Discussion Paper No. 09-043

Wage Insurance within German Firms: Do Institutions Matter?

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Non-technical summary: Drawing on a large-scale German Linked Employer-Employee data set, this paper studies the extent to which employers insure workers against firm-level shocks. Particular emphasis is given to the question of whether trade unions and works councils facilitate risk-sharing contracts between workers and firms. Given that the extent of insurance should critically depend on the frequency of the shock, we adopt the identification strategy proposed by Guiso et al. (2005), which enables us to distinguish between transitory and permanent shocks.

In addressing the role of collective bargaining coverage for the amount of wage insurance, our results offer a remarkably consistent picture. Wage insurance is found to be particularly apparent for employers who are subject to collective wage agreements. Moreover, the ability of collective contracts to provide wage insurance appears to decrease with plant size. While in small plants (plant size $\leq$ 100 employees) collective contracts are sufficient on their own to fully insure workers against transitory shocks, they provide only partial insurance in medium-sized (100 $<$ plant size $\leq$ 500) and large plants (plant size $>$ 500). At large employers, the joint existence of collective contracts and works councils helps to provide full insurance against transitory shocks, but provides only partial insurance against permanent shocks. This finding supports the view that the amount of insurance against permanent shocks should be constrained by the possibility of job losses and bankruptcy. The established differences across size classes provide some support for the notion that the degree of information asymmetries is likely to increase with firm size. This should render full insurance under collective contracts at medium-sized and large employers much more difficult and may therefore require the additional existence of a local worker representation. The fact that the latter succeeds in insuring workers only at large employers is consistent with works councils having more formal information rights in large plants.

Das Wichtigste in Kürze: Die vorliegende Studie geht der Frage nach, in welchem Ausmaß Beschäftigte von ihren Arbeitgebern gegen firmenspezifische Produktivitätsschocks versichert werden. Im Mittelpunkt des Interesses steht hierbei die Frage, inwiefern die Existenz von Tarifverträgen sowie Betriebsräten das Ausmaß der Versicherung beeinflusst. Da die Fähigkeit von Arbeitgebern, Beschäftigte gegen persistente Schwankungen zu versichern, erheblich durch mögliche Insolvenzrisiken restringiert sein sollte, unterscheidet die Analyse explizit zwischen permanenten und transitorischen Schocks. Auf Basis deutscher Linked Employer-Employee Daten wird hierzu die von Guiso et al. (2005) vorgeschlagene Identifikationsstrategie angewendet, die eine Identifikation der jeweiligen Reagibilitäten von Löhnen auf kurz- und langfristig wirkende Schocks erlaubt.

Die Ergebnisse der Untersuchung liefern deutliche Evidenz dafür, dass Tarifverträge eine erhebliche Versicherungsfunktion einnehmen, da die individuelle Entlohnung in tarifgebundenen Betrieben schwächer auf Produktivitätsschocks reagiert als die Entlohnung in nicht-tarifgebundenen Betrieben. Die Versicherungsfunktion von Tarifverträgen hängt jedoch erheblich von der Betriebsgröße ab: In kleinen Betrieben (bis zu 100 Beschäftigten) führt die Tarifbindung zu einer vollständigen Entkopplung der Entlohnung von kurzfristigen Produktivitätsschocks, während dies in mittleren (zwischen 100 und 500 Beschäftigten) und großen Betrieben (mehr als 500 Beschäftigte) nicht der Fall ist. In großen Betrieben kann die zusätzliche Existenz von Betriebsräten jedoch dazu beitragen, eine vollständige Versicherung gegen kurzfristige Schocks zu gewährleisten. Die Hypothese, dass kurzfristige Schocks mit Hilfe von Tarifverträgen und Betriebsräten vollständig, langfristige Fluktuationen hingegen nur partiell versichert werden, kann zumindest für die Gruppe der größeren Betriebe bestätigt werden. Das Ergebnis, dass in mittleren und großen Betrieben Tarifverträge allein keine vollständige Versicherung gegen kurzfristige Fluktuationen gewährleisten können, ist möglicherweise auf unterschiedlich große Informationsasymmetrien zurückzuführen, deren Beseitigung einer lokalen Arbeitnehmervertretung bedarf. Dass Betriebsräten dies jedoch nur in großen Betrieben gelingt, ist konsistent damit, dass Betriebsräte gemäß dem Betriebsverfassungsgesetz in großen Betrieben mehr Informationsrechte besitzen.

Wage Insurance within German Firms: Do Institutions Matter?

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July 2009

Abstract

Using a large linked employer-employee data set, this paper studies the extent to which employers insure workers against transitory and permanent firm-level shocks. Particular emphasis is given to the question of whether the amount of wage insurance depends on the nature of industrial relations. Adopting the identification strategy proposed by Guiso et al. (2005), it is shown that wage insurance is particularly apparent for individuals subject to collective wage agreements. While collective contracts alone are sufficient to fully insure workers against transitory shocks in small plants, they provide only partial insurance in medium-sized and large plants. At large employers, the joint existence of collective contracts and works councils helps to provide full insurance against transitory shocks, but provides only partial insurance against permanent shocks. This finding is consistent with the amount of insurance against permanent shocks being constrained by the possibility of considerable job losses and bankruptcy.

Keywords: Wage insurance, linked employer-employee data, collective bargaining

JEL-Code: J31, J51

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1 Introduction

The fact that entrepreneurs may insulate workers' earnings from shocks in the product market has long been recognised as an important determinant of the dynamics of wages. The rationale for such an insurance ultimately rests on the concept of implicit labour contracts originated by Azariadis (1975), Baily (1974) and Gordon (1974). A central empirical implication is that contract wages may entail implicit payments of insurance premiums by workers in favourable states of nature and the receipt of indemnities in unfavourable states.

In the past two decades, a great deal of empirical work has attempted to quantify the extent to which workers' wage dynamics reflect insurance contracts. Early studies date back to Gamber (1988) who uses aggregate U.S. industry-level data. Subsequent work relying on individual data has focused on the question to what extent individuals' wages are affected by external labour market conditions. While much of this work is concerned with aggregate shocks[1], the increasing availability of firm-level and linked employer-employee data has enabled researchers to address the responsiveness of wages to firm-specific conditions. Studies of this sort include e.g. Arai (2003), Hildreth and Oswald (1997) and van Reenen (1996). The firm-level focus provides a more appropriate framework for studying insurance contracts since - due to their idiosyncratic nature - firm-specific as opposed to aggregate shocks constitute diversifiable, and therefore insurable risks. Within the firm-level framework, the methodology has been recently considerably refined by Guiso et al. (2005). Based upon the notion that the extent of insurance should critically depend on the frequency of the shock, the authors propose an identification strategy that aims at explicitly distinguishing between the reaction to transitory and permanent firm-level demand shocks. Using Italian linked employer-employer data, their empirical results suggest that employers provide full insurance against transitory and only partial insurance against permanent shocks. The latter finding is consistent with

[1] The evidence by Beaudry and DiNardo (1991) indicates that workers' wages depend on the tightest labour market conditions since a worker was hired, thereby providing empirical support for an implicit contract framework with worker mobility (for similar evidence see also Grant 2003). Devereux (2005) and Weinberg (2001) also use individual level data and examine the responsiveness of wages to industry-level demand shocks.

the amount of insurance against permanent shocks being constrained by the possibility of bankruptcy. Since then some other authors have replicated their strategy for other European countries. The evidence by Cardoso and Portela (2009) yields similar results for Portugal, whereas the results of Kátay's study (2008) points to considerably less insurance in Hungarian firms.

While much of this recent literature has focused on how the amount of insurance varies across different worker and employer groups, the role of collective bargaining has received somewhat less attention.[2] The scant evidence on collective bargaining is particularly surprising as the role of trade unions as an insurance device has long been emphasised by researchers. The general argument here is that union may mitigate the enforcement problems that arise within risk-sharing agreements between workers and their employers (e.g., Horn and Svensson 1986, Malcolmson 1983). Clearly, examining the trade unions' role in providing wage insurance is crucial to an understanding of how labour market institutions affect wage dynamics.

The purpose of the present paper is therefore to explore whether trade unions facilitate risk-sharing contracts between workers and firms. To do so, we adopt the identification strategy proposed by Guiso et al. (2005) for the case of Germany. The German labour market is particularly interesting as it is characterised by institutions that are widely thought to impose substantial restrictions on the flexibility of wages. A salient feature of the German labour market is the system of widespread collective bargaining coverage. Within this system, regional and industry-wide wage agreements rank among the most important contract type. Moreover, the German labour market is characterised by the coexistence of different wage determination structures, which enables us to exploit these variations to compare outcomes under different bargaining regimes. Recent empirical evidence for Germany shows that centralised contracts decrease the responsiveness of individual wages to firm profits as compared with firm-level contracts and uncovered firms, thereby providing some support for insurance contracts (Guertzgen 2009). As the identification strategy by

[2] E.g., Guiso et al. (2005) and Cardoso and Portela (2005) address the observability of effort as well as individual risk aversion, while Devereux (2005), Grant (2003) and Weinberg (2001) look at gender-specific differences. While Cardoso and Portela (2009) consider wage insurance under firm-level and centralised contracts, they do not compare covered and uncovered firms.

Guiso et al. allows for a distinction between transitory and permanent shocks, we expand on this previous work and examine whether the amount of insurance varies with the frequency of the shock. Given the union's role in facilitating risk-sharing arrangements between workers and their employers, the first natural expectation is that collective wage contracts provide workers with full insurance against transitory shocks at the firm level. As such a full insurance is likely to induce substantial job loss if shocks have a more permanent character, the second hypothesis to be tested is that collective wage contracts should allow for a response to persistent demand shocks.

The data we use to address these questions are taken from a large-scale matched worker-firm data set for Germany, the IAB Linked Employer-Employee data set (LIAB). This data set links the *IAB-Establishment Panel* with individual data for the entire population of workers from the *Employment Statistics Register*. Due to its administrative nature, one of the major advantages of this data set is that it offers very reliable information on individual daily wages inclusive of supplemental pay as long as such pay is subject to social security contributions. Moreover, the data are especially well suited for our purposes as they offer longitudinal information on value added, collective bargaining coverage at the establishment level as well as information on a number of worker and firm characteristics. The latter are particularly important to filter out any systematic variation in workers' wages and firms' value added in order to isolate shocks to firm performance and workers' earnings.

