as we know there has been no attempt to test the general applicability of the principal-agent theory of compensation and to relate it to the dynamics of the firm's performance.

To provide a framework for the subsequent empirical analysis, consider the following standard, one-period model. Firm performance is given by:

(1) $$y = z + f(e) + \varepsilon$$

---

[3] The two models differ in their predictions concerning lay-offs. For example, following a negative shock, in the competitive model some workers are fired, while in the implicit contract model they are temporarily laid-off but receive a payment that equates the utility of those working and those laid-off (Rosen, 1985).

where $\varepsilon \sim N(0, \sigma^2)$ is a random shock, z denotes a predictable component that depends on observable characteristics of the firm (such as location, industry, size, etc.), e is an unobservable (to the firm) component that depends on the worker's effort or action, and f represents the sensitivity of performance to effort.

Following Holmstrom and Milgrom (1987), assume that risk-averse workers receive a stochastic compensation that is the sum of a fixed component, a, that may vary with observable workers' characteristics (education, experience, etc.) and a variable component that depends on the performance of the firm, i.e. a bonus tied to output y:

$$w = a + by \tag{2}$$

Firms maximize profits $\pi = y - w$, while workers maximize utility $u(w - c(e))$, where $c(e)$ is the disutility of effort in wage units. Holmstrom and Milgrom show that if utility is of the CARA type, i.e. $u(x) = -\frac{1}{\rho}\exp\{-\rho x\}$, the optimal contract is linear and specifies:

$$b = \frac{f'(e)}{1 + \rho c''(e)\sigma^2} \tag{3}$$

where $\rho = -\frac{u''(.)}{u'(.)}$ is the coefficient of absolute risk aversion and $c''(.)$ the curvature of the agent's effort function, which is assumed to be convex; thus $c''(.) > 0$. Consider the special case in which $f'(e) = 1$. If workers are risk-neutral, $\rho = 0$ and they bear all risks. In this case $b = 1$. Risk aversion makes it worthwhile to reduce the impact of risk on wages ($b < 1$); under full insurance, $b = 0$. In the general case, the greater the marginal response of performance to effort $f'(e)$, the higher b. Note that if a firm employs workers with different preferences (ρ and $c(.)$) or different impact on performance (f), then we should expect different contracts to be offered accordingly. Finally, note that the sensitivity of workers' compensation to firm performance declines with output variability: $\frac{\partial b}{\partial \sigma^2} < 0$.

The predictions of this model are the basis of our empirical tests. The model suggests that the amount of insurance offered within the firm, parameterized by b, varies in predictable ways with workers' risk aversion, their propensity to shirk and the variance of the shocks to firm performance. Our aim is to recover an estimate of b and check whether the factors that theory indicates as relevant to determining the extent of insurance are important empirically. Note that in both the frictionless spot model with infinitely elastic labor supply curve and the implicit contract model of the labor market, $b = 0$: the compensation of workers who remain with the firm is unaffected by idiosyncratic shocks to the firm's performance. The distinction between the models is in employment dynamics; our study, however focuses only on the wage relationship.[4]

3. Shocks and insurance: modelling the stochastic structure of firms' performance and workers' earnings

The principal-agent model of the previous section offers a highly stylized characterization of wage contracts. When taken to the data, however, various adaptations are needed.

First, firm performance can be measured in several ways. The market value of the firm is perhaps the best measure, but it is only available for listed firms. In this respect, sales, profits and value added are more appealing proxies; here, we elect value added, but we examine the sensitivity of results using the alternative gauges. Second, the firm-worker relationship is dynamic. In dynamic agency models there may be an additional source of variability in output: the agent's ability or permanent component, assumed to be time-invariant (i.e., a random effect) in the simple principal agent model. This is utterly restrictive, as ability may evolve stochastically over time (say, due to learning on both sides of the employment relationship), which suggests a more appealing random walk specification. We incorporate this important feature in our empirical model. Third, in its bare form the agency model assumes that wage variability can only be explained by the combination of strength of incentives and output variability. With full insurance or no variability in performance, wages should evolve deterministically. However, there may be additional sources of unexplained wage variability that have nothing to do with the pay-performance relation.[5] We augment our wage specification to account for this. Fourth, in accordance with previous empirical

---

[4] The optimal static model has no place for terminations, and, accordingly, we examine the insurance role of the firm without considering them. In reality, the risk of losing one's job is a major source of uncertainty, implying that firings should play a role in an optimal incentive scheme. Theoretical results on this issue are still limited, given the analytical challenges that an optimal dynamic incentive model with dismissals presents. Using computational techniques, Sleet and Yeltekin (2000) show that, even in the presence of dismissals, wages will be linked to performance and thus used as an incentive device; they show that the basic predictions of the simple static principle-agent model carry over to this more general setup.

[5] For example, part of vector a in (2) could be unobservable to the econometrician even though perfectly observable to the firm (unobservable human capital components, measurement error, etc.). In this case one would record unexplained wage variability even in the absence of a contractual relationship between earnings and firm performance.

studies (Blanchflower, Oswald, and Sanfey, 1996; Aggarwal and Samwick, 1999), we use total compensation as the empirical counterpart of the wage in equation (2). Finally, departing from the theoretical model of the previous section, we consider a compensation scheme that is linear in logarithms, rather than levels. This has the advantage of fitting earnings and output data better. We now turn to the statistical characterization of the model.

3.1 *Firm performance*

Consider the case where the unanticipated component of (log) firm performance contains a permanent component ζ_{jt} which follows a random walk process, and a transitory component, $v_{jt,}$ which is serially uncorrelated. Permanent shocks 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. The firm can distinguish between transitory and permanent shocks but cannot determine whether they are due to pure chance or to workers' effort. In accordance with the basic model, the likelihood of both shocks may depend on workers' effort. For example, effort may influence performance permanently when the firm is trying to secure a new contract or develop a new product, and in a transient way when it is an input of a static production function. 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.

We model firm performance according to the following stochastic process:

$$A(L,p)y_{jt} = z'_{jt}\theta + \varepsilon_{jt} \tag{4}$$

where j and t are subscripts for the j-th firm at time t, $A(L,p)$ is a lag polynomial of order $p \geq 0$ (i.e. $A(L,p)x_{jt} = \sum_{\tau=0}^{p} \alpha_\tau x_{jt-\tau}$, with $\alpha_0 \equiv 1$), y_{jt} is a measure of observed firm performance, such as the logarithm of profits, value added or output, z_{jt} a vector of observable attributes, ε_{jt} the stochastic component of firm performance, and θ and A are parameters to be estimated. We assume that the stochastic component of firm performance has the following structure:

$$\varepsilon_{jt} = \zeta_{jt} + v_{jt} \tag{5}$$

(6) $$\zeta_{jt} = \zeta_{jt-1} + u_{jt}.$$

Equations (5) and (6) decompose the disturbance into a transitory component, v_{jt}, and a permanent one, ζ_{jt}. To simplify subsequent notation, 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. We assume that the two shocks v_{jt} and u_{jt} are serially uncorrelated and uncorrelated with each other. 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).

3.2 *Workers' earnings*

Consider now workers' compensation. The standard principal-agent model described above assumes that the only source of unanticipated fluctuations in wages is variability in the firm's performance. In reality, fluctuations in individual compensation depend also on individual idiosyncratic shocks (i.e., shocks that are unrelated to unanticipated changes in firm output, such as a spell of illness affecting productivity on the job). From a purely statistical point of view, another source of random variation in wages is measurement error.

We generalize equation (2) as:

(7) $$w_{ijt} = a_{ijt}'\delta + by_{jt} + \psi_{ijt}$$

where the subscript i stands for the i-th individual and w_{ijt} is the logarithm of worker compensation.[6] The term a_{ijt} denotes a vector of systematic factors that affect individual $i's$ compensation, which can vary across workers, firms and time, while ψ_{ijt} is the stochastic component of earnings, which is unrelated to the firm's fortunes. These idiosyncratic shocks are meant to capture unanticipated fluctuations in individual ability, shocks to productivity (such as illness), idiosyncratic changes in labor supply (child-raising, family labor supply effects, etc.).

---

[6] We let earnings to depend on contemporary 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 a year (say, a quarter), annual data of the type used here will not detect deviations from the contemporaneous adjustment assumption.

Let us assume that idiosyncratic earnings shocks are in turn the sum of a permanent random walk component $\eta_{ijt} = \eta_{ijt-1} + \xi_{ijt}$ and a serially uncorrelated transitory shock μ_{ijt}.[7] Again to save on notation, it is maintained that covariance stationarity holds, i.e. $E\left(\xi_{ijt}^2\right) = \sigma_\xi^2$, and $E\left(\mu_{ijt}^2\right) = \sigma_\mu^2$ for all t. The two shocks are serially uncorrelated and uncorrelated with each other at all leads and lags.

