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Rent-Sharing and Collective Wage Contracts - Evidence from German Establishment-Level Data

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Abstract

Using German establishment-level data, this paper analyses whether wages respond to firm-specific profitability conditions. Particular emphasis is given to the question of whether the extent of rent-sharing varies with collective bargaining coverage. In this context, two conflicting hypotheses are tested. The first one asserts that unions exploit their bargaining power at the firm-level and appropriate a larger share of rents than the bargaining parties in uncovered firms. The second one states that unions favour a compressed intra-industry wage structure and suppress the responsiveness of wages to firm-specific profitability conditions. The empirical analysis provides strong support for the second hypothesis. While pooled OLS estimates yield positive estimates of the rent-sharing coefficient in covered establishments, dynamic panel data estimates accounting for unobserved heterogeneity and the endogeneity of rents point to a rent-sharing coefficient of zero.

Keywords: Rent-Sharing, Wage-Setting Structure, Unions, Panel Data
JEL Code: C23, J31, J51

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1 Introduction

The question of whether wages vary systematically with firms' ability-to-pay has long been of considerable interest to labour economists. In the literature, various theoretical explanations have been advanced for a positive relationship between wages and profits (see e.g. Blanchflower et al. 1996, Hildreth and Oswald 1997). Apart from temporary frictions and efficiency wages, a frequently invoked explanation refers to union power. Under collective bargaining, workers' remuneration may be expected to increase with profits, as unions will be able to appropriate part of the industry- or firm-specific rents. Whether wages react to industry- or firm-specific conditions should naturally depend on the level of bargaining. Intuitively, wages ought to be most responsive to firm-specific profitability conditions if wage determination allows for some adjustment to local conditions at the firm-level.

Although the bargaining structure appears to be an important determinant for the degree of rent-sharing at the firm- or industry-level, there is surprisingly little empirical evidence on this topic. While the question of whether wages vary systematically with profits has spawned a vast empirical literature (see e.g. Abowd and Lemieux 1993, van Reenen 1996, Arai 2003, Budd et al. 2005)[1], few studies explicitly address the role of the bargaining structure for rent-sharing. One exception is the study of Holmlund and Zetterberg (1991), who analyse this question based on a cross-country comparison. The authors find countries with highly centralised and coordinated bargaining institutions to exhibit less industry-level rent-sharing than countries with relatively decentralised bargaining systems. In this paper, we draw on establishment-level data from Germany and present some new evidence on rent-sharing and collective bargaining by exploiting intra-national variations in the bargaining structure. Clearly, such variations offer the advantage of controlling for a large part of the unobserved heterogeneity in institutional conditions characterising cross-country comparisons.

The German institutional environment provides a useful example for the coexis-

[1] Further studies include Christofides and Oswald (1992), Blanchflower et al. (1996), Hildreth and Oswald (1997), Abowd et al. (1999), Margolis and Salvanes (2001), Kramarz (2003) and Dobbelaere (2004) amongst others. There are only few previous studies on the relationship between wages and profits in Germany: Hübler and König (1998) and Klodt (2000) use data from the 'Hannover establishment panel' and report a significant positive impact of profits on average firm wages.

tence of different bargaining structures. Until the early 1990s, wage determination was dominated by centralised wage bargaining between industry-specific unions and employers' associations. Those industry agreements were embedded in a corporatist environment characterised by a high degree of coordination (Soskice 1990). However, in the last decade, there has been a strong tendency towards decentralisation of wage determination, since firm-specific collective wage agreements as well as wage determination without any bargaining coverage have become more and more important (Hassel 1999, Ochel 2005). Even within centralised industry agreements, there have been numerous attempts to allow for more (downward) flexibility of wages by introducing opt-out and hardship clauses. Moreover, since bargained wages in centralised agreements merely represent a lower bound for wages, there is also sufficient room for upward flexibility.

Given this intra-national variation in German wage determination regimes, the principal aim of this paper is to shed light on the following questions: Do firm-specific contracts and flexibility provisions in centralised industry agreements allow for rent-sharing at the firm-level? If yes, does the extent of rent-sharing differ from that in firms without any bargaining coverage? A striking feature of the German wage determination process is that decentralisation in collective wage determination merely refers to the *level* of bargaining and not to the *degree of coordination*. The reason is that - as will be discussed below - collective wage determination at the firm-level is generally influenced by industry-wide unions which may retain control over centralised union objectives. Bargaining power considerations lead us to expect the extent of rent-sharing to be the larger the more coordinated the wage-setting process and the more decentralised the level of wage determination. Thus, the first hypothesis to be tested is that unions are able to skim off an even larger part of rents than the bargaining parties in uncovered firms. The question of whether this notion may be confirmed empirically is of considerable interest in an institutional environment such as the German one, which has long been pointed out as corporatist with little scope for excessive rent-sharing at the firm-level. On the other hand, there are various reasons for why unions might favour a compressed intra-industry wage structure, such as high transaction costs or workers' demand for income insurance (see Agell and Lommerud 1992, Burda 1995, Agell 2002). Therefore, a countervailing hypothesis to be tested is that unions suppress any inter-firm wage dispersion due to heterogeneous firm performance.

We investigate the relationship between wages and profitability using the IAB Establishment Panel. This data set is particularly useful for our purposes as it provides detailed information on whether an establishment is subject to an industry-wide wage agreement, a firm-specific wage agreement or to no wage agreement at all. In our estimation strategy, we first focus on simple static pooled Ordinary Least Squares (POLS) estimates. The OLS estimations serve as a benchmark case and will be modified by using dynamic panel data methods. First, we will address the possibility of unobserved firm-specific time invariant factors. A second problem concerns the endogeneity of our profitability measure, since wages and profits are simultaneously determined. Third, we will consider dynamic specifications to allow for possible dynamics in the response of wages to profitability conditions. Finally, we will investigate whether our results are robust to sample selection and the endogeneity of the bargaining structure.