The remainder of the paper proceeds as follows. Section 2 contains a theoretical and institutional background discussion of how the nature of industrial relations may be expected to affect the extent of wage insurance at the firm level. Section 3 presents the empirical analysis. While Section 3.1 to 3.4 provide a description of the data set and a discussion of the basic identification strategy, Section 3.5 and 3.6 present the empirical results. The final Section 4 concludes.

2 Unions and Wage Insurance

2.1 Theoretical Background

The idea that firms may insulate workers' earnings from demand shocks has been formalised by the literature on optimal, or implicit labour contracts (Azariadis 1975, Baily 1974, and Gordon 1974).[3] At the heart of this approach is the view that, due to its long term nature, a labour contract may involve considerable risk-sharing and intertemporal utility smoothing aspects. In these models, differences in risk aversion provide the main theoretical determinant of the amount of insurance provided by the entrepreneur. A general prediction is that, if workers are sufficiently more risk averse than employers, wages will fluctuate less as compared with a pure spot market situation. A major drawback of this approach, however, is that optimal contracts typically require individual workers to have access to an unreasonable amount of information about the technological and product market conditions their employers are confronted with. This apparent deficiency has been taken up by a number of authors who have integrated ideas from the theory on optimal labour contracts with trade union theory. In general, this strand of literature emphasises the trade union's informational role in providing workers with more accurate information on the relevant state of nature and rendering implicit contracts feasible. Models of this sort include the studies by Malcolmson (1983), Horn and Svensson (1986) as well as Hogan (2001). Among these authors Malcolmson (1983) was the first to argue that under product market uncertainties unions may enable workers to enforce state-contingent efficient contracts by removing information asymmetries and imposing collective action upon the employer within an efficient bargaining framework. A similar view is expressed by Horn and Svensson (1986), who consider the union's informational role by combining a monopoly-union set-up with risk-sharing contracts. While the former models identify product market demand shocks as the main source of contracting difficulties, the contribution by Hogan (2001) introduces employers' incentives to cheat on implicit contracts striving to encourage effort provision as the main workers' contracting concern.

[3]For an overview see e.g. Rosen (1985) and Malcolmson (1999).

A second channel through which collectively bargained wages might promote wage insurance relates to the level of wage bargaining. If wages are determined at sectoral or national levels, this should open up less possibilities for local wage adjustments as compared with firm-level wage bargaining. Taken together, the overall view that emerges from these considerations is that collective bargaining may act as a substitute for legal contractual enforcement and may therefore serve as a device to promote implicit contractual arrangements when legal enforcement is otherwise unavailable.

2.2 Institutional Background

As in many other European countries, German wage determination is dominated by collective bargaining agreements. Such collective contracts are generally negotiated between industry-specific trade unions and employers' associations. While legally binding on all member firms of the employers' association and on all employees who are members of the trade union, member firms generally extend the wage settlement to the non-unionised labour force as well. The decision to join an employers' association and to apply such a centralised agreement is generally left to the firms' discretion. An exception is if an agreement is declared to be generally binding by the Federal Ministry of Labour in which case centralised wage contracts may also apply to non-member firms and their employees. Further, there are voluntary extension mechanisms, i.e. firms without any legally binding agreement may voluntarily apply a centralised industry agreement. Finally, a minor fraction of non-member firms are engaged in bilateral negotiations with a trade union and conclude firm-specific agreements. In 2004, the fraction of establishments with a legally binding industry-wide contract was 41 per cent, whereas the fraction of establishments covered by a firm-level contract was 2 per cent in western Germany.[4] Even though industry-level bargaining may be still be viewed as the predominant form of wage determination, the past two decades have seen a clear tendency towards more flexible wage-setting at the firm level. The reason is that contractual opt-out or hardship clauses have become a widespread element of centralised agreements. While opt-out clauses dele-

[4]Source: *IAB Establishment Panel*, own calculations. The figures are reported for 2004, since our data span the time period 1995 to 2004.

gate issues that are usually specified in the central agreement, such as working-time and pay-conditions, to the plant level, hardship clauses enable firms to be exempted from the centralised agreement if they are close to bankruptcy. Moreover, since bargained wages in centralised agreements merely represent a lower bound for wages, there is also sufficient scope for upward flexibility which is reflected in a major fraction of covered firms paying wages above the collectively agreed rates.

In Germany, works councils constitute the second important pillar of the industrial relations system and provide workers with the opportunity of employee representation at the establishment level.[5] The participation rights are laid down under the German *Works Constitution Act* (Betriebsverfassungsgesetz) and include consultation, co-determination and information rights, which generally increase in scope the larger the establishment becomes. For example, Section 106 of the Works Constitution Act obliges plants with more than 100 employees to set up a so-called economic committee in order to provide works councils with all relevant information about their business conditions. According to Section 100, employers with more than 1,000 employees are more formally obliged to do so by recording the required information within each annual quarter. As to wages, even though works councils are formally prohibited from negotiating over issues that are normally dealt with in collective bargaining agreements, they are widely recognised to have a substantial impact on wages for several reasons. First, works councils are traditionally involved in the implementation of collective bargaining agreements at the establishment level and have a consent right with respect to the placement of workers in certain wage groups. Second, works councils may also be expected to play a crucial role in local negotiations over the payment of wages above the collectively agreed rates. Third, since the adoption of opt-out clauses within centralised wage contracts generally requires the approval of the collective bargaining parties and union membership among works councils is typically very high, works councils are also likely to be actively engaged in implementing flexibility provisions at the plant level.[6]

[5]While being legally mandatory in all establishments with at least 5 employees, a local worker representation of this kind only takes institutional form if workers initiate a works council election.

[6]For a more detailed description of the German industrial relations system see Guertzgen (2007).

2.3 Expected Insurance Heterogeneity

As noted in Section 2.1, the trade unions' and works councils' role in removing information asymmetries leads us to expect covered firms and those with a works council to provide more insurance than uncovered firms.

When comparing industry-level with firm-level contracts, it is worth emphasising that the latter are concluded by industry-specific unions. I.e., firm-level contracts in Germany merely involve a different level of bargaining, but do not reflect a fundamentally different union structure. Thus, with respect to the trade union's role in removing information asymmetries, there is a-priori no reason to expect any differential effects under firm and industry-level contracts, as the collection of the relevant firm information ought to be equally easy to deal with under either contract type. The distinctive feature that is relevant here apparently relates to the level of wage determination. To the extent that industry-level wage bargaining makes contracts contingent on sectoral conditions, one might expect the amount of wage insurance to be stronger under centralised contracts as compared with firm-level contracts. However, as a large fraction of firm-level contracts in Germany simply adopts wage bargains negotiated in the corresponding industry agreements (*"Anerkennungstarifverträge"*), the overall differential effect is not clear-cut a-priori.

Clearly, a straightforward implication of wages being determined by sectoral conditions would be that industry-level contracts offer little scope for adjustments of individual wages to firm-specific demand shocks, even if the latter are of permanent nature. However, as demonstrated in Section 2.2, industry-level contracts do not necessarily provide an obstacle to the adjustment of wages to local conditions, as recent decentralisation tendencies in Germany have introduced the option of making such wage adjustments. Given that full insurance is likely to induce substantial job loss if demand disturbances have a more permanent character, a natural expectation is that this potential should at least have been exploited to allow for reactions to permanent firm-level demand shocks.

Given that opt-out clauses allow for wage adjustments even under centralised bargaining, centralised contracts on their own - i.e. without any additional local worker representation - are less likely to provide full wage insurance the larger the degree of information asymmetries at the firm level. The reason is that full insurance would require an industry-level union's knowledge about the technological and product market conditions a single employer is confronted with. In general, one might expect the degree of information asymmetries to increase with the size of the employer. A further hypothesis to be tested, therefore, is that the ability of centralised contracts to provide wage insurance should decrease with firm size and that full insurance at larger employers should require the additional existence of a local worker representation.

Because larger firms are much more likely to be covered by collective bargaining contracts and works councils, a closely related issue concerns the independent role of firm size in providing wage insurance. As firm size is typically viewed as a good proxy for capital market access (e.g., Gertler and Gilchrist 1994), insurance contracts should be particularly apparent for individuals working at larger employers. In our empirical analysis, we will therefore explicitly attempt to sort firm size from industrial relations explanations using evidence on differential effects of collective bargaining across size classes. When addressing this issue, two conflicting hypotheses can be tested. First, it might be conceivable that due to their better credit market access large firms provide more wage insurance than smaller firms irrespective of their collective bargaining status. A countervailing hypothesis is that collective bargaining coverage is used as an explicit device to provide wage insurance and that large firms who choose to stay uncovered might not want to commit themselves to wage insurance and provide no more wage insurance than their smaller counterparts.

3 Empirical Analysis

3.1 Data

The data used in this paper are taken from the IAB Linked Employer-Employee Panel (LIAB) which combines data from the *IAB-Establishment Panel* and the *Employment Statistics Register* (see Alda et al. 2005). The *IAB-Establishment Panel* is based on an annual survey of German establishments, whose sampling frame encompasses all German establishments that employ at least one employee paying social security contributions. New establishments are added to the survey every year to incorporate births and to correct for panel mortality and exits in order to preserve the panel's representative character. The individual data stem from the *Employment Statistics Register,* which is an administrative data set based on reports from employers in compliance with the notifying procedure for the German social security system. This procedure obliges employers to provide a notification at the beginning and the end of each employment relationship for all employees who are covered by the German social security system. In addition, there is at least one annual compulsory notification on the 31^{st} December of each year.