Taking first differences of (5) and (7), and using the stochastic structure outlined above, we get the following system of two equations:

$$A(L,p)\Delta y_{jt} = \Delta z_{jt}'\theta + u_{jt} + \Delta v_{jt} \tag{8}$$

$$\begin{aligned} A(L,p)\Delta w_{ijt} &= \Delta a_{ijt}'\lambda + \Delta z_{jt}'b\theta + bu_{jt} + b\Delta v_{jt} \\ &\quad + A(L,p)\xi_{ijt} + A(L,p)\Delta\mu_{ijt} \end{aligned} \tag{9}$$

where $\lambda = A\left(L,p\right)\delta$. 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.:

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

$$\Delta\omega_{ijt} = bu_{jt} + b\Delta v_{jt} + A(L,p)\xi_{ijt} + A(L,p)\Delta\mu_{ijt} \tag{11}$$

where, from equations (8) and (9), $\Delta\varepsilon_{jt} \equiv A(L,p)\Delta y_{jt} - \Delta z_{jt}'\theta$ and $\Delta\omega_{ijt} \equiv A(L,p)\Delta w_{ijt} - \Delta a_{ijt}'\lambda - \Delta z_{jt}'b\theta$.

Under our hypotheses, the serial correlation properties of $\Delta\varepsilon_{jt}$ are well defined: since it follows an MA(1) process (an assumption confirmed by the empirical analysis below), autocorrelations at the second or higher order are all zero. On the other hand, the serial correlation properties of $\Delta\omega_{ijt}$ depend on the order p of the lag polynomial $A(L,p)$. In general, $\Delta\omega_{ijt}$ will follow an $MA(p+1)$ process. The restrictions on the variance-covariance matrix of $\Delta\varepsilon_{jt}$ are standard and are reported below in the simple case of covariance stationarity:

[7] 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.

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

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.[8] 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.[9]

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 b coefficient is the same for the two shock components u_{jt} and Δv_{jt}. Yet we can test whether the amount of insurance varies with the temporary or permanent nature of the shock. Let b_u and b_v denote respectively the different response of wages to permanent and transitory shocks. We can distinguish various insurance regimes depending on the values of b_u and b_v. The contemporaneous covariance between shocks to performance and shocks to wage growth has the following structure:

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

where we have assumed that $E\left(\xi_{ijt}u_{js}\right)=E\left(\xi_{ijt}v_{js}\right)=0$ for all s,t, and similarly for μ_{ijt}, $E\left(\mu_{ijt}u_{js}\right)=E\left(\mu_{ijt}v_{js}\right)=0$ for all s,t. For simplicity, we have also assumed covariance stationarity. If workers are fully insured against fluctuations in the performance of the firm,

---

[8] In fact $-E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right)$ identifies the variance of the transitory shock σ_v^2 with $|\tau|=1$ for all t, while $E\left[\Delta\varepsilon_{jt}\left(\Delta\varepsilon_{jt+1}+\Delta\varepsilon_{jt}+\Delta\varepsilon_{jt-1}\right)\right]$ identifies the variance of the permanent shock σ_u^2 for all t. If there is covariance non-stationarity, then the expressions $-E\left(\Delta\varepsilon_{js}\Delta\varepsilon_{js+1}\right)$ and $E\left[\Delta\varepsilon_{js}\left(\Delta\varepsilon_{js+1}+\Delta\varepsilon_{js}+\Delta\varepsilon_{js-1}\right)\right]$ identify the variance of the transitory shock and the variance of the permanent shock at time s, respectively.

[9] If a classical measurement error $r_{it}\sim i.i.d.\left(0,\sigma_r^2\right)$ 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.

the contemporaneous covariance between shocks to performance improvement and shocks to wage growth is zero and full insurance obtains against 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 $b\left(\sigma_u^2 + 2\sigma_v^2\right)$, where $b = b_v = b_u$. We call this case "homogeneous partial insurance". This is likely to arise if firms are unable to distinguish between transitory and permanent shocks. Three other cases 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 $b_u\sigma_u^2 + 2b_v\sigma_v^2$. Two special cases occur when workers bear only transitory shocks but are insulated from permanent shocks ("permanent full insurance", characterized by $E\left(\Delta\varepsilon_{jt}\Delta\omega_{ijt}\right) = 2b_v\sigma_v^2$) or bear permanent shocks but are insured against transitory shocks ("transitory full insurance", and $E\left(\Delta\varepsilon_{jt}\Delta\omega_{ijt}\right) = b_u\sigma_u^2$).

Identification strategy Without further restrictions, from equation (13) we cannot separately identify b_u and b_v, nor can we gauge whether $b_u = b_v = b$. To see how identification of the relevant parameters is achieved, start from the general case $b_u \neq b_v$ in (11):

(14) $$\Delta\omega_{ijt} = b_u u_{jt} + b_v \Delta v_{jt} + \vartheta_{ijt}$$

where $\vartheta_{ijt} = A(L,p)\xi_{ijt} + A(L,p)\Delta\mu_{ijt}$. Subtract $b_v\Delta\varepsilon_{jt}$ from both sides to obtain:

(15) $$\Delta\omega_{ijt} - b_v\Delta\varepsilon_{jt} = (b_u - b_v)u_{jt} + \vartheta_{ijt}.$$

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

(16) $$E\left[\Delta\varepsilon_{jt+1}\left(\Delta\omega_{ijt} - b_v\Delta\varepsilon_{jt}\right)\right] = 0$$

(17) $$E\left[\Delta\varepsilon_{jt-1}\left(\Delta\omega_{ijt} - b_v\Delta\varepsilon_{jt}\right)\right] = 0.$$

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 b_v, 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}.[10] 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. At least in principle, all the variables $\Delta\varepsilon_{jt-\tau}$ (with $|\tau| \geq 1$) are uncorrelated with the error term. However, the instruments that satisfy $\Delta\varepsilon_{jt-\tau}$ with $|\tau| > 1$ are uncorrelated with the current unexplained component of value added growth, if this is an MA(1) process as in (10). Thus in estimation only $\Delta\varepsilon_{jt-1}$ and $\Delta\varepsilon_{jt+1}$ are used as instruments. Equations (16)-(17) can be used to identify the first parameter of interest b_v with one overidentification restriction. This can be tested with standard methods.

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

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

Equation (18) identifies the second parameter of interest b_u. 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 b_u, the extent of permanent insurance), what is left is uncorrelated with an MA(2) term centered in $\Delta\varepsilon_{jt}$ with unity coefficients.[11] Thus one can use $(\Delta\varepsilon_{jt+1} + \Delta\varepsilon_{jt} + \Delta\varepsilon_{jt-1})$ 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, $\sum_{k=-q}^{1}\Delta\varepsilon_{jt+k}$ (for any $q \geq 2$) is a valid instrument as well. It follows that the model can be tested *via* these additional overidentifying restrictions.

---

[10] It is worth noting that OLS estimation provides unbiased and consistent estimation if $b_u = b_v = b$. Thus an exogeneity test for $\Delta\varepsilon_{jt}$ can be used to check whether $b_u = b_v = b$.

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

$$\Delta\omega_{ijt} = b_u\Delta\varepsilon_{jt} + \left[\left(b_v - b_u\right)\Delta v_{jt} + \vartheta_{ijt}\right].$$

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

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. Also, note that if we had shocks to value added in levels (i.e., estimates of ε_{jt}), we could have many more instruments available to estimate b_u in (18). In fact, $\varepsilon_{jt+\tau}$ $(\tau > 1)$ will all be valid instruments, as can be easily checked. Finally, the moment conditions derived above are valid regardless of the covariance stationarity hypothesis, which provides a convenient level of generality.[12]

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 (if this is the only source of uncertainty in the model). The popular income process involving permanent random walk plus transitory serially independent component implies that the consumption innovation adjusts fully to permanent income shocks (u_{jt}), but only to the annuity value of transitory shocks (v_{jt}). 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.

The foregoing is a discussion of the identification of the two insurance parameters b_u and b_v. 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:

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

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

---

[12] 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 standard relation: $\Delta\varepsilon_{jt}^* = \Delta\varepsilon_{jt} + \Delta r_{jt}$) implies that the estimate of b_v is biased toward zero while that of b_u is unaffected. If the true b_v is zero, however, there is no bias. The problem is one of invalid instruments; to some extent, it is possible to check measurement error bias by checking whether overidentifying restrictions are rejected in our model.

and use minimum distance estimation similar to that suggested by 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. Under covariance stationarity $E\left(v_{jt}^2\right) = \sigma_v^2$ and $E\left(u_{jt}^2\right) = \sigma_u^2$ for all t.[13]

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 $p = 0$, with heterogeneous partial insurance and covariance stationarity, one obtains:

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

Conditioning on the estimated values for b_u, b_v, 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.

Before turning to the description of the data, it is worth remarking that the identification strategy outlined in this section is implemented in a series of steps. First, one needs to filter predictable components from both firm performance and workers' earnings. Since these are perfectly observable, incentive contracts will not be made contingent on their realizations. The observables include the autoregressive components (if any), and exogenous characteristics z_{jt} and a_{ijt}. When the empirical exercise is implemented using standard IV estimation techniques, the resulting residuals are consistent estimates of the unexplained growth rates $\Delta\varepsilon_{jt}$ and $\Delta\omega_{ijt}$.[14]

The next step is to use (16)-(18) to estimate b_u and b_v and check whether these are affected by observable firm and individual characteristics along the lines of what is predicted by agency models.