The remainder of the paper is organised as follows: the institutional background of German wage determination is presented in Section 2. Section 3 provides a theoretical discussion to derive testable hypotheses about the extent of rent-sharing under the different wage-setting regimes. These hypotheses are tested in Section 4. While Section 4.1. presents the general empirical model, Section 4.2. describes the data set and the main variables used in the empirical analysis. Section 4.3. reports the estimation results. Section 5 discusses the robustness of the results. Finally, Section 6 provides a discussion and some conclusions.

2 Institutional Background

In Germany, basically three forms of wage determination may be distinguished: central collective wage agreements, firm-specific collective wage agreements as well as wage determination without any collective bargaining coverage. Until the early 1990s, wage determination was dominated by central regional and industry-wide collective wage agreements (*Flächentarifverträge*). Such central wage agreements are negotiated between an industry-specific trade union and an employers' association. They are legally binding on all member firms of the respective employers' association and on all employees who are members of the trade union. Although the negotiated wage applies strictly speaking only to union members, member firms generally extend the wage settlement to non-member employees as well.[2]

Since the early 1990s, the German system of wage determination has experienced a considerable decline in the importance of centralised collective wage agreements (see e.g. Hassel 1999, Kohaut and Schnabel 2003). The tendency towards more decentralisation is caused by three major developments. First, firm-specific collective wage agreements have become more frequent. Those agreements are negotiated between an individual firm and a union. A noteworthy feature of those agreements is that they are concluded by industry-specific unions and do not involve uncoordinated wage bargaining of independent firm-specific unions. That is, decentralisation here merely refers to the *level* of bargaining and not to the *degree of coordination*. Second, wage determination without any bargaining coverage is becoming more relevant. In firms that are not covered by a collective agreement wage determination may either take the form of individual wage contracts or of plant-specific agreements (*Betriebsvereinbarungen*) between works councils and the management.[3] In contrast to firm-specific collective wage agreements, this kind of wage determination can be characterised as decentralised and uncoordinated. Third, there is a tendency even within centralised wage agreements to allow for more flexibility at the firm-level. In recent years, contractual opt-out clauses or hardship clauses have become a widespread element of central agreements. While opt-out clauses delegate 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. In general, the adoption of such clauses requires the approval of the collective bargaining parties (Hassel 1999, Ochel 2005). Moreover, bargained wages in centralised agreements merely represent a lower bound for wages, so that there is always sufficient room for upward flexibility. Even though the wage drift is part of the local negotiations between works councils and the firm, it is also likely to be coordinated by the centralised bargaining parties.

[2] The reason is that non-unionised employees who would receive a lower wage may be expected to join the union anyway in order to benefit from the higher union wage. In addition, central wage agreements may also apply to non-member firms and their employees if the agreement is declared to be generally binding by the Federal Ministry of Labour.

[3] According to the German Works Constitution Act, works councils are not allowed to negotiate about issues that are normally dealt with in collective agreements, even in firms that are not parties of a collective agreement. In practice, however, works councils may be expected to play a crucial role in wage determination (see e.g. Hassel 1999, Hübler and Jirjahn 2003).

The reason is that union density among works councils members is very high (Hassel 1999), and this is particularly relevant for covered firms. Thus, similar to firm-specific collective contracts, the adoption of flexibility provisions in centralised wage agreements is still coordinated by the centralised bargaining parties and involves merely a decentralisation of the level of bargaining.

3 Theoretical Considerations

The purpose of the present section is to derive testable hypotheses about the degree of rent-sharing under the different wage determination regimes. The institutional discussion in Section 2 has yielded two important insights. First, collective contracts do by no means provide an obstacle to adjust wages to local conditions at the firm-level, since recent decentralisation tendencies in Germany have introduced - at least formally - the possibility for such adjustments. Second, even if collective wage determination takes place at the firm-level, it is still influenced by industry-wide unions which may retain control over centralised union objectives.

Thus far, the theoretical literature has mainly focused on the effects of different bargaining regimes on the overall wage level (see e.g. Calmfors and Driffill 1988, Soskice 1990, Dowrick 1993). There is little theory to guide us on the expected effects on the returns to firm-specific attributes such as profits. The rent-sharing literature generally predicts a pay-performance link that depends on the relative bargaining strength of the bargaining parties (e.g. Abowd and Lemieux 1993, Blanchflower et al. 1996, van Reenen 1996). Such considerations lead us to expect the sensitivity of wages to firm-specific profits to be larger under firm-specific contracts than in uncovered firms. An important argument is that firm-specific contracts in Germany are concluded by industry-specific unions, whose bargaining power is likely to be considerably higher than that of works councils determining wages in uncovered firms. This argument is re-enforced by the fact that the wage bargaining process under firm-specific contracts is highly coordinated by an industry-wide union, whereas it is completely uncoordinated in uncovered firms. The bargaining parties in uncovered firms therefore have an incentive to cut wages in order to gain a larger share of industry-demand, and this restricts their ability to raise wages in response to more favourable profitability conditions. With an industry union, this competitive mechanism completely disappears, since a central union may coordinate wage determination at the industry-level.[4] For this reason, one might expect an industry union to capture a larger share of rents under firm-specific contracts than works-councils or individuals in uncovered firms.