To construct the linked employer-employee data set, we first select establishments from the establishment panel data. From the available waves, we use the years 1995 to 2005. Since information on a number of variables, such as investment expenditures and sales are gathered retrospectively for the preceding year, we lose information on the last wave. Moreover, we restrict our sample to western German establishments from the mining and manufacturing sector with at least two employees. From the establishment level data we gain information on a number of establishment characteristics, such as establishment size, collective bargaining coverage and the existence of a works council. To capture technological differences, we also construct a measure for the capital-labour-ratio. Following Guiso et al. (2005), we use per-capita value added as a proxy for demand shocks, which is constructed as the (per-capita) difference between annual sales and material costs. Table A1 in the appendix provides a detailed description of the construction of the establishment variables. From the establishment data, we first construct a sample in order to identify value added shocks at the plant level. As we apply dynamic panel data methods to identify these shocks, this sample comprises establishments with consistent information on the establishment characteristics of interest and at least three consecutive annual time-series observations. The resulting sample contains 1,354 establishments, yielding an unbalanced panel containing 6,332 establishment-observations with, on average, 4.7 years of data.

In a second step, we merge the establishment panel data with individual data for the entire population of workers who are employed by the selected establishments by using a unique establishment identifier, which is available from both data sets. In particular, the data allow us to merge the selected establishment data with notifications for all employment spells comprising the June 30^{th} of each year. Similar to Guiso et al. (2005), we select our sample so as to focus on stable employment patterns. To do so, we exclude observations for apprentices, part-time and home-workers as well as workers younger than 19 and older than 55 from the individual data. We further eliminate those individuals who move between sample establishments, in order to exclude workers with multiple employers over the observation period.[7] Moreover, since we consider only full-time workers, we eliminate those whose wage is less than twice the lower social security contribution limit. In order to apply dynamic panel data methods we keep those workers who are tracked over at least three consecutive time periods. The resulting sample comprises 435,556 individuals in 1,263 establishments with a total of 2,153,723 individual observations. The individual data provide information on the gross daily wage, age, gender, nationality, employment status (blue/white-collar), educational status (six categories)[8] and on the date of entry into the establishment. Since there is an upper contribution limit to the social security system, gross daily wages are top-coded. In our sample, top-coding affects 14 per cent of all observations. Following Gartner (2005), right-censored observations are replaced by imputed wages. The latter are randomly drawn from a truncated normal distribution whose moments are constructed by the predicted values from Tobit regressions and whose (lower) truncation point is given by the contribution limit to the social security system. Table A2 in the appendix contains a more detailed description of the individual characteristics gained from the *Employment Statistics Register*.

[7] For those workers who separate and do not move between sample establishments, our data unfortunately lack information on their subsequent employment status.

[8] The categories are: No degree, vocational training degree, highschool degree (*Abitur*), highschool degree and vocational training, technical college degree and university degree. Missing and inconsistent data on education are corrected according to the imputation procedure described in Fitzenberger et al. (2006). This procedure relies, roughly speaking, on the assumption that individuals cannot lose their educational degrees.

3.2 Modelling Shocks to Firm Performance

Following Guiso et al. (2005), we isolate idiosyncratic shocks to firm performance by modelling firm performance according to the following process:

$$(1 - \rho L) \cdot y_{jt} = Z'_{jt} \cdot \gamma + \varphi_j + \varepsilon_{jt}, \tag{1}$$

where y_{jt} is the log of per-capita value added in establishment j at time t, which has been deflated by a sector-specific producer price index. L is the lag operator, and the parameter ρ is intended to capture the extent of autroregressive predictable dynamics in the evolution of y_{jt}. In order to control for aggregate non-idiosyncratic shocks, Z'_{jt} includes a full set of time dummies.[9] To capture variation in value added due to changes in capital-input, Z'_{jt} contains as a further control the plant-specific (log) capital-labour ratio. φ_j is a plant-specific fixed effect and ε_{jt} reflects the shock to value added against which firms may insure their employees. Taking first-differences of eq. (1) sweeps out the plant-specific fixed effect and yields:

$$(1 - \rho L) \cdot \Delta y_{jt} = \Delta Z'_{jt} \cdot \gamma + \Delta\varepsilon_{jt}, \tag{2}$$

In eq. (2), first differencing causes the lagged dependent variable Δy_{jt-1} to become correlated with the error term $\Delta\varepsilon_{jt}$, so that it is necessary to instrument lagged value added. In the absence of second-order correlation in the error term, y_{jt-2} and earlier lags provide suitable instruments, since they do not correlate with $\Delta\varepsilon_{jt}$. The same is true for other endogenous variables in $\Delta Z'_{jt}$ which are likely to be correlated with the differenced error term. To estimate eq. (2), we apply the differenced Generalised Methods of Moments (GMM) estimator as proposed by Arellano and Bond (1991). This estimator exploits all available moment conditions around the error term as specified above. Apart from instrumenting endogenous and lagged dependent variables by their lagged values in $t - 2$, the GMM estimator provides an appropriate treatment of predetermined variables which are assumed to be uncorrelated with ε_{jt} and ε_{jt+1}, but are correlated with ε_{jt-1}. These are typically variables whose values in subsequent time periods later than t are likely to be affected by value added in

[9] Other aggregate shocks may be represented by industry-specific and regional dummies. However, the latter are time constant in our data set and are captured by the establishment fixed effect.

period t. As first differencing causes such variables to become correlated with the error term $\Delta\varepsilon_{jt}$, they are instrumented by lagged values in $t-1$ and earlier. To test the validity of the moment conditions, we present the J-Test (generalised Sargan/Hansen test) of overidentifying restrictions. This test statistic calculates the correlation of the error terms with the instrument matrix and has an asymptotic χ^2 distribution under the null that the moment conditions are valid. Moreover, we report diagnostics for second and (higher)-order serial correlation of the error terms (testing the null of no serial correlation).[10]

Table 1 reports the results from the GMM regressions, using y_{jt-2} and y_{jt-3} as instruments for lagged value added. The estimate of ρ is 0.301 with a standard error of 0.066. Moreover, the log of the capital-labour ratio enters the equation with a positive and significant sign. The capital-labour ratio and the time dummies are treated as exogenous variables. A difference Sargan/Hansen test confirms the additional moment restrictions as compared to a specification that treats these variables as predetermined (with a $p-$value of 0.23). Overall, the test statistics in Table 1 indicate that the specifications pass the test of overidentifying restrictions and the AR(2)-Test, thereby confirming the validity of the instruments.

Table 1: Value added GMM regressions

Variable	Coefficient	Standard error
Δlog Value added$(t-1)$	0.301***	(0.066)
Δlog K/L	0.301***	(0.085)
Year dummies ($\chi^2(k)$, $p-$value)	28.57 (*8*)	0.000
J-Test ($\chi^2(k)$, $p-$value)	17.65 (*20*)	0.610
AR(2)-Test ($p-$value)		0.373
AR(3)-Test ($p-$value)		0.862
Observations (Plants)	3,624 (1,354)	

Note: The dependent variable is log (per-capita) value added. Results are reported for the one-step GMM estimator. Robust standard errors are in parentheses. **significant at 5%-level, ***significant at 1%-level.

[10] Note that the GMM estimator may also help to reduce a potential endogeneity problem that arises from measurement error. Measurement error is likely to be relevant since value added and the capital-labour ratio are constructed using the employment level. Thus, measurement error in this variable can induce spurious correlations between the capital-labour ratio and the dependent variable.

In a second step, we use the residuals from the GMM estimations in order to construct a consistent estimate of $\Delta\varepsilon_{jt}$. Table 2 reports estimates of the autocovariances $E(\Delta\varepsilon_{jt}, \Delta\varepsilon_{jt-\tau})$ of the differenced error terms. The figures show that there appears to be no statistically significant correlation at lags greater than one, confirming again the validity of the instruments used in Table 1.

Table 2: Autocovariance structure of value added GMM residuals

Order (τ)	$E(\Delta\varepsilon_{jt}, \Delta\varepsilon_{jt-\tau})$	Standard error
0	0.297***	0.025
1	-.148***	0.023
2	0.012	0.014
3	0.002	0.010
4	0.018	0.010
5	0.030	0.016
6	0.036	0.028

The table reports estimates of the autocovariances $E(\Delta\varepsilon_{jt}, \Delta\varepsilon_{jt-\tau})$ along with their standard errors.
Data are pooled over all years.
**significant at 5%-level, ***significant at 1%-level.

Following Guiso et al. (2005), we specify the error term ε_{jt} as the sum of a transitory and a permanent shock, where the latter is assumed to follow a random walk process. We further assume that the transitory shock follows an $MA(q)$-process, whose order may be recovered from the autocorrelation structure of $\Delta\varepsilon_{jt}$. From Table 2 it can be seen that the estimates of the autocovariances at lags greater than one provide evidence of no large and statistically significant autocorrelation. This leads us to conclude that the autocorrelation structure is in line with an $MA(1)$-process of $\Delta\varepsilon_{jt}$ and an $MA(0)$-process of ε_{jt}. A representation of ε_{jt} that is consistent with the data is therefore

$$\varepsilon_{jt} = \zeta_{jt} + \widetilde{\upsilon}_{jt}, \tag{3}$$

with ζ_{jt} denoting the permanent component which follows a random walk process

$$\zeta_{jt} = \zeta_{jt-1} + \widetilde{u}_{jt}, \tag{4}$$

and $\widetilde{\upsilon}_{jt}$ representing the transitory component which follows an $MA(0)$-process. We further assume $E(\widetilde{u}_{jt}^2) = \sigma_{\widetilde{u}}^2$, $E(\widetilde{\upsilon}_{jt}^2) = \sigma_{\widetilde{\upsilon}}^2$ and $E(\widetilde{u}_{jt}\widetilde{u}_{js}) = E(\widetilde{\upsilon}_{jt}\widetilde{\upsilon}_{js}) = 0$ for $s \neq t$ as well as $E(\widetilde{\upsilon}_{jt}\widetilde{u}_{js}) = 0$ for all s, t. As $\Delta\varepsilon_{jt} \sim MA(1)$ even in the absence of a random walk, we test for the existence of the random walk component by testing whether $E(\Delta\varepsilon_{jt}\left(\sum_{\tau=-1}^{1}\Delta\varepsilon_{jt+\tau}\right)) = 0$. The latter condition amounts to testing the null that $\Delta\varepsilon_{jt} = \Delta\widetilde{\upsilon}_{jt}$ against the alternative that $\Delta\varepsilon_{jt} = \widetilde{u}_{jt} + \Delta\widetilde{\upsilon}_{jt}$.[11] On the basis of the estimated residuals $\Delta\varepsilon_{jt}$ this hypothesis can be rejected (with a p-value < 0.001).