We then calculate the sample analogs of the theoretical autocovariances $E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right)$ and $E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau}\right)$. For more technical details see Appendix B; for a more thorough

---

[13] In the empirical analysis for brevity we report estimates obtained under the assumption of covariance stationarity. Those obtained under covariance non-stationarity are available on request.

[14] 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).

discussion of covariance estimation see Chamberlain (1984). Estimated autocovariances are then used as inputs for the minimum distance estimation of the variances of shocks to value added and earnings conditioning on insurance/incentive arrangements.

4. 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 for brevity). Since for each worker we can identify the firm, we combine the two data sets and use them in a matched employer-employee framework.[15] There is a burgeoning empirical literature on the use of matched employer-employee data sets (see Hamermesh, 2000, 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 approximately 30,000 firms 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 firms, 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. The INPS lacks information on self-employment and on public employment, which is also excluded from the CAD. As we describe in Appendix A, the INPS data set derives from forms filled out by the employer

[15] The INPS data set has been used by Casavola, Cipollone and Sestito (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.

that are roughly comparable to those collected by the Internal Revenue Service in the US.[16] Misreporting is prosecuted.

Given that the INPS data set includes a fiscal identifier 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. In particular, the data set reports total earnings and the number of weeks worked in each year.

Table 1 reports various descriptive statistics for the firms (Panel A) and workers (Panel B) present in our sample. Panel A shows the main characteristics for the sample of firms in the CAD data set. From an initial sample of 177,654 firm/year observations, we end up with a sample of 116,809, excluding firms with intermittent participation (40,225 observations) and those with missing values on the variables used in the empirical analysis (20,620 observations).[17]

The sample ranges from very small firms to firms with almost 180,000 employees, with an average of 204 and a median of 60. As expected, most of the firms are in the North (75 percent). As for the distribution by industry, firms in the chemical, metal production 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 1974-1994 INPS sample. We start with an initial sample of 383,985 worker/year observations and end up with 186,715. 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 (5,564

---

[16] 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.

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

observations), circumventing the problem of modelling human capital accumulation and retirement decisions. To avoid dealing with wage changes that are due to job termination (registration or layoffs) or unstable employment patterns, we excluded workers with part-time employment, those who change position during a year, and those with multiple jobs (127,403 observations). For similar reasons, we dropped individuals who worked for less than 12 months (63,347 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. Finally, we kept only individuals with non-zero recorded earnings in all years (148 observations lost) and eliminated some outliers (808 observations)[18].

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.[19] For workers with intermittent participation we 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 25 million lire on average (at 1991 prices and exchange rates, roughly \$17,300), with a median of 23 million (\$15,600).[20]

5. Estimation of the stochastic structure of firm performance and workers' earnings

5.1 *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). Value added is the closest measure to the theoretical concept of firm performance y in the model described

---

18 An observation is classified as an outlier if (a) real earnings are below 1 million lire (500 Euro), (2) real earnings are below 6 million *and* the growth rate is below -200 percent, or (c) real earnings exceed 100 million lire *and* the growth rate is greater than 200 percent.

19 Results are very similar if we use gross income as a measure of earnings.

20 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.

in Section 2. In our view, it constitutes a better gauge of y than gross output or sales, since it 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).[21]

To identify shocks to firm performance we proceed along the lines of Section 2. The first step is to filter out the predictable component. To this end we consider equation (4):

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

where y_{jt} is the log of real value added for firm j at time t. In the presence of a random walk permanent component and/or serially correlated transitory effects (for instance, an $MA(q)$ process), estimates in levels are inconsistent because of a short T problem. Taking the first difference of the data eliminates this problem and yields:

$$A(L,p)\Delta y_{it} = \Delta z'_{jt}\theta + \Delta\varepsilon_{jt}. \tag{22}$$

Also notice that the first difference eliminates firm-fixed effects (if any) that may influence the level of the firm's output, such as size. Included in $\Delta z'_{jt}$ is a full set of time dummies, sector dummies, location dummies, year/sector and year/location interactions.[22] The orthogonality condition that identifies the parameters of (22) in the general case is:

$$E\left(\Delta\varepsilon_{jt}\left|\Omega_{t-q-2}\right.\right) = 0 \tag{23}$$

[21] When alternative measures of performance (such as gross output or sales) are used, the results are very similar.

[22] 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 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. In terms of the agency model these components should not enter the optimal contract, because they can be separated from effort, given that they are common to all firms in a given location or sector at each point in time.

where Ω_{t-q-2} denotes the information set on which the agents condition. The residual from (22) 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.[23]

We use the IV estimator suggested by Anderson and Hsiao (1982), which provides consistent estimates of the parameters of interest in dynamic panel data models.[24] The results of the IV regression are reported in Table 2. We assume that $p = 1$ and $q = 0$, and check whether such restrictions are consistent in the data (see below). A look at the reduced form shows that these instruments are powerful for identification (a p-value on the excluded instruments below 0.1 percent). We find an AR parameter of 0.28 with a standard error of 0.02. Time, area and sector effects are not statistically significant; interactions between time and area and between time and sector, however, are jointly 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 $(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau})$, reported in Table 3 pooling over all years, reveals the absence of any large or statistically significant correlation at lags greater than one, consistent with $\Delta\varepsilon_{jt}$ being an MA(1) process. This can be tested more formally using the zero restriction test proposed by Abowd and Card (1989). We find that the null that $\Delta\varepsilon_{jt}$ is an MA(0) process is overwhelmingly rejected (p-value $<$0.0001), while the null of MA(1) has a borderline p-value of 4.5 percent. The p-value of the test increases slightly with the order of the MA process being tested. A difference test MA(0) vs. MA(1) rejects the null (p-value$<$0.0001), while a difference test MA(1) vs. MA(2) supports the null (p-value 28 percent) and the proposition above that $\Delta\varepsilon_{jt} \sim MA(1)$. This makes us confident that in the estimation of b_u and b_v below, one need not be concerned about instrument validity, just about power. Thus one autoregressive lag is sufficient to characterize the predictable dynamics in the growth rate of firm value added (alongside the indicators for a given time/sector/location configuration); there is thus evidence for value added growth being an ARMA(1,1) process.

---

23 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.

24 More efficient estimates can be obtained using the two-step procedure suggested by Arellano and Bond (1991). This, however, may have severe finite sample bias, as noted by Blundell and Bond (1999) among others. We thus resort to a simple one-step estimator.

Overall, these results suggest that the random walk plus serially uncorrelated transitory shock specification is a reasonable representation of the stochastic component of value added data.[25]

5.2 *Workers' earnings*

For workers' earnings we consider a logarithmic specification of the process (7), in which the first difference of log annual net real earnings is regressed on a set of observable attributes: a fourth-order polynomial in age, education (here proxied by a set of occupation dummies), gender, area of residence dummies, and time dummies (the vector a_{ijt}). 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. Our estimated regression is:

$$A(L,p)\Delta \ln w_{ijt} = \Delta a'_{ijt}\lambda + \Delta\omega_{ijt}. \tag{24}$$

Recall from Section 3 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. Moreover, under the same hypothesis, $\Delta\omega_{ijt}$ is no longer an $MA(1)$ process but an $MA(2)$ process if $p = 1$, as we have assessed in the previous section. We thus impose $p = 1$ and estimate the specification (24) by IV using the orthogonality condition:

$$E\left(\Delta\omega_{ijt}\left|\Omega_{t-3}\right.\right) = 0. \tag{25}$$

The results from estimating equation (24) are reported in Table 4. The AR(1) coefficient takes on a value of 0.33, with a standard error of 0.03. The reduced form regression (not reported here) shows that the instruments have sufficient predictive power: the *p*-value of the *F* test is in fact below 0.1 percent.

---

[25] 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 would not be an extremely unlikely characterization of the data. The full matrix of estimated autocovariances is available on request.

As in the case of firms, we use the residual of the IV 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 5.

A thorough examination of the estimated autocovariances of the unanticipated component of the rate of growth of earnings reveals that there is no large or statistically significant covariance at lags greater than one. This evidence is not entirely consistent with $\Delta\omega_{ijt}$ being an MA(2) process of the form:

$$\Delta\omega_{jt} = b_u u_{jt} + b_v \Delta v_{jt} + A(L,p)\xi_{ijt} + A(L,p)\Delta\mu_{ijt} \tag{26}$$

as in the modified version of equation (11) with $p = 1$.[26] On average, the autocovariance of order zero is 0.016, while the autocovariance of order one is -0.006.[27] Autocovariances of order higher than two are economically very small (between -0.0005 and 0.0004) and mostly insignificant.

Armed with these results we can now recover the implied values of parameters b_v and b_u following the procedure described in Section 3.2 and focusing on the matched employer-employee data set.

6. Shocks and insurance: the estimates

The matched data set includes 39,930 individual/year observations for 8,228 workers and 4,194 firms. It is an unbalanced matched panel of firms and workers. The mean number of matches (i.e., the number of workers) per firm is 1.96, with a minimum of 1 and a maximum of 397 per year. Table 6 reports characteristics for the firms and the workers in the 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 103 in the matched sample, compared with 60 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. Other characteristics (such

---

[26] There are two possible explanations for this. First, $\sigma^2_\mu = 0$ (absence of a transitory component once the AR component is removed from 24). Second, a low value of the AR coefficient may make an MA(2) hard to detect in the data.