The extent of rent-sharing under centralised contracts ultimately depends on whether the bargaining parties make use of flexibility provisions. If such provisions are exploited, the extent of rent-sharing should be larger under industry-contracts than in uncovered firms. The argument here is similar to the reasoning for firm-specific contracts, since the institutional discussion has shown that any adjustment to local conditions at the firm-level is still highly coordinated by the centralised bargaining parties. At this point, it is worthy to note that the question of whether flexibility provisions are used to adjust wages to local profitability conditions still remains to be answered empirically. For example, even though contractual opt-out and hardship clauses have become an important (formal) element of centralised agreements, empirical evidence on the use of such clauses is rather scarce.

Note, in this context, that there are various reasons for why unions (and possibly employers) might favour a compressed intra-industry wage structure. First, transaction costs that are incurred when adjusting wages to firm-level profitability may be high and outweigh any gain involved with wage differentiation. Second, a further rationale for unions to maintain a compressed wage structure might be workers' demand for income insurance (see Agell and Lommerud 1992, Burda 1995, Agell 2002). In our context, intra-industry wage compression provides insurance against two dimensions of uncertainties. First, wage compression between firms at a given point in time may reduce income risk if workers face uncertainties over the allocation to more or less profitable firms. Second, with a compressed intra-industry wage structure wage growth is likely to depend on average sector performance, so that workers' wages at a given employer should also be sheltered against fluctuations in firm-level profitability over time.[5] Thus, the countervailing hypothesis to be tested is that unions suppress inter-firm wage dispersion due to heterogeneous firm performance.

[4] We have formalised this argument elsewhere (Guertzgen 2005).

[5] The issue of wage insurance at the firm level has been taken up recently by Guiso et al. (2005) and Cardoso and Portela (2005). However, these empirical studies do not distinguish different bargaining regimes.

4 Empirical Analysis

4.1 Empirical Model and Testable Hypotheses

In order to quantify the relationship between firm-specific profitability and wages across different wage-setting regimes, we impose a wage equation taking the basic form

$$w_{it} = \alpha + \beta_\pi \cdot \pi_{it} + \gamma \cdot \mathbf{x}'_{it} + \delta \cdot \mathbf{s}'_{it} + f_i + u_{it}. \quad (1)$$

Since we will use establishment level panel data, all variables are subscripted by a establishment-index i and a time index t. The dependent variable, w, is the establishment-specific average wage per worker. The explanatory variable of main interest is π, measuring establishment-specific per-capita profitability.[6] Following the majority of the rent-sharing literature (see e.g. see Abowd and Lemieux 1993, van Reenen 1996), profitability, π, is measured by per-capita quasi-rents. We choose quasi-rents - defined as value-added minus the opportunity cost of labour - for two reasons. First, from a theoretical perspective quasi-rents may be interpreted as representing the 'pie' to be divided between the bargaining parties. Second, from an econometric perspective, the use of quasi-rents instead of profits enables us to circumvent the endogeneity problem induced by the accounting relationship between wages and profits.

In eq. (1), $\mathbf{x}'$ represents a (column) vector of further establishment characteristics with a coefficient vector γ, while $\mathbf{s}'$ denotes a vector of industry characteristics with a coefficient vector δ. For $\mathbf{s}'$ we include the average sectoral wage as well as industry dummies. The latter are supposed to capture industry-specific factors, such as the overall level of industry demand and the degree of competition. The vector of establishment-specific characteristics, $\mathbf{x}'$, includes among other variables dummies for the three wage-setting regimes since the bargaining regime is likely to affect not only the extent of rent-sharing but also the overall wage level. Moreover, $\mathbf{x}'$ contains shares of different skill groups and shares of female workers to control for establishment-specific compositions of the workforce. To account for unobserved differences in worker quality and differences in technologies, further explanatory

[6] Particularly in case of multi-plant firms, one might object that firm-level profitability provides a more appropriate measure than establishment-level profitability. However, since we do only have access to the establishment-level measures, those are taken as a proxy for firm-level profitability.

variables include firm size and the capital-labour ratio. Establishment-specific fixed effects f_i are added to eq. (1) in order to capture unobserved time-invariant factors. Finally, time dummies are included to capture common macroeconomic shocks, and u_{it} is a serially uncorrelated white-noise error term.

Since the emphasis of our analysis is on the impact of different wage-setting regimes on the sensitivity of wages to local profitability conditions, the rent coefficient β_π is specified to depend on the wage-setting regime:

$$\beta_\pi = \beta_0 + \beta_{\pi\_CENT} \cdot CENT_{it} + \beta_{\pi\_DECENT} \cdot DECENT_{it}, \tag{2}$$

where $CENT$ is a dummy taking the value of unity if an establishment adopts a centralised collective wage agreement and $DECENT$ takes on the value of unity if a firm is party to a firm-specific collective wage contract. Recall that according to our first hypothesis, β_{π_DECENT} and β_{π_CENT} should be positive, if firm-specific contracts and flexibility provisions are used to adjust wages to local firm performance. Conversely, testing $\beta_{\pi\_CENT} = -\beta_0$ (and $\beta_{\pi\_DECENT} = -\beta_0$) provides a direct test of the second hypothesis, according to which unions enforce a compressed intra-industry wage structure.

4.2 Data and Variable Description

The empirical analysis uses data from the IAB Establishment Panel. This data set is based on an annual survey of West-German establishments administered since 1993. Eastern German establishments entered the panel in 1996. The data base is a representative sample of German establishments employing at least one employee who pays social security contributions. The survey data provide numerous information on establishment structure and performance, as for example the aggregate wagebill, sales, size and composition of the workforce (see e.g. Bellmann et al. 2002). Moreover, the data contain information on whether an establishment is covered by an industry-wide collective wage agreement, a firm-specific wage agreement or by no collective agreement at all.