Based upon the established representation of the error term ε_{jt}, value added may be decomposed into a deterministic component, D_{jt}, a permanent, P_{jt}, and a transitory shock, T_{jt}, such that

$$y_{jt} = D_{jt} + P_{jt} + T_{jt}, \tag{5}$$

where $D_{jt} = (1-\rho L)^{-1}(Z'_{jt}\cdot\gamma+\varphi_j)$, $P_{jt} = (1-\rho)^{-1}\zeta_{jt}$ and $T_{jt} = (1-\rho L)^{-1}[\widetilde{\upsilon}_{jt} - (1-\rho)^{-1}\rho\cdot\widetilde{u}_{jt}]$. First-differencing eq. (5) and pre-multiplying by $(1-\rho L)$, eq. (2) can be rewritten as

$$(1-\rho L)\cdot\Delta y_{jt} = \Delta Z'_{jt}\cdot\gamma + (1-\rho L)\cdot u_{jt} + \Delta\upsilon_{jt}, \tag{6}$$

with $u_{jt} = (1-\rho)^{-1}\widetilde{u}_{jt}$ and $\upsilon_{jt} = \widetilde{\upsilon}_{jt} - (1-\rho)^{-1}\rho\cdot\widetilde{u}_{jt}$ denoting the innovations to the permanent and transitory components in eq. (5).

3.3 Modelling Shocks to Individual Earnings

Workers' earnings are modelled according to the following process:

$$\ln w_{ijt} = X'_{ijt}\cdot\delta + \alpha P_{jt} + \beta T_{jt} + \phi_i + \psi_{ijt}, \tag{7}$$

with $i = 1,..., N$ individuals and a total of $N^* = \sum T_i$ total worker-year observations. j refers to the establishment which employs individual i at time t. The dependent variable, $\ln w_{ijt}$, is the individual log gross daily wage. The explanatory variables consist of a vector of covariates, X'_{ijt}, with a coefficient vector δ. X'_{ijt} includes individual and plant-level characteristics (including those captured by D_{jt}) as well as time dummies in order to filter out any systematic variation in workers' wages. To model the dependence of earnings on stochastic shocks to firm performance, the permanent as well as transitory shocks to value added, P_{jt} and T_{jt},

[11] As shown by Guiso et al. (2005) under the alternative hypothesis $E(\Delta\varepsilon_{jt}\left(\sum_{\tau=-1}^{1}\Delta\varepsilon_{jt+\tau}\right)) = \sigma^2_{\widetilde{u}}$ (see also Meghir and Pistaferri 2004).

from eq. (5) are assumed to enter the wage equation with parameters α and β, respectively. Finally, ϕ_i denotes an individual unobserved effect, whereas ψ_{ijt} represents a time-specific error term that is unrelated to any idiosyncratic shocks to firm performance.

On the basis of eq. (5), eq. (7) can be rewritten as

$$\ln w_{ijt} = X'_{ijt} \cdot \delta + \alpha \cdot (1-\rho)^{-1}\zeta_{jt} + \beta \cdot (1-\rho L)^{-1}\upsilon_{jt} + \phi_i + \psi_{ijt}. \tag{8}$$

First-differencing and pre-multiplying eq. (8) by $(1-\rho L)$ gives:

$$\begin{aligned}(1-\rho L)\Delta \ln w_{ijt} &= (1-\rho L)\Delta X'_{ijt} \cdot \delta + \alpha \cdot (1-\rho L) \cdot u_{jt} + \beta \cdot \Delta\upsilon_{jt} + (1-\rho L)\Delta\psi_{ijt} \\ &= (1-\rho L)\Delta X'_{ijt} \cdot \delta + \Delta\omega_{ijt}. \end{aligned} \tag{9}$$

Similar to eq. (2), eq. (9) is estimated by applying the Generalised Method of Moments by Arellano and Bond (1991). X'_{ijt} includes individual covariates such as age and tenure, a quadratic in age and tenure, qualification levels, employment status, as well as the log of plant size and the capital labour ratio. The choice of the instruments for the lagged dependent variable depends on the autocorrelation structure of the differenced error term $\Delta\omega_{ijt}$. Because the AR-Tests indicate that $\Delta\omega_{ijt}$ and $\Delta\omega_{ijt-3}$ are serially uncorrelated, the differenced lagged wage is instrumented using lagged values in $t-3$ and $t-4$. The log of plant size and the capital-labour ratio are treated as endogenous variables and are instrumented by lagged values in $t-2$ and $t-3$.

Table 3 displays the results from the individual earnings GMM estimations. The estimate of the autoregressive coefficient is 0.130 with a standard error of 0.008. With some exceptions, the remaining covariates enter the equation with their expected sign and are significant at conventional levels. Finally, Table 4 reports estimates of the autocovariances $E(\Delta\omega_{ijt}, \Delta\omega_{ijt-\tau})$, where $\Delta\omega_{ijt}$ is constructed from the estimated differenced residuals from eq. (9). The figures show that - except at lag six - there appears to be no large and statistically significant correlation at lags greater than 2, confirming the validity of the instruments used in Table 3.

As the autocorrelation structure of $\Delta\omega_{ijt}$ is consistent with an $MA(2)$-process, a representation of ψ_{ijt} that is consistent with the data is

Table 3: Individual wage GMM regressions

Variable	Coefficient	Standard error
Δlog wage$(t-1)$	0.130***	(0.008)
ΔAge	-.101***	(0.002)
ΔAge^2	-.000***	($6.03e^{-06}$)
ΔTenure	0.010***	($1.2e^{-04}$)
ΔTenure^2	-1.04e^{-06***}	($4.25e^{-08}$)
ΔWhite-collar	0.052***	(0.001)
ΔHighschool	0.011	(0.011)
ΔVocational training	0.006**	(0.003)
ΔVocational training + highschool	0.010**	(0.004)
ΔTechnical college	0.110***	(0.010)
ΔUniversity	0.111***	(0.012)
Δlog Plant size	0.074***	(0.006)
$\Delta\ K/L$	$6.71e^{-06}$	($8.30e^{-06}$)
ΔIndustry-level contract (Cent)	0.008***	(0.001)
ΔFirm-level contract (Firm)	-.002**	(0.001)
ΔWorks council	0.026***	(0.000)
Year dummies ($\chi^2(k)$, $p-$value)	7748.04 (*7*)	0.000
J-Test ($\chi^2(k), p-$value)	8770.09 (*46*)	0.000
AR(2)-Test ($p-$value)	11.78	0.000
AR(3)-Test ($p-$value)	-0.81	0.419
AR(4)-Test ($p-$value)	1.22	0.222
Observations (Individuals)	1,250,755	(435,556)

Note: The dependent variable is the log daily wage. Results are reported for the one-step GMM estimator.
Robust standard errors are in parentheses.
**significant at 5%-level, ***significant at 1%-level.

Table 4: Autocovariance structure of wage GMM residuals

Order (τ)	$E(\Delta\omega_{ijt}, \Delta\omega_{ijt-\tau})$	Standard error
0	0.013***	0.001
1	-0.006***	0.001
2	6.5e^{-04***}	1.6e^{-04}
3	-0.4e^{-04}	0.6e^{-04}
4	-0.7e^{-04}	-0.7e^{-04}
5	-1.9e^{-04}	1.6e^{-04}
6	0.3e^{-04**}	0.1e^{-04}

The table reports estimates of the autocovariances $E(\Delta\omega_{ijt}, \Delta\omega_{ijt-\tau})$ along with their standard errors.
Data are pooled over all years.
**significant at 5%-level, ***significant at 1%-level.

$$\psi_{ijt} = \vartheta_{ijt} + \xi_{ijt}, \tag{10}$$

with ϑ_{ijt} representing a permanent component following a random walk process

$$\vartheta_{ijt} = \vartheta_{ijt-1} + \mu_{ijt}, \tag{11}$$

and ξ_{ijt} denoting a transitory component following an $MA(0)$-process.

3.4 Identification

In this section, we look at the relationship between the residual component of individual wage growth, $\Delta\omega_{ijt}$, and the shock to firm performance, $\Delta\varepsilon_{jt}$. By virtue of eq. (9), $\Delta\omega_{ijt}$ may be specified as

$$\Delta\omega_{ijt} = \alpha \cdot (1 - \rho L) \cdot u_{jt} + \beta \cdot \Delta v_{jt} + (1 - \rho L)\Delta\psi_{ijt}. \tag{12}$$

According to this representation individual wages may respond differently to the transitory and permanent component of firm-specific shocks. As we do only observe a consistent estimate of $\Delta\varepsilon_{jt}$, i.e. the sum of $(1 - \rho L) \cdot u_{jt}$ and Δv_{jt}, identification of α and β requires orthogonality conditions for the residuals of the equations

$$\Delta\omega_{ijt} = \beta \cdot \Delta\varepsilon_{jt} \tag{13}$$

and

$$\Delta\omega_{ijt} = \alpha \cdot \Delta\varepsilon_{jt}. \tag{14}$$

Subtracting $\beta \cdot \Delta\varepsilon_{jt}$ from eq. (12), we obtain

$$\Delta\omega_{ijt} - \beta \cdot \Delta\varepsilon_{jt} = (\alpha - \beta) \cdot (1 - \rho L) \cdot u_{jt} + (1 - \rho L)\Delta\psi_{ijt}. \tag{15}$$

It can be shown that

$$E(\Delta\varepsilon_{jt+1}, \Delta\omega_{ijt} - \beta \cdot \Delta\varepsilon_{jt}) = 0, \tag{16}$$

i.e. in a regression of $\Delta\omega_{ijt}$ on $\Delta\varepsilon_{jt}$, $\Delta\varepsilon_{jt+1}$ and any power $[\Delta\varepsilon_{jt+1}]^k$ with $k \geq 1$ may serve as an instrument for $\Delta\varepsilon_{jt}$ to identify the parameter β, since it is uncorrelated with the residual in eq. (13) and correlated with $\Delta\varepsilon_{jt}$, since $E(\Delta\varepsilon_{jt+1}, \Delta\varepsilon_{jt}) = -\sigma^2_{\widetilde{v}}$.[12]

Correspondingly, under the assumption of covariance stationarity and using the fact that $\Delta\omega_{ijt} - \alpha \cdot \Delta\varepsilon_{jt} = (\beta - a)\ \Delta\nu_{jt}$, a may be identified exploiting the orthogonality condition

$$E(\left(\sum_{\tau=-1}^{1} \Delta\varepsilon_{jt+\tau}\right), \Delta\omega_{ijt} - \alpha \cdot \Delta\varepsilon_{jt}) = 0. \tag{17}$$

Hence, all terms of the form $(\sum_{\tau=-1}^{1} \Delta\varepsilon_{jt+\tau})^m$ with $m \geq 1$, may be used as an instrument for $\Delta\varepsilon_{jt}$ to identify the parameter α.