[27] These are much lower than the estimates for the US using the PSID (Meghir and Pistaferri, 2000), perhaps reflecting the fact that measurement error is less of a problem in this data set.

as location, industry, and workers' demographic characteristics), are fairly similar in the two samples.

We apply the identification strategy outlined in Section 4 to recover the parameters of interest, b_u (the sensitivity of earnings shocks to permanent shocks to value added) and b_v (the sensitivity to transitory shocks). In both cases, our estimating equation is:

$$\Delta\omega_{ijt} = \beta\Delta\varepsilon_{jt} + \vartheta_{ijt}. \tag{27}$$

As explained in Section 4, parameter b_v is identified using $\Delta\varepsilon_{jt-1}$ and $\Delta\varepsilon_{jt+1}$ as instruments, while parameter b_u is identified using $\sum_{k=-q}^{1}\Delta\varepsilon_{jt+k}$ (with $q = 1, 2, 3$) 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 instruments excluded. Finally, an exogeneity test for $\Delta\varepsilon_{jt}$ (Davidson and MacKinnon, 1993) is an implicit test for $b_u = b_v = b$.

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 $\frac{\sqrt{b_u^2\sigma_u^2+2b_v^2\sigma_v^2}}{\sqrt{E\left[(\Delta\omega_{ijt})^2\right]}}$, 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.

6.1 *Main results*

Table 7 shows the results of our exercise. The first thing to emerge from the table is that workers' wages do reflect shocks to the firm's value added: $E(\omega_{ijt}\varepsilon_{jt}) = 0.0019$ 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 b_v (which measures the sensitivity of workers' earnings to transitory shocks) is economically small (point estimate 0.007) and not statistically different from zero (standard error 0.0051). The estimated value of b_u (responsiveness to permanent shocks) is 0.06, an order of magnitude greater, with a small standard error of 0.0154.[28] Joint consideration of the point estimates and of the standard errors of the two parameters suggests that $b_u \neq b_v$.[29] 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 (and that measurement error bias in the estimation of b_v is negligible, see note 12). Instruments' power is not a concern, as is shown by the low p-value of the F-test in the reduced form regressions.

Our findings imply that a 10 percent permanent change in firm performance induces a 0.6 percent permanent variation in earnings for those employed at the same firm on a continuing basis.[30] To get a sense of the economic significance of this effect consider the median firm (value added of 3.49 million euro, 103 employees, paying an average salary of 12,409 euro per year). Evaluated at the sample median, a permanent decrease in value added of 349,500 euro (10 percent) - equivalent to a 3,413 euro drop in value added per (initial) worker - would permanently lower the earnings of the continuing workers by 203 euro.

Table 7, Panel B, also reports the estimated value of the relevant moments of the shocks to output and wages. One can notice that while these are lower than for the full sample (see Tables 3 and 5), they are not dramatically different.[31] Consistently with the estimates of b_u

---

[28] 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: b_u is estimated with a slightly higher standard error of 0.0252, but the associated p-value is still only 1.7 percent.

[29] 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 $b_u = b_v = b$.

[30] Using gross rather than net earnings produces a slightly higher coefficient of 0.0692, implying less insurance than when net earnings are used. This is consistent with the view that tax progressivity provides implicit insurance.

[31] For example, $E\left(\Delta\omega_{ijt}\Delta\omega_{ijt}\right)$ is 0.0139 in the matched sample and 0.0165 in the full sample; $E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-1}\right)$ is -0.0055 in the matched sample and -0.0065 in the full sample.

and b_v, as seen before we find that the estimate of $E\left(\Delta\omega_{ijt}\Delta\varepsilon_{jt}\right)$ is positive and statistically significant, while that of $E\left(\Delta\omega_{ijt}\Delta\varepsilon_{jt-1}\right)$ is economically minuscule and 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 b_u and b_v) the variances of idiosyncratic shocks to earnings.[32] The estimate of the variance of the permanent shock to value added, σ_u^2, is 0.0229 (with a standard error of 0.0035), while the estimated variance of the transitory shock, σ_v^2, is 0.0334 (with a standard error of 0.0048). These are both sizeable and imply standard deviations of 15 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:

$$\widetilde{\Delta\omega_{jt}} \equiv \Delta\omega_{ijt} - b_u u_{jt} - b_v \Delta v_{jt} = \xi_{ijt} + \varrho\xi_{ijt-1} + \Delta\mu_{ijt} + \varrho\Delta\mu_{ijt-1} \tag{28}$$

i.e., $\widetilde{\Delta\omega_{jt}}$ follows a composite MA(2) process. In part, the coefficient ϱ will reflect the legacy of the autoregressive process of the value added (see above); 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 0.0058 (standard error 0.0015), while σ_μ^2 is 0.0034 (s.e. 0.0011). The MA coefficient ϱ in the stochastic process of earnings is negative (-0.16) and marginally significant.

The variability in compensation induced by the incentive scheme adopted depends on $b_u\sigma_u$ (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 0.1179, one can infer that roughly 8 percent of the total earnings variability can be explained by firm-specific risk (see the last row of Table 7), while the remaining component is related to workers specific shocks.

---

[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.

6.2 *Insurance and firm-worker characteristics*

As we saw in Section 2 the principal-agent model of wage determination implies that the response of wages to performance varies in predictable ways with workers' risk aversion, the curvature of the effort function, the variance of shocks to firm performance and the elasticity of performance with respect to effort. As Aggarwal and Samwick (1999), among others, argue forcefully, these implications are the most useful for the empirical assessment of the model, particularly to discriminate it from alternative models that predict a correlation between firm performance and worker compensation. We now address this issue directly, using the wealth of matched employer-employee data at our disposal. In particular, we go beyond the prediction tested by Aggarwal and Samwick that the sensitivity of compensation to performance should decline with the variance of performance and consider also the other implications of the model. In addition, we also consider further predictions coming from generalizations of the basic principal-agent model that can be tested with our data set.

For workers characteristics, we can exploit outside information on risk aversion derived from experiments to construct a measure of risk aversion and study its relation with the level of wage insurance. The degree of insurance might also depend on the occupational status. In particular, the optimal contract for CEOs and executives should prescribe more incentives in the light of the greater responsiveness of firm performance to their effort. Another characteristic that may help explain differences in the amount of insurance received is tenure. Gibbons and Murphy (1992) show that when there are career considerations, i.e. concerns about the effect of current performance on future compensation, the explicit incentives from the optimal compensation contract should be strongest for workers close to retirement, whose career concerns are weakest.

Firm characteristics are important as well. The first, direct implication of the model is that firms with more noisy performance should rely less on incentive schemes and provide more insurance, a proposition that can be tested by considering the firm-level variance of performance. A second set of implications relates to size, which is likely to matter for several reasons. First, larger firms have easier access to financial markets, allowing them to buffer shocks, which is likely to make them more willing to provide insurance (i.e., to bear risk). In particular, larger firms have easier access to equity capital and can thus transfer risk to the market. Second, insofar as larger firms produce multiple goods in multiple locations with diversified processes (or belong to a conglomerate), they might be more willing to assume risk. Third, the strength of incentives itself may depend on size. If output depends on aggregate workers' effort and it is impossible to disentangle individual contributions, the incentive mechanism becomes less effective as the number of workers increases.[33] These considerations imply more extensive insurance provisions in larger firms (i.e., lower b_u and b_v). However, smaller firms should be characterized by a larger variance of shocks to output which, according to equation (3), should imply less reliance on incentive schemes.[34] Moreover, (surviving) small firms tend to show a higher growth rate, which allows for a more extensive use of career promises with respect to monetary rewards to motivate workers. In addition, only large firms might be able to implement sophisticated contracts.[35] The effect of size may also pick up the fact that small firms can use the threat of dismissal as a more effective discipline device. This is quite likely, as in Italy firing and hiring rules are much stricter for large firms. Finally, using a variant of the standard agency model, Schaefer (1998) shows that the sensitivity of wages to firm performance will be increasing with firm size if the marginal productivity of effort increases with size more rapidly than the amount of risk faced by workers. Larger firms will therefore find it optimal to forgo risk-sharing in favor of more powerful incentives. This discussion indicates that size might be important for insurance provision, but that its effect can only be signed empirically.

The last prediction we consider is based on an extension of the basic model that emphasizes the importance of comparative performance to optimize the wage contract (see, e.g., Holmstrom and Milgrom, 1987). If shocks to performance are correlated across firms, then the uncertainty about the unobservable component of output growth can be reduced by looking at the performance of similar firms. An interesting feature of the Italian economy is the widespread presence of industrial districts. These are groupings of small- and medium-sized firms specializing in a particular product and located in a circumscribed area, ordinarily within a few miles one another. Firms within a district should be better able to extract valuable

---

[33] This can be mitigated by the presence of strong peer pressure. As Prendergast (1999) points out, however, there is little empirical evidence that peer pressure circumvents free riding in large production units.