In our analysis we use data for the years 1995 to 2002, since detailed information on bargaining coverage is available only from 1995 onwards. Because information on a number of variables, as e.g. sales and the share of materials in total sales are gathered retrospectively for the preceding year, we lose information on the last year.

Moreover, we restrict our sample to mining and manufacturing establishments with at least two employees. We focus on these sectors, since the introduction of opt-out and hardship clauses has been particularly relevant in central collective wage agreements of these industries. These sectors therefore provide a particularly interesting case for testing the empirical relevance of the use of such clauses. As we will apply dynamic panel data methods, only establishments with consistent information on the variables of interest and at least four consecutive time series observations are included in our sample. This results in a sample of 661 establishments with 3,411 observations, yielding an unbalanced panel containing establishment observations with, on average, 5.16 years of data.[7]

The variables used in the subsequent empirical analysis are defined as follows. The dependent variable, w, is defined as the annual aggregate wagebill divided by the number of employees. The number of employees and the wagebill are reported for the month June, where the wagebill is defined exclusive of employers' mandatory social security contributions as well as fringe benefits. Per capita quasi-rents are constructed as the difference between annual sales, material costs and the alternative annual wagebill divided by establishment size, so that

$$\pi = \frac{SALES - MATERIALCOST - \overline{w} \cdot SIZE}{SIZE}. \tag{3}$$

Establishment size ($SIZE$) is calculated as the reported number of employees averaged over the present and preceding year. The alternative wagebill, $\overline{w} \cdot SIZE$, is defined as the annual wagebill which each establishment would incur if it had to pay the average industrial wage. Thus, we approximate $\overline{w}$ by the weighted average of industry-specific annual wages (separately for East and West Germany) for blue- and white-collar workers with the weights being the establishment-specific shares of those worker groups in the total work force.[8] All monetary values are expressed as

[7] Originally, the sample includes 3,546 establishments with consistent information on all the variables of interest. 21 observations were dropped due to suspected errors in the establishment size variable. These observations featured per-capita values of rents of above 1 million DM. For the same reason, 81 observations with a per-capita wagebill of below 8,000 DM were discarded from the sample. This results in a sample of 3,515 establishments with a total of 8,617 observations. Only 661 establishments feature at least four consecutive time-series observations.

[8] We convert average hourly industrial wages of blue collar workers into monthly wages by multiplying them with establishment-specific average working time. Since information on average sectoral wages of white-collar workers is available only on a monthly basis, we

real values by deflating them with a sector-specific producer price index normalised to 1 in 2000. Industry-specific price indices and wages are obtained from the Federal Statistical Office Germany and are matched to the establishment data on the basis of a two-digit sector classification.

Further variables include the share of high-skilled workers (defined as skilled white-collar workers), the share of skilled blue-collar workers, the share of female workers and the share of apprentices in the total work force. Because we do not directly observe the capital stock, we need to construct a proxy. We measure capital by using the perpetual inventory method starting from the capital value in the first observation year and using the information on expansion investment in the following years. 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, δ.[9] Capital-stocks in subsequent periods are calculated by adding real expansion investment expenditures.[10] To obtain real values, nominal investment expenditures are deflated by the producer price index of investment goods of the Federal Statistical Office Germany. The capital-labour ratio, K/L, is constructed by dividing the resulting capital proxy by establishment size. An ownership dummy variable indicates whether the establishment is part of a company owned by persons with unlimited liabilities.

Table 1 presents sample statistics for the main variables used in the subsequent analysis. The figures disclose that quasi-rents vary considerably more than average wages. With respect to collective bargaining coverage, the fraction of observations covered by an industry-wide wage agreement amounts to about 62 per cent, while the fraction of observations with a firm-specific agreement is only 11 per cent. 27 per cent of all observations are subject to no agreement at all. Breaking down the sample into those establishments adopting an industry-wide agreement, a firm-specific

are not able to adjust those wages for establishment-specific average working time. As with the dependent variable, monthly values are converted into annual values by multiplying them with the factor 12.

[9] This involves the assumption that investment expenditures on capital have grown at a constant average rate, g, so that the capital stock in the base year is $K_1 = I_0 + (1-\delta)I_{-1} + (1-\delta)^2 I_{-2} + ... = I_1 \sum_{s=0}^{\infty}[\frac{1-\delta}{1+g}]^s = I_1/(\delta + g)$. In particular, to calculate K_1, we set $\delta = 0.1$ and $g = 0.05$ (see Hempell 2005).

[10] More specifically, $K_t = K_{t-1}(1-\delta) + I_{t-1} = K_{t-1} + EI_{t-1}$,where K_t is the capital stock at the beginning of period t, i.e. at the end of period $t-1$, and EI_t are expansion investment expenditures in period t.

agreement and into those without any bargaining coverage reveals that average wages are highest under industry-wide agreements and lowest without bargaining coverage (see Table A2a in the Appendix). The variability in wages is higher in uncovered establishments with a coefficient of variation of 0.46 as compared to 0.32 and 0.33 in covered establishments. Moreover, establishments under centralised agreements outperform those under firm-specific contracts and those without bargaining coverage in terms of per-capita quasi-rents. Establishments adopting industry-wide agreements also have more employees and exhibit the largest fraction of high-skilled workers, while establishments without bargaining coverage employ on average more women than those covered by a collective wage agreement. Finally, establishments with firm-specific contracts feature the largest capital-labour ratio.