3.5 Results

We begin by presenting workers' and firms' autocovariances as well as worker-firm cross covariances in Panel A of Table 5 for the matched worker-establishment sample. The estimated values for the moments of the shocks to value added are similar to those reported in Table 1, which are based on the full sample of establishments. From the estimated cross covariances one can see that there appears to be no significant correlation between shocks to workers' wages and establishments' value added for the full sample. The estimated value of $E(\Delta\omega_{ijt}, \Delta\varepsilon_{jt})$ is 0.0002 and not statistically different from zero. The point estimate of $E(\Delta\omega_{ijt}, \Delta\varepsilon_{jt-1})$ is even smaller and is also very imprecisely estimated. Panel B reports estimates of α and β based upon the identification strategy described in Section 3.4. As we use the first three powers of the instruments described in the previous section, we have two overidentifying restrictions for each equation.[13] To estimate α and β we adopt the feasible efficient GMM procedure since the Pagan-Hall-statistic consistently rejects the null of homoskedastic error terms (see also Baum et al. 2003). In Panel B, the estimate for β is 0.011 with a standard error of 0.008. The estimate for α is even negative (with a point estimate of -.0091) and is not statistically significant either (with a standard error of 0.034). In both specifications, the generalised Sargan tests of overidentify-

[12] Note that $\Delta\varepsilon_{jt-1}$ does not serve as an instrument as u_{jt-1} enters eq. (13). The same is true for lags greater than one since they do not correlate with $\Delta\varepsilon_{jt}$ (see Table 2).

[13] Note that we lose some further observations as we use appropriate lags and leads to construct the instruments described in Section 3.4.

Table 5: Responsiveness of wages to value added shocks

A. Autocovariances					
Workers' Autocovariances		Firms' Autocovariances		Worker-Firm Cross Covariances	
$E(\Delta\omega_{ijt}, \Delta\omega_{ijt})$	$S.E.$	$E(\Delta\varepsilon_{jt}, \Delta\varepsilon_{jt})$	$S.E.$	$E(\Delta\omega_{ijt}, \Delta\varepsilon_{jt})$	$S.E.$
0.0127	(0.0012)	0.2997	(0.0266)	0.0002	(0.0009)
$E(\Delta\omega_{ijt}, \Delta\omega_{ijt-1})$	$S.E.$	$E(\Delta\varepsilon_{jt}, \Delta\varepsilon_{jt-1})$	$S.E.$	$E(\Delta\omega_{ijt}, \Delta\varepsilon_{jt-1})$	$S.E.$
-.0061	(0.0008)	0.1513	(0.0244)	$8.0e^{-05}$	(0.0007)

B. Instrumental Variable Estimates				
	Transitory Shock (β)		Permanent Shock (α)	
Sensitivity to shock	0.0111	(0.0081)	-.0091	(0.0335)
J-Test ($p-$value)		0.255		0.337
F-Test ($p-$value)		0.000		0.000
Observations	872,778		581,900	

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors (S.E.) are in parentheses and are adjusted for clustering at the establishment level. ** significant at 5%-level, *** significant at 1%-level.

ing restrictions (J-Test) fail to reject the null hypothesis that the models are not misspecified. (with p-values of 0.255 and 0.337, respectively). Also, the low p-values of the F-Tests confirm that the excluded instruments are jointly significant in the first-stage regressions.

As discussed earlier, capital market access and industrial relations considerations suggest that the extent of insurance might differ systematically by firm size. To test this notion, we split up the sample into workers employed by small (less than or equal to 100 employees), medium-sized (between 100 and 500 employees) and large (more than 500 employees) establishments. The choice of the smallest size class is motivated by the discussion from Section 2.2, which suggests that works councils' information rights are relatively weak in these plants.[14] Panel A of Table 6 presents estimates of the worker-firm cross covariances for the different size classes. The figures strongly indicate that our failure to find evidence of significant cross

[14] In line with the discussion from Section 2.2, one would ideally choose plants with more than 1,000 employees as the largest size class. However, among these plants all employers without collective contracts are covered by a works council, making it impossible to infer the differential insurance effect of works councils versus uncovered plants from our sample establishments.

Table 6: Responsiveness of wages to value added shocks across size classes

A. Cross covariances Across Size Classes

Worker-Firm Cross Covariances	*Small* Size ≤ 100		*Medium* $100 < \text{Size} \leq 500$		*Large* Size > 500	
$E(\Delta\omega_{ijt}, \Delta\omega_{ijt})$	0.0023^{***}	(0.0007)	0.0024^{**}	(0.0006)	-.0002	(0.0009)

B. IV Estimates Across Size Classes

	Small		*Medium*		*Large*	
Explanatory Variable	Transitory Shock (β)					
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
$\Delta\varepsilon_{jt}$	0.0070^{**}	(0.0036)	0.0051	(0.0040)	0.0085	(0.0117)
J-Test ($p-$value)		0.980		0.395		0.440
F-Test ($p-$value)		0.000		0.000		0.000
Exogeneity-Test ($p-$value)						
Observations	17,312		98,656		756,810	
Explanatory Variable	Permanent Shock (α)					
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
$\Delta\varepsilon_{jt}$	0.1103	(0.1154)	-.0665	(0.0544)	-.0030	(0.0326)
J-Test ($p-$value)		0.962		0.649		0.276
F-Test ($p-$value)		0.000		0.000		0.000
Exogeneity-Test ($p-$value)	0.000		0.000		0.000	
Observations	9,726		57,142		515,032	

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors (S.E.) are in parentheses and are adjusted for clustering at the establishment level. **significant at 5%-level, ***significant at 1%-level.

covariances appears to be the result of different patterns across size classes:

While for large employers the estimated moment is negative and not statistically significant, small and medium-sized establishments exhibit significant and positive cross covariances (with point estimates of 0.0023 and 0.0024, respectively). Panel B reports the results from estimating the IV regressions. For small establishments, the response to transitory shocks, parametrised by β, is estimated to be positive and is statistically significant at the 5 per cent level (with a point estimate of 0.0070 and a standard error of 0.0036). The estimates for medium-sized and large firms, in contrast, are also positive but not statistically significant. Referring to the lower part of Panel B, the estimate for α exhibits its expected sign for small plants, but has a very large standard error. In all specifications, the J-Tests confirm the validity of the moment restrictions. Correspondingly, the F-Tests show that the instruments have sufficient explanatory power in the first stage regressions. Overall, the results for the different subsamples show that there appear to be different patterns of wage insurance against transitory shocks in smaller and larger establishments.

We now address the implications of different types of industrial relations for the extent of wage insurance across different size classes. Table A4 in the appendix reports descriptive statistics tabulated by size classes. The figures show that compared with small plants a much larger fraction of medium-sized and large establishments is covered by collective contracts as well as works councils. In order to investigate whether the insignificant responses in Table 6 are driven by suppressed differential industrial relations effects, Table 7 reports results from including interactions between $\Delta\varepsilon_{jt}$ and dummy variables representing the existence of an industry-level ($CENT$), a firm-level contract ($FIRM$) and a works council ($WCOUNCIL$), respectively. The set of instruments is extended by adding the interactions between the original instruments and the respective indicator variables. Referring to the estimates for β in columns (1) and (3) in Table 7, the figures show that industry as well as firm-level contracts appear to suppress the response of workers' wages to transitory shocks in small and medium-sized plants. While the baseline responses in uncovered plants are positive and significant at conventional levels, the coefficients on the interaction terms are estimated to be significantly negative. For both small and medium-sized plants, a Wald test fails to reject the null that the overall effect is zero for both contract types (with p-values of 0.387 and 0.724 for industry-level and 0.384 and 0.438 for firm-level contracts, respectively). Interestingly, among medium-sized plants, works councils appear to have no differential impact on the extent of insurance as the interaction terms are consistently found to be insignificant. This contrasts sharply with small and large plants (Column (5)), where the interaction term of works councils enters the equation with a negative and highly significant coefficient. In column (5), the interaction effect of works councils is more precisely estimated than the (negative) interaction terms of centralised and firm-level contracts. Moreover, a Wald test of $\beta = -\beta\_CENT - \beta\_WCOUNCIL$ confirms the hypothesis that the overall effect is zero (with a p-value of 0.903).

Turning to the responses to permanent shocks in columns (2), (4) and (6), the figures reveal that among medium-sized and large plants, the insignificant estimates of α from Table 6 are also driven by differential collective bargaining effects (Columns (4) and (6)). This contrasts with small plants, for whom the coefficients for α and its interactions with collective bargaining do not exhibit their expected sign and are found to be insignificant (Column (2)). In uncovered medium-sized plants (Column (4)), individuals' wages are found to respond positively to permanent shocks (with a significant point estimate of 0.088). The coefficients on the interaction term of industry-level contracts and works councils show that these institutions appear to significantly reduce the sensitivity of wages to permanent shocks. The same is true for large plants (Column (6)), where the estimate of α for uncovered plants is considerably larger (with a significant point estimate of 0.521). A Wald test of $\alpha = -\alpha\_CENT$ shows that in medium-sized plants centralised contracts are sufficient on their own to suppress the responsiveness to permanent shocks. At large employers, in contrast, only the joint existence of works councils and industry-level contracts helps to fully insure workers against permanent shocks (with a p-value of 0.209).