[34] Jovanovic (1982) presents a selection model that delivers a positive association between size and the variance of the rate of growth in output. Such predictions find widespread empirical support (e.g., Dunne, Roberts and Samuelson, 1989).

[35] For example, FIAT (the largest private employer in Italy) initiated profit-sharing in the mid-1980s in order to motivate workers and implement Japanese-style production strategies.

information from the performance of closely related firms compared to those outside a district, and therefore rely more on incentives.

To check all the above implications we modify our IV estimation strategy allowing the sensitivity coefficients b_u and b_v to depend on observable worker and firm characteristics. Thus we estimate by IV:

$$\Delta\omega_{ijt} = \beta\left(X_{ijt}\right)\Delta\varepsilon_{jt} + \psi_{ijt} \tag{29}$$

with $\beta\left(X_{ijt}\right)$ being linear in the set of individual-firm variables X_{ijt} (risk aversion, job type, tenure, firm size, location in an industrial district, historical variability of firm performance, and a constant term). This amounts to including interactions of such variables with value added growth shocks, $\Delta\varepsilon_{jt}$, and augmenting the set of instruments by the interactions of the original instruments with the relevant worker and firm characteristics.[36]

As far as worker characteristics are concerned, we create dummies for production workers, clerical workers and managers, and use age as an (imperfect) indicator of seniority. Risk aversion, however, is not available in the CAD data set. 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.[37] Following Guiso and Paiella (2001), we use the answers to obtain a measure of the Arrow-Pratt index of relative risk

---

[36] Our estimates of $b\left(X_{ijt}\right)$ are not affected by the relationship between $\Delta\omega_{ijt}$ and X_{ijt} that may happen to exist in the cross-section. Including main effects has virtually no effect on the estimates of $b\left(X_{ijt}\right)$ (results available on request).

[37] 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 Euro 5,000 (or $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.

aversion for each consumer.[38] 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 about 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.[39] 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.[40]

As for firm characteristics, the historical performance 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 also construct a dummy for location in an industrial district and use a quadratic in log firm size (number of employees). Using a polynomial in log firm size has the advantage that when evaluated at the minimum size observed in our sample (one employee) the wage-performance sensitivity is zero, which can be thought of as the baseline. Table 8 reports the results. Column (1) shows the effects of worker and firm characteristics on

---

[38] Let Z_i be the maximum amount consumer i is willing to pay to enter the lottery; c_i the endowment and u_i the utility function. The maximum price for participating in the lottery is then defined by:

$$Eu_i\left(c_i\right) = \frac{1}{2}u_i\left(c_i + x\right) + \frac{1}{2}u_i\left(c_i - Z_i\right)$$

where E is the expectations operator and x what the agent gains in the favorable state (i..e, $x = 10$ million lire). Taking a second order Taylor expansion and solving for the Arrow-Pratt measure of absolute risk aversion $A_i\left(c_i\right)$ gives

$$A_i\left(c_i\right) = -\frac{u_i''\left(c_i\right)}{u_i'\left(c_i\right)} = \frac{2\left(10 - Z_i\right)}{\left(100 + Z_i^2\right)}$$

Given that Z_i is known, this expression can be recovered for all those who answer the survey question on the lottery. Relative risk aversion $R_i\left(c_i\right)$ is obtained by multiplying A_i by individual i's consumption c_i.

[39] 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.

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

the sensitivity of wages to permanent shocks to performance; column (2), 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.

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 b_u). Overall, there is a quite sizable sensitivity differential due to risk aversion (-0.07). In the same direction, managers have less insurance than either white collar or blue collar workers (the excluded category). However, 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). Finally, we find no solid evidence of a relation between tenure (age) and incentive schemes.

In terms of firms, consistently with the predictions of the basic agency model, those with higher variability in performance provide more insurance and less incentives: the coefficient is negative (-0.0338) and highly 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 theory. Predictions coming from extensions of the basic model also find empirical support. Firms located within a district provide less insurance and more incentives than others, and the difference is statistically significant. This is consistent with the idea that district firms can rely on their neighbors to improve the precision of their inference and according to the agency model this should make it easier to motivate workers and thus result in less insurance, other things being equal. Size also matters. The pattern of incentives is increasing and concave with the size of the firm, a finding that is not novel to our study (see Gibbons and Murphy, 1997, for similar evidence).[41] This goes against the idea that incentive mechanisms are less effective in large production units (perhaps because of free riding effects), and that insurance is more costly to small firms.

To get a sense of the results contained in Table 8, consider a 30 years-old, highly risk-averse production worker employed in a medium-sized firm (50 employees) located in

---

[41] Note that the pay-performance elasticity with respect to permanent shocks ($\frac{\partial \Delta \omega_{ijt}}{\partial u_{jt}}$) equals $\frac{b_u (\ln S)}{S} + 2b_u \left[(\ln S)^2\right] \frac{\ln S}{S}$. Given the empirical estimates of $b_u (\ln S)$ and $b_u \left[(\ln S)^2\right]$, the implied sensitivity increases with firm size at a decreasing rate.

a district, and with a historical performance variability of 20 percent. For this worker, the sensitivity to firm permanent shocks is 0.06. For an individual with the same characteristics $-$except risk aversion$-$ the coefficient becomes as high as 0.14. For a firm with the same characteristics $-$but a larger standard deviation of, say, 50 percent$-$ the coefficient declines to 0.05. In line with the predictions of the agency model, changes in worker and firm characteristics may thus impart a wide range of variability in b_u.

Note finally that the p-value of the J-test does not point to misspecification of the model (11 percent), and that in all cases the power of the instruments (as measured by the partial R^2) is high enough to allow identification of the relevant parameters and to dismiss the possibility of finite sample bias and inconsistency.

In column (2) we repeat the estimation exercise for the sensitivity of earnings to transitory shocks. In accordance with the results reported in Table 7, neither worker nor firm characteristics appear to be statistically significant. This implies that insurance of transitory shocks to value added is pervasive involving all types of workers and all types of firms. The J-test has a high p-value of 36 percent, which suggests that it is unlikely that the model is misspecified due to measurement error. The high values of the partial R^2 for the reduced-form regressions, on the other hand, suggest that the lack of a relation between transitory shocks and wages cannot be explained by low power of the instruments.

7. Discussion

Our evidence points to three important considerations. First, workers do share the firm's fortunes, at least in part. The unexplained earnings variability that can be attributed to incentive schemes is nearly 10 percent. Second, while transitory shocks to value added do not affect wages, permanent shocks are partly transferred to earnings. Finally, our search for heterogeneity in the sensitivity of earnings to shocks shows that insurance increases with workers' risk aversion, declines with the amount of responsibility within the firm (as measured by job position), is less for firms located in an industrial district and greater for those with high output variability, and is a decreasing and convex function of firm size. Most of these results are consistent with the predictions of the agency model but very hard to reconcile with competitive models of wage determination even when some friction is allowed for.

Consider a competitive model of the labor market, in which price-taking firms choose employment to equate the marginal product of labor to the market wage. Wages should not respond at all to idiosyncratic shocks to the firm, which faces an infinitely elastic labor supply at the prevailing wage. This is the version tested by Blanchflower, Oswald and Sanfey (1996), although they correlate wages with *industry* rather than firm profitability. Our test of the perfect competition model is less stringent, in that we do let wages respond to industry shocks - as any competitive model of the labor market with limited mobility across industries would imply. All that we require for the perfectly competitive model to hold is that wages not respond to firm-specific shocks, as is implied by competition *within* the industry. Obviously, this is a much weaker requirement than the null hypothesis that workers wages do not respond to shocks to the industry. Rejecting perfect competition using industry data may not be surprising in the presence of, say, some market segmentation or industry-specific skills. Thus our results can be seen as a more robust rejection of the perfect competition model.

As Blanchflower, Oswald and Sanfey (1996) note, however, the short-run labor supply might be upward sloping, due to temporary friction. For instance, in the presence of workers' mobility costs of the type considered by Bertola (1999), firms face an upward-sloping labor supply curve. In these models, the responsiveness of wages to idiosyncratic shocks to firm performance is determined by the slope of the labor supply curve: the further away from infinite elasticity, the stronger the implied correlation between performance and earnings. Thus, one could argue that while the correlation that we find between wages and firm-specific shocks is inconsistent with the frictionless, perfectly competitive model, it could be made consistent with an extended version that allows for some temporary friction. Yet our evidence is difficult to reconcile with these models as well. Competitive models with either firing or mobility costs carry direct implications for the response of wages to shocks of different duration. Specifically, since adjusting employment rather than wages is relatively more advantageous vis-à-vis permanent shocks, wages should respond *more* to temporary than to permanent shocks, a prediction that is strongly at variance with our empirical findings.

Apart from this direct evidence, the competitive model fails to fit with some other features of our results. First, whereas in principal-agent models the response of wages to performance depends in predictable ways on well identified firm and worker characteristics, competitive models (with or without friction) have no clear implication of how the correlation between wages and firm performance varies with those attributes. Second, one could argue that an environment corresponding more closely to the competitive paradigm is that of industrial district firms. In particular, firms are highly similar and employ workers with similar characteristics; workers' mobility and search costs are negligible (firms being located very close to one another); firms are small, implying price-taking behavior. Thus, if the competitive model were valid one should expect lesser sensitivity of wages to performance among district firms, since they are likely to have a more elastic short run labor supply. Again, however, this is the opposite of what we actually find.