Variable	*Definition*	*Mean*	*Std.-Dev.*	*Obs.*
w	Per-capita wagebill	49.74	18.84	3,411
π	Per-capita quasi-rents	70.57	94.59	3,411
$\overline{w}$	Alternative wage	51.19	11.57	3,411
HIGHSHARE	Share of skilled white-collar workers	0.25	0.20	3,411
BLUESHARE	Share of skilled blue-collar workers	0.42	0.23	3,411
APPSHARE	Share of apprentices	0.05	0.06	3,411
FEMSHARE	Share of female workers	0.27	0.21	3,411
SIZE	Establishment size	605.80	2505.35	3,411
CENT	Centralised collective agreement	0.62	0.49	3,411
DECENT	Firm-specific collective agreement	0.11	0.32	3,411
WCOUNCIL	Works council	0.64	0.48	3,411
K/L	Capital-labour ratio	249.94	1344.08	3,411
EAST	Eastern Germany	0.42	0.49	3,411
OWN	Private ownership	0.21	0.41	3,411

Source: IAB Establishment Panel 1995-2002.
Note: All monetary values are measured in 1,000 DM. 1 € corresponds to 1.95583 DM.

Table 1: Descriptive statistics

4.3 Results

4.3.1 Estimation Strategy

We first focus on a simple static pooled Ordinary Least Squares (POLS) specification of eq. (1). The POLS estimations serve as a benchmark case and will be modified in various respects: first, we will address the possibility of unobserved establishment-specific time invariant factors. In our context, the presence of unobserved heterogeneity may result from neglected capital costs in the rent measure as well as from differences in technological conditions[11] and worker quality that are not captured by our control variables. As such unobserved factors are likely to be correlated with our profitability measure, simple POLS estimates may be expected to yield biased estimates of β_π. A second problem concerns the endogeneity of per-capita rents. A first source of bias is a standard simultaneity bias which occurs if wages, output and quasi-rents are jointly determined. In general, the direction of bias can go either way and largely depends on the underlying relationship between output and employment (see Abowd and Lemieux 1993). In addition, because alternative wages and establishment wages are likely to be positively correlated, there will always be some source of downward bias. Third, we will consider more dynamic specifications and will include lagged wages and rents as explaining variables in our wage regression. The inclusion of lagged rent measures and lagged wages is meant to allow for possible dynamics in the reaction of wages to profitability conditions and sluggish adjustment of wages.

4.3.2 Pooled OLS-Results

Table 2 reports results from POLS estimations of the impact of quasi-rents per worker on wages. The variables are specified in levels rather than logs, since the use of logs would have required discarding all observations with negative quasi-rents. The estimate of quasi-rents per employee on the average wage is 0.042 when including only the alternative wage in the regression. Adding worker characteristics reduces the coefficient to 0.036, suggesting that around 14 per cent of the correlation between rents and wages is due to systematic sorting of workers across establishments (Model (2)). In particular, high-qualified workers appear to be associated with more profitable establishments. The effects of rents on wages are further reduced when including other establishment characteristics, such as establishment size, bargaining coverage, the existence of a works council and ownership status (Model (3)). Apart from $APPSHARE$ (fraction of apprentices), the capital-labour ratio K/L and $DECENT$, all control variables enter the regression with their expected

[11] With respect to differences in technologies, establishment-specific fixed effects capture e.g. production processes that provide firms with higher rents and which may require compensating wage differentials (e.g. processes involving dangerous work). Such differences might lead to a positive wage-rent correlation which would not be due to rent-sharing (see e.g. Margolis and Salvanes 2001).

sign and are all significant at conventional levels. Establishment size is found to have a significant positive effect on average wages, a result which is consistent with earlier evidence.[12] In the literature, various explanations have been advanced for a positive relationship between firm size and wages, such as differences in profits, capital equipment, worker quality and monitoring costs among others (e.g. Oi and Idson 1999). As we control explicitly for differences in the work force composition, the capital-labour ratio and quasi-rents, the establishment size variable may be interpreted as capturing some part of unobserved worker quality and technology differences.

In Model (4), including industry and time dummies leaves the coefficient on rents largely unchanged. Adding industry and time dummies changes the coefficient on the capital-labour ratio to its expected sign, indicating some systematic differences in capital-intensities across industries. Including an east-west dummy does not change the coefficient on rents either (Model (5)). As far as the bargaining coverage effects are concerned, the coefficients on centralised contracts are always significantly positive, whereas decentralised contracts seem to have no significant impact on wages. In addition to the collective bargaining regime, we control for the existence of a works council. Those are found to exert a positive impact on averages wages, which is also in line with earlier studies (see e.g. Addison et al. 2001, Hübler and Jirjahn 2003).

Finally, our main interest concerns the question whether the rent-coefficient differs systematically across the three wage-setting structures. Model (6) includes interactions between collective bargaining coverage and quasi-rents. The inclusion of interactions leads to a larger and more precise estimate of the coefficient on the dummy for firm-level agreements ($DECENT$). In sum, the results indicate that the extent to which wages react to local profitability conditions is significantly lower in establishments that are covered by a collective wage agreement. Even in establishments covered by a firm-specific contract wages appear to be less sensitive to rents. Moreover, the adoption of a centralised wage agreement seems to reduce the magnitude of rent-sharing to a slightly larger extent, as a Wald-Test of $\beta_{\pi\_CENT} = \beta_{\pi\_DECENT}$ can be rejected at the 10 per cent level (with a $p-$value of 0.078). However, the null hypotheses of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$

[12] For German evidence on employer size effects see e.g. Schmidt and Zimmermann (1991) and Gerlach and Hübler (1998).