Table 7: Results by plant size interacted with industrial relations

	Small				*Medium*				*Large*			
	Size $\leq$ 100				100 $<$ Size $\leq$ 500				Size $>$ 500			
Explanatory Variable	Transitory (β)		Permanent (α)		Transitory (β)		Permanent (α)		Transitory (β)		Permanent (α)	
	(1)		*(2)*		*(3)*		*(4)*		*(5)*		*(6)*	
$\Delta\varepsilon_{jt}$	0.0179***	(0.0065)	-.0070	(0.0284)	0.0518***	(0.0158)	0.0881**	(0.0357)	0.2329**	(0.1023)	0.5211**	(0.2246)
		[0.2720]		[0.0398]		[0.1215]		[0.2499]		[0.4261]		[0.0411]
$\Delta\varepsilon_{jt}$_Cent	-.0140***	(0.0052)	-.0113	(0.0156)	-.0463***	(0.0095)	-.0544**	(0.0257)	-.1488	(0.1002)	-.4372**	(0.2244)
		[0.3589]		[0.0882]		[0.1478]		[0.1371]		[0.3828]		[0.0370]
$\Delta\varepsilon_{jt}$_Firm	-.0202***	(0.0071)	0.0093	(0.0358)	-.0369***	(0.0139)	0.0356***	(0.0102)	-.1246	(0.1007)	-.3875*	(0.2277)
		[0.3813]		[0.2020]		[0.1922]		[0.1754]		[0.3781]		[0.0381]
$\Delta\varepsilon_{jt}$_Wcouncil	-.0124**	(0.0062)	0.0081	(0.0287)	-.0019	(0.0157)	-.0632*	(0.0324)	-.0825***	(0.0128)	-.0947***	(0.0304)
		[0.2823]		[0.0411]		[0.1253]		[0.2610]		[0.9321]		[0.6024]
$\beta = -\beta$_Cent, $\alpha = -\alpha$_Cent												
($p-$value)		0.387		0.474		0.724		0.209		0.000		0.000
$\beta = -\beta$_Firm, $\alpha = -\alpha$_Firm												
($p-$value)		0.384		0.905		0.438		0.000		0.000		0.017
$\beta(\alpha) = -\beta(\alpha)$_Cent-$\beta(\alpha)$_Wcouncil												
($p-$value)		0.074		0.542		0.332		0.174		0.903		0.714
$\beta(\alpha) = -\beta(\alpha)$_Firm-$\beta(\alpha)$_Wcouncil												
($p-$value)		0.029		0.771		0.244		0.000		0.123		0.399
J-Test ($p-$value)		0.635		0.701		0.387		0.5180		0.340		0.392
Observations	17,312		9,726		98,656		57,142		756,810		515,032	

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors are in parentheses and are adjusted for clustering at the establishment level. The partial R^2 for the reduced form regression is reported in brackets.
*significant at 10%-level, **significant at 5%-level, ***significant at 1%-level.

In order to compare the responses to transitory and permanent shocks, we also perform an exogeneity test for our regressors, which is based upon the difference in the J-statistic from a model where the regressors are assumed to be exogenous and an alternative model where they are taken as endogenous (see Baum et al. 2003). This test can also be taken as an indirect test for the equality of all coefficients across the models identifying α and β. As the $p-$values from these difference tests are 0.138, 0.128 and 0.183 for small, medium-sized and large plants, respectively, we are not able to formally reject the null that the responses to transitory and permanent shocks do not significantly differ from each other. However, given that the $p-$values still border significance for small and medium-sized plants, we prefer to present separate estimates for α and β. Finally, it is worth mentioning that all specifications pass the $J-$Test of overidentifying restrictions with sufficiently large p-values. Further, as we have multiple endogenous regressors, the terms in brackets report the partial R^2 of the reduced form regressions (see Shea 1997). The figures show that in all cases the power of the instruments is sufficient to identify the parameters of interest.

Taken together, the estimates indicate that - consistent with our expectations - collective contracts in small and medium-sized plants seem to provide full insurance against transitory shocks. Somewhat unexpectedly, a similar result holds for insurance against permanent shocks in medium-sized plants. Moreover, in large plants who are neither covered by a collective contract nor by a works council the sensitivity of wages to shocks is considerably more pronounced than in their small and medium-sized counterparts, with the differences in the coefficients for α and β across size classes being statistically significant. While in large plants collective bargaining coverage alone fails to provide full insurance, the joint coverage by a collective contract as well as a works council helps to fully insure workers against either type of shock.

Thus far, we have only considered insurance heterogeneity induced by different industrial relations. To check whether the pattern of results derived in Table 7 is robust to the inclusion of further interactions, we next turn to the implications of the risk sharing literature that (i) the amount of wage insurance should decrease with the sensitivity of firms' performance to workers' effort (Holmström and Milgrom 1997), (ii) increase with the individuals' degree of risk aversion and (iii) increase with the variability of value added as an inverse measure for the precision of the signal on workers' effort (see also Guiso et al. 2005). To do so, we include interactions between $\Delta\varepsilon_{jt}$ and dummy variables taking on the value of unity for white-collar workers and skilled individuals, whose effort might be expected to be more relevant to firms' performance than that of their blue-collar and low-skilled counterparts. As to risk aversion, the individual data lack explicit information on workers' risk preferences. Recent evidence from the German Socioeconomic Panel suggests that risk aversion is generally higher among females and increases significantly with age (Dohmen et al. 2005). As a proxy for risk aversion, we therefore include interactions between $\Delta\varepsilon_{jt}$ and age as well as a dummy for female workers for whom we expect the amount of insurance to be larger. To measure differential effects with respect to the variability in firm performance, we include also an interaction between $\Delta\varepsilon_{jt}$ and the standard deviation of log real value added over each plant's observation period.

The results from including these additional interactions are shown in Table 8. Overall, the results for these interactions appear to be somewhat mixed, as the estimated coefficients exhibit their expected sign only in some few specifications. For example, while the female interactions enter with their expected (negative) sign in columns (1), (2) and (6), the coefficients are estimated to be significantly positive in columns (3) and (5). The same is true for skilled workers, whose wages are found to be more responsive to transitory shocks only in large plants. A similar picture emerges for white-collar workers who receive less wage insurance in small and large plants only. In a similar vein, the coefficient on the interaction between the variability in value added enters with a negative significant sign only in column (6).

Despite this mixed picture, the clear pattern that emerges from Table 8 is that the inclusion of the additional interactions appears to preserve the industrial relation pattern found in Table 7. In small firms, both contract types are found to provide full insurance against transitory shocks (Column (1)). On the contrary, the coefficients for α and its industrial relations interactions are again found to be insignificant

(Column (2)). In medium-sized and large firms, centralised contracts suppress the responsiveness of wages to transitory shocks - but not to a full extent as a Wald test of $\beta = -\beta\_CENT$ can be rejected at conventional levels (Columns (3) and (5)). The interaction term of works councils is negative, but insignificant in medium-sized plants. This contrasts with large plants, where the interaction terms enters the equation with a significantly negative sign. Even though in large plants the overall response is estimated to be negative (with a point estimate of -0.06 and -0.01 under industry and firm-level contracts), this result supports the view that both contract types along with works councils helps to suppress a positive responsiveness of wages to transitory shocks at large employers. The coefficients on the interaction terms of firm-level contracts in medium-sized plants do not alter their sign, but are estimated with much less precision. Compared with centralised contracts, the role of firm-level contracts seems to be confined to wage insurance against transitory shocks at smaller and large employers. Similar to centralised contracts, firm-level contracts on their own fail to provide full insurance against transitory at large employers as a Wald test rejects the null of $\beta = -\beta\_FIRM$ with a p-value < 0.001 in column (5).

The result that - compared with smaller plants - collective contracts at medium-sized and large employers do not succeed in insuring workers against transitory shocks is supportive of the notion that the degree of information asymmetries is likely to increase with plant size. This should render full insurance under collective contracts at medium-sized and large employers much more difficult and may therefore require the additional existence of a local worker representation. The fact that the latter are able to insure workers against transitory shocks only at large employers (compare Columns (3) and (5)) is perfectly consistent with works councils having more formal information rights in larger plants (see Section 2.2). Note that this result does not hold for insurance against permanent shocks in medium-sized plants, where the interaction effect of works councils is estimated to be significantly negative (Column (4)).