Overall, we take our results on wage determination to be remarkably consistent with the predictions of the agency model and inconsistent with competitive theories, not only in the extreme version characterized by continuously clearing markets but also in more realistic versions allowing for rigidities in the form of firing or mobility costs.

8. Conclusions

We began by observing that in the principal-agent model there is a trade-off between earnings incentives and wage insurance against the vagaries of the product market. We offer empirical evidence on the extent of insurance and incentive provision within the firm, based on a matched employer-employee data set that spans almost 15 years, from the early 1980s to the mid-1990s, in Italy.

The main empirical finding is that the provision of incentives is not limited to managers and executives, the two groups on which the empirical literature has mainly focused. We document the existence of incentives for all categories of employees, including production and clerical workers. However, while full insurance is provided against temporary shocks, lasting disturbances to output are only partially insured.

The sensitivity of workers' wages to shocks to the firm varies systematically with firm and worker attributes. In particular, it is a concave function of firm size, it is greater for firms located within an industrial district, where output is less noisy and information about workers' effort is easier to obtain and process, and it is lesser for firms with high overall performance variability. Our estimates also suggest that the responsiveness to shocks depends on the employee's position within the firm: managers' compensation is more reactive than that of white-collar or blue-collar workers, although estimates are not at all precise. In addition, workers risk aversion is negatively correlated with such sensitivity, consistent with the agency model of wage determination.

These findings are sufficiently robust for us to draw a few conclusions. First, all workers share at least partly the fortunes of their company, to an extent that depends on their relationship with the firm (i.e., position and tenure), and - more importantly from an economic and statistical point of view - on 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. Let us remark that the distinction made here between transitory and permanent shocks is not found in the theory, nor has been taken into account in previous empirical work. This can perhaps explain why several studies have found that wages are little responsive to measures of performance, i.e. that insurance appears to dominate incentives. In fact, ignoring the distinction between transitory and permanent shocks biases the estimate of the pay-per-performance coefficient towards full insurance if transitory shocks are more likely to be insured than permanent shocks (a solid conclusion of our empirical analysis).[42] Third, the supply of insurance depends on firm characteristics other than size (a well known fact of the

---

42 To show this notice that the unexplained component of earnings growth can be written as in equation (26):

$$\Delta\omega_{ijt} = b_u u_{jt} + b_v \Delta v_{ijt} + \varphi_{ijt}$$

which explicitly takes into account the different nature of the shocks to output and the fact that they may have a different impact on wages. The term φ_{ijt} reflects residual unexplained variation not accounted for by shocks. Our empirical analysis suggests that $b_u > 0$ and $b_v = 0$. Thus in this model $\Delta\omega_{ijt} = b_u u_{jt} + \varphi_{ijt}$ and transitory shocks to firm performance do not affect wages because they are smoothed away at the firm level. Suppose that the distinction between transitory and permanent shocks is ignored and that a common sensitivity factor b_u is imposed. In this case:

$$\Delta\omega_{ijt} = b_u \left(u_{jt} + \Delta v_{jt}\right) + \vartheta_{ijt}.$$

The residual term ϑ_{ijt} includes residual unexplained variation not accounted for shocks, φ_{ijt}, and $-b_u \Delta v_{jt}$, a term that reflects the failure to account for the different sensitivity of wages to shocks of different nature. In this simple univariate case the OLS estimate of b_u has probability limit:

$$p\lim \widehat{b}_u = b_u \left(\frac{\sigma_u^2}{\sigma_u^2 + 2\sigma_v^2}\right).$$

The OLS estimate of b_u is therefore downward-biased. The sign of the bias remains negative also when $b_v > 0$ but $b_u > b_v$.

labor market), such as location within an industrial district, which presumably helps employers to disentangle random common fluctuations from idiosyncratic fluctuations in output.

Overall, the firm proves to be an important provider of insurance for individuals. The average standard deviation of wage growth shocks is about 10 percent while that of shocks to value added growth is 30 percent; about one-tenth of the wage variability is due to workers sharing the firm's (permanent) risk. If temporary shocks were transferred to workers in the same proportion as permanent shocks, the variability of earnings would increase by as much as 15 percent.

Appendix

A. The data

A.1 The INPS data set

The Italian National Institute for Social Security (Istituto Nazionale della Previdenza Sociale) requires firms to file 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.

This database covers the universe of employees in the private sector (thus excluding the self-employed, public employees, and off-the books work). Our data set is a sub-sample of the universe, based on workers born on two particular days of the year; our data refer to 1974-1994. The form reports information on annual earnings and on the number of weeks worked. Earnings are divided into two components: normal and occasional. Occasional earnings includes sums drawn from the wage supplementation fund laid-off 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 mangers. Unfortunately, information on education is missing. 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 firm.

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 firms. 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 firm demographics (year of foundation, location, type of organization, ownership status, structure of control, group membership etc.), on employment, and on flow of funds. Balance sheets are reclassified to reduce dependence on the accounting conventions. Balance sheets for the banks' major clients (defined according to the level of borrowing) are collected by the banks. The focus on the level of borrowing skews the sample towards larger firms. Furthermore, because most of the leading banks are in the northern part of the country, the sample has more firms headquartered in the North than in the South. Finally, since banks mainly deal with firms that are creditworthy, firms 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 firms larger than 1,000 employees).

B. 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:

$$\boldsymbol{\Delta\varepsilon}_j = \begin{pmatrix} \Delta\varepsilon_{jT} \\ \Delta\varepsilon_{jT-1} \\ \ldots \\ \Delta\varepsilon_{j1} \end{pmatrix}.$$

If the $\Delta\varepsilon_{jt}$ observation is missing, it is replaced by zero. Conformably with $\boldsymbol{\Delta\varepsilon}_j$, define with $\mathbf{d}_j$ a vector of 0-1 dummy variables. The dummy is 0 if the observation for $\Delta\varepsilon_{jt}$ is missing, 1 otherwise. All the autocovariances of the type $E\left(\Delta\varepsilon_{js}\Delta\varepsilon_{jt}\right)$ are consistently estimated by the sample analogs collected in the following autocovariance matrix:

$$\mathbf{C} = \sum_{j=1}^{F} \boldsymbol{\Delta\varepsilon}_j \boldsymbol{\Delta\varepsilon}_j' ./ \sum_{j=1}^{F} \mathbf{d}_j \mathbf{d}_j'$$

where F is the number of firms always present in the data set and $./$ denotes an element-by-element division.

Define with $\mathbf{m}$ the vector of all the distinct elements of $\mathbf{C}$, i.e. $\mathbf{m} = vech\left(\mathbf{C}\right)$. Since $\mathbf{C}$ is a symmetric matrix, the number of distinct elements in it is $\frac{T(T+1)}{2}$. Conformably with $\mathbf{m}$, define $\mathbf{m}_j = vech\left(\boldsymbol{\Delta\varepsilon}_j \boldsymbol{\Delta\varepsilon}_j'\right)$, and $\mathbf{D} = vech\left(\sum_{j=1}^{M} \mathbf{d}_j \mathbf{d}_j'\right)$. The standard errors of the $\frac{T(T+1)}{2}$ autocovariances can be retrieved by the variance-covariance matrix of $\mathbf{C}$, i.e.:

$$\mathbf{V} = \sum_{j=1}^{M} \left[\left(\mathbf{m}_j - \mathbf{m}\right)\left(\mathbf{m}_j - \mathbf{m}\right)' . * \mathbf{d}_j \mathbf{d}_j'\right] ./\mathbf{DD}'.$$

The standard errors of the estimated moments are simply the square roots of the elements in the main diagonal of $\mathbf{V}$. A similar strategy is used to obtain an estimate of $\mathbf{V}$ (and corresponding standard errors) for workers' earnings.

Table 1

Firms' and workers' characteristics in the full sample

Panel A: Firms' characteristics

	Mean	Median	Stand. dev.
Value added (millions of lire)	16,879	3,935	245,773
Employees	204	60	2,353
South	0.0887	0	0.2844
North	0.7485	1	0.4339
Agriculture and Fishery	0.0028	0	0.0527
Mining	0.0054	0	0.0731
Food and tobacco products	0.0485	0	0.2149
Textiles and leather products	0.1225	0	0.3278
Paper, wood products and publishing	0.0931	0	0.2906
Chemicals and petroleum	0.1308	0	0.3372
Primary and fabricated metal products	0.1080	0	0.3104
Machinery and electrical/electronic	0.1934	0	0.3950
Energy, gas and water	0.0027	0	0.0518
Construction	0.0758	0	0.2647
Retail and wholesale trade, hotels	0.1548	0	0.3617
Transport and telecommunications	0.0254	0	0.1572
Credit, insurance and business services	0.0186	0	0.1351
Other private services	0.0182	0	0.1337

Panel B: Workers' characteristics

	Mean	Median	Stand. dev.
Earnings (millions of lire)	25.49	23.15	11.27
Age	40.52	41	9.88
Male	0.7410	1	0.4381
Blue Collars	0.6393	1	0.4802
Clericals	0.3460	0	0.4657
Managers	0.0146	0	0.1201
South	0.1529	0	0.3599
North	0.6610	1	0.4734

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

Table 2

The value added regression

Variable	Estimate
$\Delta \ln$ (Value added$_{t-1}$)	0.2759 (0.0227)
Center	0.0192 (0.0198)
North	0.0082 (0.0171)
Year dummies	Yes [0.6441]
Sector dummies	Yes [0.4217]
Year*Sector dummies	Yes [0.0000]
Year*Area dummies	Yes [0.0027]
Number of observations	74,359

This table reports the results of the IV regression for value added growth, $\Delta \ln(\text{Value added})_t$. Excluded instruments are lags of (log) value added dated $t-2$ and $t-3$. Values in round brackets are asymptotic standard errors; values in square brackets are p-values.