Model	*1*	*2*	*3*	*4*	*5*	*6*
π	0.042***	0.036***	0.024***	0.025***	0.025***	0.061***
	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.010)
$\pi *$ CENT						-0.042***
						(0.010)
$\pi *$ DECENT						-0.030***
						(0.012)
$\overline{w}$	0.992***	0.815***	0.718***	0.797***	0.650***	0.654***
	(0.024)	(0.031)	(0.031)	(0.038)	(0.065)	(0.064)
HIGHSHARE		11.484***	7.420***	6.727***	10.179***	9.952***
		(1.721)	(1.572)	(1.644)	(1.953)	(1.923)
BLUESHARE		1.433	2.206**	4.308***	4.862***	4.830***
		(1.207)	(1.098)	(1.117)	(1.146)	(1.148)
APPSHARE		-3.183	3.392	7.628**	1.841	1.980
		(3.626)	(3.215)	(3.525)	(3.922)	(3.901)
FEMSHARE		-17.374***	-14.990***	-15.339***	-15.484***	-15.370***
		(1.225)	(1.107)	(1.250)	(1.245)	(1.233)
SIZE			0.001***	0.002***	0.002***	0.002***
			(0.0003)	(0.0003)	(0.0003)	(0.0003)
SIZE^2			-2.17e^{-08}*	-3.03e^{-08}***	-2.94e^{-08}**	-2.91e^{-08}**
			(1.19e^{-08})	(1.16e^{-08})	(1.17e^{-08})	(1.18e^{-08})
CENT			3.772***	4.045***	3.692***	5.619***
			(0.613)	(0.628)	(0.635)	(0.700)
DECENT			0.671	1.195	1.046	2.058**
			(0.754)	(0.745)	(0.744)	(0.886)
WCOUNCIL			7.149***	7.401***	7.116***	7.033***
			(0.632)	(0.631)	(0.632)	(0.627)
K/L			-6.57e^{-06}	0.0002*	0.0002	0.0002
			(-9.55e^{-05})	(0.0001)	(0.0001)	(0.0001)
OWN			-4.108***	-4.267***	-4.225***	-4.189***
			(0.535)	(0.539)	(0.534)	(0.534)
EAST					-2.849***	-2.803***
					(0.995)	(0.988)
Ind.-/Time	No	No	No	Yes	Yes	Yes
Adj. R^2	0.482	0.522	0.597	0.610	0.610	0.615
Observations	3,411	3,411	3,411	3,411	3,411	3,411
Establishments	661	661	661	661	661	661

Note: The dependent variable is the aggregate per-capita wagebill. Heteroscedasticity-robust standard errors are in parentheses. Models (4) - (6) include time dummies and 15 industry dummies.
*Significant at 10%-level ** Significant at 5%-level ***Significant at 1%-level.

Table 2: Pooled OLS regression results

are also rejected (with $p-$values close to zero), suggesting that the overall impact of rents on wages is still positive under both regimes.

4.3.3 Dynamic Specifications

This section addresses potential econometric problems, such as the possibility of unobserved establishment-specific time invariant factors as well as the endogeneity of rents. A further possible endogenous regressor is firm size, as higher wages are likely to induce firms to reduce their labour force. To allow for sluggish adjustment of wages and time lags in the response of wages to profitability conditions, we add lagged wages and quasi-rents as explanatory variables to our regression. The wage equation then takes the following form

$$w_{it} = \alpha + \beta_w w_{it-1} + \sum_{k=0}^{K} \beta_{\pi t-k} \cdot \pi_{it-k} + \gamma \cdot \mathbf{x}'_{it} + \delta \cdot \mathbf{s}'_{it} + f_i + u_{it}, \qquad (4)$$

where the coefficients $\beta_{\pi t-k}$ are specified as in eq. (2). First differencing eq. (4) eliminates time-invariant establishment-specific effects.[13] In eq. (4), first differencing causes the lagged dependent variable Δw_{it-1} to become correlated with the error term Δu_{it}, so that it is necessary to instrument lagged wages. In the absence of second-order correlation in the error term, w_{it-2} and earlier lags will provide suitable instruments, since they will be uncorrelated with Δu_{it}. Because rents, their interactions with the wage-setting regimes and establishment size are likely to be endogenous, they are to be instrumented as well. As with the lagged dependent variable, suitable candidates are lagged rents and establishment size in $t-2$ and earlier provided they do not enter eq. (4) as explanatory variables. Since this might be particularly relevant for lagged rents, we test for the significance of rents up to $t-2$.

To estimate eq. (4), we first apply the differenced Generalized 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

[13] First-differenced estimates of specification (6) in Table 2 yield rent-coefficients of 0.025, -0.034 and -0.018 with standard errors of 0.013, 0.015 and 0.014 (for no-coverage, interactions with centralised contracts and firm-specific contracts, respectively), suggesting a considerable upward bias of the POLS estimates. For the sake of expositional brevity, we do not report the full first-differenced specifications. The estimates are available on request.

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 u_{it} and u_{it+1}, but are correlated with u_{it-1}. As first differencing causes such variables to become correlated with the error term Δu_{it}, they are instrumented by lagged values in $t-1$ and earlier. In particular, we allow all human capital variables, the capital-labour ratio and the alternative wage to be predetermined in order to capture potential feedback effects from wages in period t on those covariates in subsequent periods. To test the validity of the moment conditions, we present the 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-order serial correlation of the error terms (testing the null of no second-order serial correlation).