Table 8: IV Estimates with further interactions

Explanatory Variable	*Small* Size ≤ 100 Transitory (β) *(1)*		Permanent (α) *(2)*		*Medium* $100 < \text{Size} \leq 500$ Transitory (β) *(3)*		Permanent (α) *(4)*		*Large* Size > 500 Transitory (β) *(5)*		Permanent (α) *(6)*	
$\Delta\varepsilon_{jt}$	0.0332**	(0.0162)	0.0016	(0.0513)	0.0470***	(0.0106)	0.0719*	(0.0374)	0.1878***	(0.0385)	0.3812*	(0.1998)
		[0.2383]		[0.0181]		[0.1652]		[0.0551]		[0.4026]		[0.0400]
$\Delta\varepsilon_{jt}$_Cent	-.0223***	(0.0053)	-.0151	(0.0116)	-.0226***	(0.0032)	-.0788***	(0.0258)	-.1390***	(0.0386)	-.2805	(0.1977)
		[0.2980]		[0.0489]		[0.2170]		[0.1518]		[0.3807]		[0.0371]
$\Delta\varepsilon_{jt}$_Firm	-.0149**	(0.0052)	0.0371	(0.0297)	-.0133	(0.0096)	0.008	(0.0207)	-.0931**	(0.0407)	-.2901	(0.2027)
		[0.3912]		[0.1501]		[0.2619]		[0.1126]		[0.3774]		[0.0382]
$\Delta\varepsilon_{jt}$_Works council	0.0013	(0.0049)	0.0141	(0.0249)	-.0072	(0.0103)	-.0540*	(0.0283)	-.1076***	(0.0049)	-.0431***	(0.0139)
		[0.3077]		[0.0460]		[0.1316]		[0.1975]		[0.8293]		[0.3801]
$\Delta\varepsilon_{jt}$_Female	-.0097***	(0.0032)	-.0606*	(0.0356)	0.0154***	(0.0025)	0.0249	(0.0261)	0.0126***	(0.0020)	-.0198***	(0.0396)
		[0.2931]		[0.0085]		[0.2575]		[0.0149]		[0.2347]		[0.5503]
$\Delta\varepsilon_{jt}$_Skilled	-.0217***	(0.0067)	-.0696***	(0.0174)	-.0044	(0.0059)	0.0454	(0.0315)	0.0172***	(0.0025)	-.1105***	(0.0267)
		[0.4409]		[0.0495]		[0.3626]		[0.0267]		[0.2637]		[0.0404]
$\Delta\varepsilon_{jt}$_Age	-.0002	(0.0003)	-.0003	(0.0014)	-.0001	(0.0001)	0.0009	(0.0006)	0.0006***	(0.0001)	-.0003	(0.0004)
		[0.2565]		[0.0114]		[0.3123]		[0.0171]		[0.2203]		[0.0274]
$\Delta\varepsilon_{jt}$_White-collar	-.0022	(0.0029)	0.0994***	(0.0290)	-.0087***	(0.0024)	-.0915***	(0.0293)	0.0239***	(0.0020)	0.0197	(0.0146)
		[0.2713]		[0.0093]		[0.2926]		[0.0195]		[0.2481]		[0.0354]
$\Delta\varepsilon_{jt}$_SD_Value	-.0074	(0.0046)	-.0178	(0.0148)	-.0043	(0.0040)	0.0329	(0.0310)	0.0249***	(0.0049)	-.0439*	(0.0234)
		[0.2842]		[0.0577]		[0.4392]		[0.0203]		[0.3602]		[0.1324]
$\beta = -\beta$_Cent, $\alpha = -\alpha$_Cent		0.495		0.801		0.025		0.863		0.000		0.000
$\beta = -\beta$_Firm, $\alpha = -\alpha$_Firm		0.199		0.787		0.016		0.003		0.037		0.032
$\beta(\alpha) = -\beta(\alpha)$_Cent-$\beta(\alpha)$_Wcouncil		0.400		0.990		0.009		0.269		0.224		0.039
$\beta(\alpha) = -\beta(\alpha)$_Firm-$\beta(\alpha)$_Wcouncil		0.157		0.807		0.006		0.403		0.000		0.269
J-Test ($p-$value)		0.672		0.849		0.009		0.608		0.005		0.070
Observations		17,312		9,726		98,656		57,142		756,810		515,032

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors are in parentheses and are adjusted for clustering at the establishment level. The partial R^2 for the reduced form regression is reported in brackets. *significant at 10%-, ** at 5%-, ***at 1%-level.

A possible explanation for this result may relate to the fact that permanent as opposed to transitory shocks are easier to monitor. As a result, despite their weaker information rights, works councils in medium-sized plants might be equally successful in dealing with these shocks as compared with their counterparts at large employers.

Comparing the results with those from Table 7, a further important difference concerns the insignificant interaction terms for α under firm and industry-level contracts and the considerably smaller interaction effect for works councils in large plants (Column (6)). This latter finding suggests that once differences in the amount of insurance according to different individual and employer characteristics are taken into account, collective contracts along with works councils at large employers still provide partial insurance, but fail to provide full insurance against permanent shocks. This result contrasts sharply with medium-sized plants where collective contracts alone succeed in fully insuring workers against permanent shocks (Column (4)). This latter result is somewhat counterintuitive since we expected the amount of insurance against permanent shocks to be constrained by the possibility of considerable job losses and bankruptcy. While the established differences across medium-sized and larger plants might either reflect different preferences for employment or, alternatively, differences in the amount of information asymmetries, our data unfortunately do not allow us to favour either of the two explanations.

3.6 Robustness Checks - Comparison to other Estimates

In this section we conduct some robustness checks and compare our findings to previous results from the literature. As a first restriction, we have excluded all workers who move between sample establishments over the observation period from our estimation sample. To check whether this exclusion biases our results, we reestimated the model after including these movers in our sample. The results corresponding to those in Table 8 are shown in Table A5 in the appendix. For the sake of expositional brevity we confine the presentation to the industrial relations interactions. Even though the inclusion of movers leads to less precise estimates of α in medium-sized plants, the pattern of results is very similar to that in Table 8. Exceptions are firm-level contracts in medium-sized plants (centralised contracts in larger plants), which now significantly reduce the sensitivity of wages to transitory (permanent) shocks. Overall, the baseline point estimates in small and large uncovered plants are somewhat larger than those in Table 8. Even though the differences are not statistically significant, this finding may be taken as weak evidence that movers might be less risk averse than stayers.

A second concern is that we had to impute wages for those workers whose wages are top-coded. To assess the sensitivity of our findings with respect to top-coded wage observations, we re-ran the specifications from Table 8 excluding the observations with imputed wages. The results are shown in Table A6 in the appendix. Even though the estimates of α in medium-sized plants perform somewhat unsatisfactorily, the figures again corroborate the pattern of results that has been found earlier. Exceptions are firm-level contracts in medium-sized and large plants whose interaction effects for α are estimated with more precision. Further, compared with the estimates from Table 8, the baseline point estimates of α and β in uncovered plants turn out to be somewhat smaller - but again not significantly so. A possible explanation for the lower point estimates might relate to the fact that workers with top-coded wage observations are characterised by a larger amount of observed and unobserved productivity. As a result, employers' performance should be more sensitive to the effort of workers whose wages are top-coded, thereby giving rise to a more pronounced responsiveness of wages to value added shocks for this group. Taken together, the above exercises lead us to conclude that the overall pattern of results is fairly robust to the inclusion of movers as well to the exclusion of workers with censored wage information.

Finally, it is interesting to compare our results to other estimates from the literature. Using Italian data, Guiso et al. (2005) find an elasticity of wages to permanent shocks ranging from 0.05 to 0.09, while Cardoso and Portela (2009) report an estimate of 0.09 for Portugal. Both studies find full insurance against transitory shocks. Katay's (2008) results, in contrast, point to a somewhat larger responsiveness of wage to permanent as well as transitory shocks in Hungary, as the author's estimates of α (β) range from 0.07 to 0.12 (0.04 to 0.06). Compared to these figures, our estimated value of α for uncovered medium-sized firms is within a similar range (0.07-0.08), whereas our estimates of α and β for large uncovered firms are rather on the high side. However, it needs to be emphasised that - except for Cardoso and Portela (2009) - the cited studies do not allow the elasticity to vary with collective bargaining coverage and firm size. Thus, our findings for large uncovered firms may reflect the fact that large firms who choose to stay uncovered might not want to commit themselves to wage insurance and provide even less wage insurance than their smaller counterparts.

4 Summary and Conclusions

Drawing on a large-scale Linked Employer-Employee data set, this paper studies the extent to which employers insure workers against firm-level shocks. Particular emphasis is given to the question of whether the amount of wage insurance depends on collective bargaining coverage. Adopting the identification strategy proposed by Guiso et al. (2005), the analysis distinguishes between transitory and permanent shocks. In addressing the role of collective bargaining coverage for the amount of wage insurance, our results offer a remarkably consistent picture. Wage insurance is found to be particularly apparent for employers who are subject to collective wage agreements. Moreover, the ability of collective contracts to insulate workers' wages from shocks appears to decrease with plant size. While in small plants collective contracts alone are sufficient to fully insure workers against transitory shocks, they provide only partial insurance in medium-sized and large plants. At large employers, the joint existence of collective contracts and works councils helps to provide full insurance against transitory shocks, but provides only partial insurance against permanent shocks. Note that this finding is consistent with the amount of insurance against permanent shocks being constrained by the possibility of job losses and bankruptcy. The established differences across size classes provide some support for the notion that the degree of information asymmetries is likely to increase with firm size. This should render full insurance under collective contracts at medium-sized and large employers much more difficult and may therefore require the additional existence of a local worker representation. The fact that the latter help to insure workers particularly at large employers is consistent with works councils having more formal information rights at large employers. In sorting firm size from collective bargaining explanations, we find that large uncovered employers provide even less wage insurance than their smaller counterparts. This lends support to the hypothesis that employers use collective bargaining coverage as an explicit device to provide wage insurance and that large firms who choose to stay uncovered might not want to commit themselves to wage insurance.

Finally, there are potential directions for future research. The established insensitivity of wages to permanent shocks in medium-sized plants raises particular concerns about employers' ability to adjust to shocks in the long run. Future research should address the question as to how the heterogeneity in insurance translates into different amounts of job creation and destruction. A closely related issue concerns the probability of plant closure. As full insurance against permanent shocks holds the risk of bankruptcy, further investigations should go into the long-run employment effects and explore whether a less pronounced responsiveness of wages to permanent shocks is associated with a larger risk of plant closure.

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5 Appendix

Variable	Definition
Value added	Value added is constructed by subtracting material costs from annual sales. Per-capita values are obtained by dividing by average establishment size. The latter is calculated by averaging the number of employees for the month June over the present and preceding year. Nominal values are deflated by the sector-specific producer price index obtained from the German Federal Statistical Office, which is merged to the data based upon a two-digit sector classification.
K/L *Capital-labour ratio*	Constructed by using the perpetual inventory method starting from the capital value in the first observation year and using the information on expansion investments. The initial capital value is proxied by dividing investment expenditures in each establishment's first observation year by a pre-period growth rate of investment, g, and a depreciation rate of capital, d.*) Capital-stocks in subsequent periods are calculated by adding real expansion investment expenditures. Nominal investment expenditures are deflated by the producer price index of investment goods of the Federal Statistical Office Germany. The capital-labour ratio is constructed by dividing the resulting capital proxy by establishment size.
Works council	Dummy=1 if works council is present. In some years (1995 and 1997) only those plants who enter the panel are asked to report the existence of a works council. For the remaining establishments the missing information is imputed based upon the information in the following year.
Firm-level contract	Dummy=1 if establishment is covered by a firm-specific agreement.
Industry-level contract	Dummy=1 if establishment is covered by an industry-specific agreement.

Note: *) To calculate the capital stock in the first period, we set d=0.1 and g=0.05.