Table 3

The autocovariance structure of shocks to value added

Order	All years	Order	All years
0	0.1188 (0.0031)	6	0.0002 (0.0010)
1	-0.0435 (0.0022)	7	-0.0024 (0.0012)
2	0.0008 (0.0008)	8	0.0011 (0.0011)
3	0.0001 (0.0008)	9	-0.0042 (0.0023)
4	0.0010 (0.0009)	10	0.0109 (0.0041)
5	0.0006 (0.0008)		

The table 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 $E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right)$. The data are pooled over all years.

Table 4

The earnings equation

Variable	Estimate
$\Delta \ln(\text{earnings}_{t-1})$	0.3309 (0.0345)
Male	0.0063 (0.0010)
Age	0.0016 (0.0004)
$\text{Age}^2/100$	−0.0208 (0.0046)
Productions	−0.0264 (0.0040)
Clerical	−0.0210 (0.0040)
South	−0.0030 (0.0015)
North	−0.0012 (0.0011)
Year dummies	Yes [0.0000]
Number of observations	91,575

This table reports the results of the IV regression for $\Delta \ln(\text{Earnings}_t)$, earnings growth. Excluded instruments are lags of (log) earnings dated $t-3$ and $t-4$. Values in round brackets are asymptotic standard errors; values in square brackets are p-values.

Table 5

The autocovariance structure of shocks to earnings

Order	All years	Order	All years
0	0.0165	10	−0.0002
	(0.0003)		(0.0001)
1	−0.0065	11	0.0002
	(0.0002)		(0.0001)
2	0.0001	12	0.0001
	(0.0001)		(0.0002)
3	0.0003	13	−0.0005
	(0.0001)		(0.0002)
4	−0.0001	14	0.0003
	(0.0001)		(0.0002)
5	−0.0000	15	−0.0004
	(0.0001)		(0.0002)
6	−0.0000	16	0.0001
	(0.0001)		(0.0002)
7	−0.0000	17	−0.0002
	(0.0001)		(0.0003)
8	−0.0001	18	0.0004
	(0.0001)		(0.0005)
9	0.0001		
	(0.0001)		

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 $E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau}\right)$. Data are pooled over all years.

Table 6

Firms' and workers' characteristics in the matched sample

Panel A: Firm characteristics

	Mean	Median	Stand. dev.
Value added (millions lire)	33,966	6,757	320,159
Employees	413	103	3,724
South	0.1098	0	0.3128
North	0.7156	1	0.4512
Agriculture and Fishery	0.0016	0	0.0403
Mining	0.0042	0	0.0649
Food and tobacco products	0.0526	0	0.2234
Textiles and leather products	0.1251	0	0.3309
Paper, wood products and publishing	0.0907	0	0.2872
Chemicals and petroleum	0.1524	0	0.3595
Primary and fabricated metal products	0.1137	0	0.3176
Machinery and electrical/electronic	0.2142	0	0.4103
Energy, gas and water	0.0039	0	0.0623
Construction	0.0562	0	0.2304
Retail and wholesale trade, hotels	0.1183	0	0.3230
Transport and telecommunications	0.0227	0	0.1491
Credit, insurance and business services	0.0221	0	0.1470
Other private services	0.0221	0	0.1470

Panel B: Workers' characteristics

	Mean	Median	Stand. dev.
Earnings (millions of lire)	26.02	23.81	10.90
Age	40.77	41	9.79
Male	0.7627	1	0.4255
Blue Collars	0.6351	1	0.4815
Clericals	0.3523	0	0.4777
Managers	0.0127	0	0.1118
South	0.1345	0	0.3413
North	0.6750	1	0.4684

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

The sensitivity of earnings to value added shocks

Panel A

	Permanent shock	Transitory shock
Sensitivity	0.0599 (0.0154)	0.0072 (0.0051)
J-test (p-value)	0.1161	0.3777
F-test (p-value)	<0.0001	<0.0001

Panel B

Moment	Estimate	Parameter	Estimate
$E\left(\Delta\omega_{ijt}\Delta\omega_{ijt}\right)$	0.0139 (0.0004)	σ_u^2	0.0229 (0.0035)
$E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-1}\right)$	−0.0055 (0.0003)	σ_v^2	0.0334 (0.0048)
$E\left(\Delta\omega_{ijt}\Delta\varepsilon_{jt}\right)$	0.0019 (0.0002)	σ_ξ^2	0.0058 (0.0015)
$E\left(\Delta\omega_{ijt}\Delta\varepsilon_{jt-1}\right)$	−0.0002 (0.0002)	σ_μ^2	0.0034 (0.0011)
$E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt}\right)$	0.0820 (0.0052)	ϱ	−0.1606 (0.0863)
$E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-1}\right)$	−0.0310 (0.0039)	*Ratio*	0.0769

The first row in Panel A reports the IV estimate of the sensitivity of wages to value added shocks (b_u for permanent shocks and b_v for transitory shocks). J-test is the test of overidentifying restrictions. F-test is the test of joint insignificance of excluded instruments. $E\left(\Delta\omega_{ijt}\Delta\omega_{ijt-\tau}\right)$ is an estimate of the autocovariance of wage shocks of order τ; $E\left(\Delta\varepsilon_{jt}\Delta\varepsilon_{jt-\tau}\right)$ an estimate of the autocovariance of value added shocks of order τ; $E\left(\Delta\omega_{ijt}\Delta\varepsilon_{jt-\tau}\right)$ an estimate of the cross-covariance of wage and value added shocks of order τ. σ_u^2, σ_v^2, σ_ξ^2 and σ_μ^2 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: $\frac{b_u\sigma_u}{\sqrt{E\left[(\Delta\omega_{ijt})^2\right]}}$ and measures the amount of earnings variability attributable to value added shocks.

Table 8

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

	Sensitivity to permanent shocks (1)	Sensitivity to transitory shocks (2)
$\Delta\varepsilon_{jt}$	-0.1885	-0.0460
	(0.1367)	(0.0491)
	[0.0417]	[0.2513]
$\Delta\varepsilon_{jt}*$High risk aversion	-0.0764	0.0080
	(0.0304)	(0.0161)
	[0.0706]	[0.1987]
$\Delta\varepsilon_{jt}*$Clerical worker	-0.0026	-0.0130
	(0.0311)	(0.0165)
	[0.0713]	[0.2055]
$\Delta\varepsilon_{jt}*$Manager	0.0815	-0.0344
	(0.1131)	(0.0390)
	[0.0627]	[0.3187]
$\Delta\varepsilon_{jt}*$Worker Age	0.0002	0.0002
	(0.0018)	(0.0006)
	[0.0341]	[0.2497]
$\Delta\varepsilon_{jt}*$s.d.[ln (VA_{jt})]	-0.0338	-0.0041
	(0.0148)	(0.0070)
	[0.0312]	[0.3624]
$\Delta\varepsilon_{jt}*$District	0.0712	0.0065
	(0.0286)	(0.0127)
	[0.0666]	[0.2437]
$\Delta\varepsilon_{jt} *$ [ln (Firm size)]	0.0906	0.0250
	(0.0360)	(0.0160)
	[0.0521]	[0.2179]
$\Delta\varepsilon_{jt} *$ [ln (Firm size)]^2	-0.0065	-0.0028
	(0.0026)	(0.0015)
	[0.0723]	[0.1872]
J-test (p-value)	0.0851	0.1459

The omitted characteristic is the interaction with “Production worker”. Asymptotic standard errors are reported in parenthesis; the partial R^2 for the reduced-form regression is reported in square brackets (see Shea, 1997). J-test is the test of overidentifying restrictions.

References

ABOWD, J., and D. CARD (1989), "On the covariance structure of earnings and hours changes", *Econometrica*, 57, 411-45.

AGGARWAL, R., and A. SAMWICK (1999), "The other side of the trade-off: The impact of risk on executive compensation", *Journal of Political Economy* 107, 65-105.

ALTONJI, J.G., and L.M. SEGAL (1996), "Small-sample bias in GMM estimation of covariance structures", *Journal of Business and Economic Statistics* 14, 353-66.

ANDERSON, T., and C. HSIAO (1982), "Formulation and estimation of dynamic models using panel data", *Journal of Econometrics* 18, 67-82.