Table A3 in the Appendix gives the results of the differenced GMM estimates.[14] While Model (1) contains the static specification, Model (2) contains the simplest dynamic specification adding solely the lagged wage to the explanatory variables. Model (3) additionally includes lagged rents, while Model (4) contains lags of rents up to $t-2$. Table A3 contains estimates for time-varying regressors only, since first-differencing eliminates all time-invariant explanatory variables.[15]

Turning to the main variables of interest, the signs of the rent-coefficients exhibit the same pattern as the POLS-estimates of Model (6) in Table 2. While the rent-coefficient is always significantly positive for uncovered establishments, wages appear to be less sensitive to rents in establishments that are covered by a collective wage agreement. Including the lagged wage as a further explanatory variable in Model (2) reduces the rent-coefficients somewhat. As mentioned earlier, using lagged rents in $t-2$ as instruments for contemporaneous rents requires that they do not enter eq. (4) as explanatory variables. To check the robustness of our findings, we therefore include lagged rents up to $t-2$ in Model (3) and (4). While lags of rents in $t-1$ are

[14] All estimations have been carried out using the "XTABOND2"-procedure in STATA 8.0 SE.

[15] In our sample, time-invariant variables are the ownership dummy, the east-west dummy and the industry dummies. The collective bargaining dummies and the works council dummy are time varying binary regressors.

found to be insignificant in Model (3), lagged rents in $t-2$ enter Model (4) significantly, indicating that wages do not only respond to contemporary establishment performance, but also to past profitability conditions. In specifications (3) and (4), the effects of (contemporaneous) rents on wages in uncovered establishments are reduced, but remain still significant, once lagged rents up to $t-2$ are controlled for. Moreover, from the last rows in the second part of Table A3 it can be seen that all specifications pass the test of overidentifying restrictions and the AR(2)-test. The last two rows in the first part of Table A3 report p-values of Wald-statistics testing the null of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$ for the contemporaneous rent coefficients. The values indicate that wages appear to be completely insensitive to profitability conditions in establishments that are covered by a collective agreement, irrespective of whether the agreement is industry or firm-specific.

With respect to the remaining covariates, the performance of the differenced GMM estimates turns out to be rather unsatisfactory: although the lagged wage enters specification (2) and (3) with its expected sign, it is not significant and its point estimates appear to be implausibly low. In Model (4), the estimate is even negative. In all specifications, establishment size and the works-council dummy are always insignificant and for the most part incorrectly signed. The capital-labour ratio is found to be significant, but with a negative sign. As regards the workforce composition, the estimates of $HIGHSHARE$ and $APPSHARE$ also seem to be poorly determined, as they enter almost all regressions with an unexpected sign. The remaining controls for the workforce composition enter with their expected sign (except for $BLUESHARE$ in Model (1)), but are not statistically significant.

In light of the poor performance of the differenced GMM-estimates, Table A4 reports results using the System-GMM (SYS-GMM) estimator as proposed by Arellano and Bover (1995). This estimator is motivated by the problem that lagged levels of a variable are likely to be weak instruments for the equation in first differences if the individual time series exhibits near unit root properties. Closer inspection of the time-series properties of the explanatory variables reveals that particularly the size variable and the capital-labour ratio appear to be close to a random walk.[16] The SYS-GMM estimator exploits additional moment conditions for the equation in levels using lagged differences as instruments in the levels equation. In particu-

[16] SYS-GMM estimates of a simple AR(1)-process yield a coefficient of about 0.94 for establishment size and of 0.91 for the capital-labour ratio.

lar, predetermined variables are instrumented by contemporaneous first differences in the levels equation, whereas endogenous and lagged dependent variables are instrumented by lagged first differences (Bond 2002). To test the additional moment conditions implied by the SYS-GMM estimator as compared to the differenced GMM estimates in Table A3, we present in each column difference tests which refer to the respective specifications in Table A3. The test statistics are calculated as the differences between the Sargan/Hansen statistics of the SYS-GMM and those of the differenced GMM estimates and have an asymptotic χ^2 distribution under the null that the additional moment restrictions are valid.

Overall, the SYS-GMM estimates appear to be more satisfactory than the differenced GMM results. The lagged wage enters all specifications with its expected sign and its estimates are considerably higher than the differenced GMM estimates, suggesting that the latter are severely downward biased. In all specifications, establishment size is found to have a significantly positive impact on average wages and is estimated much more precisely than in the differenced GMM specification. This is consistent with the random-walk property of this variable, indicating that the lagged level of establishment size is a weak instrument for first-differences.

From the human capital covariates, only $FEMSHARE$ and $BLUESHARE$ enter all regressions with their expected sign. The remaining worker controls are mostly incorrectly signed and not significant. Turning to the impact of rents on average wages, the estimates offer a similar picture as the differenced GMM results: in uncovered establishments, quasi-rents exert a positive impact on wages, while wages are generally found to be less sensitive to rents in establishments that are covered by a collective wage agreement. In all specifications, a Wald-Test fails to reject the null of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$. Similar to the differenced GMM estimates, the effects of contemporaneous quasi-rents on wages in uncovered establishments are further reduced but remain still significant, once lagged wages and lagged quasi-rents up to $t-1$ are controlled for. However, controlling for lagged rents up to $t-2$ leads to an insignificant rent-coefficient in uncovered establishments, which is slightly lower than that obtained by the differenced GMM estimates (Model (4)). All specifications pass the test of overidentifying restrictions and the AR(2)-test. Moreover, the difference Sargan/Hansen statistic testing the additional moment restrictions as compared to Table A3 confirms their validity in all specifications except for Model (3) (with a $p-$value of 0.056).