Table A1: Construction of establishment variables from the *IAB-Establishment Panel*

Variable	Definition
Gross daily wage	Reported up the contribution limit of the German social security system. Top-coded wages are imputed by drawing from a truncated normal distribution whose moments are generated by Tobit estimations (see Gartner 2005).
Female	Dummy = 1 if female
Highschool degree	Dummy = 1 if highschool (upper secondary) degree
Vocational degree	Dummy = 1 if completed vocational training
Vocational plus highschool	Dummy = 1 if vocational plus highschool degree
Technical college degree	Dummy = 1 if technical college degree
University degree	Dummy = 1 if university degree
White-collar	Dummy = 1 if white-collar worker
Age	Age (in years)
Tenure	End of spell date minus date of entry into the establishment (measured in months)
Foreign	Dummy = 1 if nationality Non-German

Table A2: Description of individual characteristics gained from the *Employment Statistics Register*

Variable	Definition	Mean	Std.-Dev.	Mean	Std.-Dev.
		Individual level		Establishm. level	
Individual characteristics		*(1)*		*(2)*	
ln$w$	Real log daily wage in €	4.626	0.305	4.450	0.248
FEMALE	Female worker	0.167	—	0.212	—
AGE	Age in years	39.422	8.561	39.366	3.230
TENURE	Tenure in months	148.400	89.995	128.951	51.659
FOREIGN	Foreign worker	0.107	—	0.078	—
WHITECOLL	White-collar worker	0.373	—	0.349	—
VOCATIO	Vocational Degree	0.667	—	0.714	—
HIGHSCHOOL	Highschool Degree	0.005	—	0.004	—
VOC-HIGH	Voc. and Highschool Degree	0.034	—	0.030	—
TECHN-UNI	Technical Univ. Degree	0.057	—	0.036	—
UNI	University Degree	0.061	—	0.025	—
Establishment characteristics					
VALUE ADDED	Per-capita value added	0.892	0.439	0.680	0.449
SIZE	Establishment size	5,651.455	10,510.390	513,462	1,771.789
CENT	Centralised agreement	0.844	—	0.673	—
FIRM	Firm-specific agreement	0.104	—	0.088	—
WCOUNCIL	Works council	0.967	—	0.663	—
K/L	Capital-labour ratio	1.059	1.608	0.940	4.354
Individuals		435,556			
Establishments				1,263	

Source: LIAB 1995-2005. 1,263 establishments, 435,556 individuals, 2,153,723 observations.

Note: Per-capita value added and the capital-labour ratio are measured in 100,000 €.

Table A3: Descriptive Statistics

Variable	Definition	Mean	Std.-Dev.	Mean	Std.-Dev.	Mean	Std.-Dev.
		Small		Medium		Large	
		Size ≤ 100		$100 < \text{Size} \leq 500$		Size > 500	
Individual characteristics		*(1)*		*(2)*		*(3)*	
ln$w$	Real log daily wage in €	4.336	0.363	4.498	0.327	4.605	0303
FEMALE	Female worker	0.232	—	0.207	—	0.174	—
AGE	Age in years	39.297	8.568	40.237	8.545	39.337	8.222
TENURE	Tenure in months	117.592	86.698	134.753	92.179	148.876	90.876
FOREIGN	Foreign worker	0.048	—	0.096	—	0.098	—
WHITECOLL	White-collar worker	0.336	—	0.379	—	0.350	—
VOCATIO	Vocational Degree	0.775	—	0.673	—	0.655	—
HIGHSCHOOL	Highschool Degree	0.004	—	0.0043	—	0.003	—
VOC-HIGH	Voc. and Highschool Degree	0.028	—	0.027	—	0.037	—
TECHN-UNI	Technical Univ. Degree	0.019	—	0.042	—	0.052	—
UNI	University Degree	0.014	—	0.026	—	0.051	—
Establishment characteristics							
VALUE ADDED	Per-capita value added	0.562	0.494	0.738	0.508	0.848	0.508
SIZE	Establishment size	38.374	28.013	242.113	103.101	2020,845	3,917.784
CENT	Centralised agreement	0.556	—	0.724	—	0.867	—
FIRM	Firm-specific agreement	0.055	—	0.117	—	0.091	—
WCOUNCIL	Works council	0.322	—	0.885	—	0.988	—
K/L	Capital-labour ratio	1.321	8.358	0.903	1.853	0.965	1.173
Observations		2,483		2,027		1,456	

Source: LIAB 1995-2005. All variables are averages over establishments.

Note: Per-capita value added and the capital-labour ratio are measured in 100,000 €.

Table A4: Descriptive Statistics by Establishment Size

	Small				*Medium*				*Large*			
	Size $\leq$ 100				100 $<$ Size $\leq$ 500				Size $>$ 500			
Explanatory Variable	Transitory (β)		Permanent (α)		Transitory (β)		Permanent (α)		Transitory (β)		Permanent (α)	
	(1)		*(2)*		*(3)*		*(4)*		*(5)*		*(6)*	
$\Delta\varepsilon_{jt}$	0.0514***	(0.0176)	0.0121	(0.0865)	0.0410**	(0.0168)	0.0435	(0.0466)	0.2065***	(0.0575)	0.5296**	(0.2182)
		[0.2382]		[0.0179]		[0.1636]		[0.0547]		[0.4019]		[0.0400]
$\Delta\varepsilon_{jt}$_Cent	-.0237***	(0.0068)	-.0096	(0.0219)	-.0116***	(0.0029)	-.0917***	(0.0253)	-.1301**	(0.0575)	-.3798*	(0.2155)
		[0.2997]		[0.0520]		[0.2137]		[0.1525]		[0.3797]		[0.0371]
$\Delta\varepsilon_{jt}$_Firm	-.0257***	(0.0080)	0.0641	(0.0427)	-.0360***	(0.0086)	-.0019	(0.0228)	-.0541	(0.0582)	-.3548	(0.2167)
		[0.3867]		[0.1496]		[0.2611]		[0.1157]		[0.3768]		[0.0382]
$\Delta\varepsilon_{jt}$_Wcouncil	-.0107	(0.0069)	0.0723**	(0.0354)	0.0040	(0.0173)	-.0800**	(0.0358)	-.1361***	(0.0059)	-.0612***	(0.0148)
		[0.3016]		[0.0457]		[0.1307]		[0.1930]		[0.8299]		[0.3827]
$\beta = -\beta$_Cent, $\alpha = -\alpha$_Cent												
($p-$value)		0.096		0.978		0.092		0.387		0.000		0.000
$\beta = -\beta$_Firm, $\alpha = -\alpha$_Firm												
($p-$value)		0.099		0.493		0.795		0.231		0.000		0.000
$\beta(\alpha) = -\beta(\alpha)$_Cent-$\beta(\alpha)$_Wcouncil												
($p-$value)		0.277		0.457		0.000		0.082		0.000		0.009
$\beta(\alpha) = -\beta(\alpha)$_Firm-$\beta(\alpha)$_Wcouncil												
($p-$value)		0.398		0.244		0.339		0.349		0.064		0.003
J-Test ($p-$value)		0.846		0.852		0.054		0.781		0.010		0.258
Observations	17,489		9,842		99,813		57,791		761,572		518,197	

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors are in parentheses and are adjusted for clustering at the establishment level. The partial R^2 for the reduced form regression is reported in brackets. The specifications correspond to those in Table 8.
*significant at 10%-level, **significant at 5%-level, ***significant at 1%-level.

Table A5: Robustness checks including movers

	Small		*Medium*		*Large*	
	Size $\leq$ 100		100 $<$ Size $\leq$ 500		Size $>$ 500	
Explanatory Variable	Transitory (β)	Permanent (α)	Transitory (β)	Permanent (α)	Transitory (β)	Permanent (α)
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
$\Delta\varepsilon_{jt}$	0.0192 (0.0161)	0.0220 (0.0479)	0.0288** (0.0136)	-.0258 (0.0388)	0.1470*** (0.0449)	0.3763** (0.1704)
	[0.2303]	[0.0196]	[0.1615]	[0.0626]	[0.4059]	[0.0451]
$\Delta\varepsilon_{jt}$_Cent	-.0144** (0.0058)	-.0167 (0.0166)	-.0128*** (0.0035)	-.0671*** (0.0194)	-.0926** (0.0449)	-.2697 (0.1710)
	[0.2780]	[0.0502]	[0.2315]	[0.1527]	[0.3907]	[0.0429]
$\Delta\varepsilon_{jt}$_Firm	-.0104** (0.0051)	-.0044 (0.0244)	-.0044 (0.0094)	0.0422*** (0.0151)	-.0388 (0.0459)	-.2979* (0.1761)
	[0.4054]	[0.1340]	[0.2713]	[0.1088]	[0.3871]	[0.0443]
$\Delta\varepsilon_{jt}$_Wcouncil	-.0005 (0.0043)	-.0014 (0.0193)	-.0111 (0.0130)	-.0323 (0.0261)	-.1117*** (0.0063)	-.0437*** (0.0133)
	[0.2999]	[0.0390]	[0.1276]	[0.2016]	[0.8118]	[0.3530]
$\beta = -\beta$_Cent, $\alpha = -\alpha$_Cent						
($p-$value)	0.733	0.913	0.237	0.046	0.000	0.000
$\beta = -\beta$_Firm, $\alpha = -\alpha$_Firm						
($p-$value)	0.523	0.760	0.132	0.605	0.000	0.087
$\beta(\alpha) = -\beta(\alpha)$_Cent-$\beta(\alpha)$_Wcouncil						
($p-$value)	0.743	0.941	0.410	0.025	0.000	0.028
$\beta(\alpha) = -\beta(\alpha)$_Firm-$\beta(\alpha)$_Wcouncil						
($p-$value)	0.537	0.807	0.176	0.647	0.733	0.452
J-Test ($p-$value)	0.337	0.800	0.410	0.508	0.016	0.373
Observations	16,217	9,098	89,476	51,994	624,255	421,706

The dependent variable is $\Delta\omega_{ijt}$, which is regressed on $\Delta\varepsilon_{jt}$ using the instruments as described in the main text. Standard errors are in parentheses and are adjusted for clustering at the establishment level. The partial R^2 for the reduced form regression is reported in brackets. The specifications correspond to those in Table 8.
*significant at 10%-level, **significant at 5%-level, ***significant at 1%-level.

Table A6: Robustness checks excluding observations with imputed wages