ARELLANO, M., and S. BOND (1991), "Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations", *Review of Economic Studies* 58, 277-97.

BERTOLA, G. (1999), "A pure theory of job security and labor income risk", mimeo, European Institute University.

BLANCHFLOWER, D., A. OSWALD, and P. SANFEY (1996), "Wages, profits, and rent-sharing", *Quarterly Journal of Economics* 111, 227-52.

BLUNDELL, R.W., and S. BOND (1999), "Initial conditions and moment restrictions in dynamic panel data models", *Journal of Econometrics* 87, 115-43.

BLUNDELL, R.W., and I. PRESTON (1998), "Consumption inequality and income uncertainty", *Quarterly Journal of Economics* 113, 603-40.

CASAVOLA, P., P. CIPOLLONE, and P. SESTITO (1999) "Caratteristiche delle imprese e caratteristiche dei lavoratori. Da cosa dipendono i salari in Italia?", *Economia e Lavoro* 33, 91-99.

CHAMBERLAIN, G. (1984), "Panel data", in *Handbook of Econometrics*, vol. 2, edited by Z. GRILICHES and M.D. INTRILIGATOR, Amsterdam: North-Holland.

DAVIDSON, R., and J. MACKINNON . (1993), *Estimation and Inference in Econometrics*, New York: Oxford University Press.

DUNNE, T., M. ROBERTS, and L. SAMUELSON (1989) "The growth and failure of U.S. manufacturing plants", *Quarterly Journal of Economics* 104, 671-98.

FARBER, H.S., and R. GIBBONS (1995), "Learning and wage dynamics", *Quarterly Journal of Economics*, 1007-46.

GALIZZI, M., and K. LANG (1998), "Relative wages, wage growth, and quit behavior", *Journal of Labor Economics* 16, 367-91.

GIBBONS, R. (1998) “Incentives in organizations”, *Journal of Economic Perspectives* 12, 115-32

GIBBONS, R., and K. MURPHY (1992), “Optimal incentive contracts in the presence of career concerns: theory and evidence”, *Journal of Political Economy* 100, 468-505.

GIBBONS, R., and M. WALDMAN (1999), “Careers in organizations”, in *Handbook of Labor Economics*, vol. 3b, edited by O. ASHENFELTER and D. CARD, Amsterdam: North Holland.

GOVINDARAJ, S., and R.T.S. RAMAKHRISHNAN (2000), “Accounting earnings processes, intertemporal incentives and their implications for valuation”, mimeo, Columbia University.

GUISO, L., and F. SCHIVARDI (1999),“Information spillovers and factor adjustment”, Banca d’Italia, Temi di Discussione 368.

GUISO, L., and M. PAIELLA (2001), “Risk aversion, wealth and background risk”, CEPR Discussion Paper.

HAMERMESH, D. (1999), “LEEping into the future of labor economics: the research potential of linking employer and employee data”, *Labor Economics* 6, 25-41.

HOLMSTROM, B., and P. MILGROM (1987), “Aggregation and linearity in the provision of intertemporal incentives”, *Econometrica*, 55, 303-28.

JENSEN, M., and MURPHY, K. (1990), “Performance pay and top managment incentives”, *Journal of Political Economy* 98, 225-264

JOVANOVIC, B. (1982), “Selection and the evolution of industry,” *Econometrica*, 50, 649-70.

LAZEAR, E. (1999), “Personnel economics: past lessons and future directions”, *Journal of Labor Economics* 17, 199-236.

MACURDY, T.E. (1982), “The use of time series processes to model the error structure of earnings in a longitudinal data analysis”, *Journal of Econometrics* 18, 82-114.

MALCOLMSON, J. (1997), “Contracts, hold-up, and labor markets”, *Journal of Economic Literature* XXXV, 1916-57.

MARGIOTTA, M., and R. MILLER (2000), “Managerial compensation and the cost of moral hazard”, *International Economic Reiew* 41, 669-719.

MEGHIR, C., and L. PISTAFERRI (2001), “Income variance dynamics and heterogeneity”, mimeo, University College London and Stanford University.

MOULTON, B.R. (1986), “Random group effects and the precision of regression estimates”, *Journal of Econometrics* 32, 385-97.

PRENDERGAST, C. (1999), "The provision of incentives in firms," *Journal of Economic Literature* XXXVII, 7-63.

ROSEN, S. (1985), "Implicit contracts: A survey", *Journal of Economic Literature* XXIII, 1144-75.

ROSEN, S. (1992), "Contracts and the market for executives", in *Contract Economics,* edited by L. WERIN and H. WIJKANDER (eds.), Cambridge, MA: Blackwell.

SCHAEFER, S. (1998), "The dependence of pay-performance sensitivity on the size of the firm", *Review of Economics and Statistics,* 436-43.

SHEA, J. (1997), "Instrument relevance in multivariate linear models: A simple measure", *Review of Economics and Statistics.*

SLEET, C., and S. YELTEKIN (2000) "Dynamic labor contracts with temporary layoffs and permanent separations", mimeo, Northwestern University.

SLITCHER, S. (1950), "Notes on the structure of wages", *Review of Economics and Statistics* 32, 80-91.

RECENTLY PUBLISHED "TEMI" (*)

No. 389 — *The monetary trasmission mechanism: evidence from the industries of five OECD countries,* by L. DEDOLA and F. LIPPI (December 2000).

No. 390 — *Disuguaglianza dei redditi individuali e ruolo della famiglia in Italia,* by G. D'ALESSIO and L. F. SIGNORINI (December 2000).

No. 391 — *Expectations and information in second generation currency crises models,* by M. SBRACIA and A. ZAGHINI (December 2000).

No. 392 — *Unobserved Factor Utilization, Technology Shocks and Business Cycles,* by D. J. MARCHETTI and F. NUCCI (February 2001).

No. 393 — *The Stability of the Relation between the Stock Market and Macroeconomic Forces,* by F. PANETTA (February 2001).

No. 394 — *Firm Size Distribution and Growth,* by P. PAGANO and F. SCHIVARDI (February 2001).

No. 395 — *Macroeconomic Forecasting: Debunking a Few Old Wives' Tales,* by S. SIVIERO and D. TERLIZZESE (February 2001).

No. 396 — *Recovering the Probability Density Function of Asset Prices Using GARCH as Diffusion Approximations,* by F. FORNARI and A. MELE (February 2001).

No. 397 — *A Simple Approach to the Estimation of Continuous Time CEV Stochastic Volatility Models of the Short-Term Rate,* by F. FORNARI and A. MELE (February 2001).

No. 398 — *La convergenza dei salari manifatturieri in Europa,* by P. CIPOLLONE (February 2001).

No. 399 — *Labor Income and Risky Assets under Market Incompleteness: Evidence from Italian Data*, by G. GRANDE and L. VENTURA (March 2001).

No. 400 — *Is the Italian Labour Market Segmented?*, by P. CIPOLLONE (March 2001).

No. 401 — *Optimal Debt Maturity under EMU*, by R. GIORDANO (March 2001).

No. 402 — *Il modello di specializzazione internazionale dell'area dell'euro e dei principali paesi europei: omogeneità e convergenza,* by M. BUGAMELLI (March 2001).

No. 403 — *Seasonality and Capacity: an Application to Italy,* by G. DE BLASIO and F. MINI (June 2001).

No. 404 — *The Effects of Bank Consolidation and Market Entry on Small Business Lending,* by E. BONACCORSI DI PATTI and G. GOBBI (June 2001).

No. 405 — *Money Demand in the Euro Area: Do National Differences Matter?,* by L. DEDOLA, E. GAIOTTI and L. SILIPO (June 2001).

No. 406 — *The Evolution of Confidence for European Consumers and Businesses in France, Germany and Italy*, by P. CARNAZZA and G. PARIGI (June 2001).

No. 407 — *A Primer on Financial Contagion*, by M. PERICOLI and M. SBRACIA (June 2001).

No. 408 — *Correlation Analysis of Financial Contagion: What One Should Know before Running a Test*, by G. CORSETTI, M. PERICOLI and M. SBRACIA (June 2001).

No. 409 — *The Role of the Banking System in the International Transmission of Shocks,* by M. SBRACIA and A. ZAGHINI (June 2001).

No. 410 — *International Transmission Via Trade Links: Theoretically Consistent Indicators of Interdependence for Latin America and South-East Asia,* by C. BENTIVOGLI and P. MONTI (June 2001).

No. 411 — *Why is the Business-Cycle Behavior of Fundamentals Alike Across Exchange-Rate Regimes?,* by L. DEDOLA and S. LEDUC (August 2001).

No. 412 — *Political Institutions and Policy Outcomes: What are the Stylized Facts?,* by T. PERSSON and G. TABELLINI (August 2001).

No. 413 — *Equilibrium Welfare and Government Policy with Quasi-Geometric Discounting,* by P. KRUSELL, B. KURUŞÇU and A. A. SMITH, Jr. (August 2001).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Studi – Divisione Biblioteca e pubblicazioni – Via Nazionale, 91 – 00184 Rome
(fax 0039 06 47922059). They are available on the Internet at www.bancaditalia.it