Finally, all specifications were re-run assuming that all predetermined explanatory variables are uncorrelated with the time-invariant establishment-specific effect. At least in specification (4), this causes all human capital and establishment covariates to enter with their expected sign. However, the remaining results from these regressions do only slightly differ from those displayed in Table A4, so that we do not report them here. Most importantly, the estimates of the rent-coefficients are very similar to Table A4. In specifications (2) to (4), a Wald-Test again fails to reject the null of a zero-coefficient on contemporaneous rents under centralised as well as firm-specific agreements ($\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$). Only in Model (1), the null of a zero-coefficient can be rejected under centralised agreements at the 10 per cent level (with a $p-$value of 0.068). In uncovered establishments, the coefficients on contemporaneous rents are slightly larger than those in Table A4, ranging between 0.095 in Model (1) and 0.056 in Model (4).

Comparing the GMM-estimates of the rent-sharing coefficients to the POLS-estimates reveals that the POLS-estimates still yield positive estimates of the rent-sharing coefficient in covered establishments, whereas the SYS-GMM-results accounting for unobserved heterogeneity and endogeneity of rents point to a rent-sharing coefficient of zero. This finding is indicative of the presence of unobserved factors in covered establishments which are positively correlated with profits and impact positively upon wages. One such factor may be that a compressed wage structure under centralised wage contracts causes firms to upgrade the quality of their workforce. This might lead to higher unobserved worker quality in such firms and therefore to upward-biased estimates in the simple POLS-specification. Comparing the GMM-estimates of the rent-sharing coefficients to the POLS estimates in uncovered establishments points to similar figures. While the POLS-coefficient in uncovered establishments amounts to 0.061, the SYS-GMM-estimates range between 0.044 and 0.095. Given these coefficients and mean wages and quasi-rents per employee of 38.78 and 37.15 in uncovered establishments, the elasticity of the average wage with respect to contemporaneous quasi-rents is of the magnitude 0.042 to 0.091. How do these results compare to other estimates for Germany? Hübler and König (1998) use data from the Hannover establishment panel and report an elasticity of about 0.12, while Klodt (2000: pp.172-182) finds an elasticity of 0.14 using the same data set. However, those studies do not allow the rent-coefficient to vary with collective bargaining coverage. Compared to these figures, our estimate of the contemporaneous rent-coefficient in uncovered establishments therefore appears to be rather low. However, given the variability of rents, our results suggest that the quantitative role of rent-sharing in wage determination is nevertheless substantial: calculating the share of variance in the distribution of wages due to the variability in rents, it can be shown that the variability in per-capita rents explains about 15.7 to 33.9 per cent of the variability in (average) establishment wages.[17]

For centralised wage-agreements, the invariance of wages against establishment-specific profitability indicates that the fraction of establishments making use of flexibility provisions seems to be rather negligible. Even though firms may pay wages above the going rate and may adopt opt-out clauses, this potential for adjustments to local profitability conditions appears to be largely unused.[18] Even more striking is the invariance of wages against local profits in establishments that are subject to a firm-specific wage contract. Although this result is to be interpreted with caution as the number of observations with a firm-specific wage contract is rather small, it does not seem to confirm our first hypothesis which led us to expect the sensitivity of wages to profits under firm-specific contracts to be larger than in uncovered establishments. In sum, these findings lend support to our second hypothesis that unions favour a compressed intra-industry wage structure and suppress inter-firm wage differentials.

5 Robustness Checks

5.1 Sample Selection

The use of dynamic panel data methods imposes strong restrictions on the size of our final sample, since we have to exclude all establishments featuring less than 4 consecutive time series observations. Tables A1a and A1b in the Appendix compare sample statistics for the original sample and the final sample used in the preceding

[17] This calculation is performed under the assumption that 95 per cent of the mass of a symmetric distribution is within plus or minus 2 standard deviations of the mean. The contribution of the variability of rents to the variability of wages can then be calculated as:
$\frac{\beta_\pi(\overline{\pi}+2\sigma_\pi)-\beta_\pi(\overline{\pi}-2\sigma_\pi)}{(\overline{w}+2\sigma_w)-(\overline{w}-2\sigma_w)} = \frac{\beta_\pi \cdot \sigma_\pi}{\sigma_w}$ (see e.g. Margolis and Salvanes 2001).

[18] This finding corroborates the results of Franz and Pfeiffer (2003), which are based on an employer survey of about 800 German firms. Their results indicate that only 18 per cent of those employers that covered by a collective contract allowing for hardship clauses make use of such provisions.

analysis. The figures show that establishments subject to a collective contract are on average considerably larger and more capital-intensive in the final sample than in the original sample. The differences for uncovered establishments mainly concern the qualification structure, with a larger fraction of qualified blue-collar employees in the final sample as compared to the original statistics. It is clear that this sample selection might bias our estimates, although the direction of bias is not clear a-priori. For example, unions might want to suppress rent-sharing in large establishments due to high transaction costs. Efficiency wage considerations, which might also play a role in explaining a positive profit effect, lead us to expect the wage-profit correlation to increase with establishment size and capital-intensity. To assess the importance and direction of bias involved with our sample selection, we re-ran the POLS regressions using the original sample of 3,515 establishments.

Re-running the POLS regressions separately by bargaining coverage on the original sample gives point estimates of the rent-coefficients of 0.045, 0.020 and 0.024 (for no-coverage, centralised contracts and firm-specific contracts, respectively) as compared to 0.071, 0.013 and 0.032 for the final sample. However, the differences in the rent-coefficient are statistically significant only for uncovered establishments. This leads us to conclude that for covered establishments our results are fairly robust to sample selection, whereas the selection appears to involve an upward bias of the rent-coefficient for uncovered establishments. Note that this might be caused by the differences in qualification structures. If rent-sharing is more relevant for blue-collar workers, then the overrepresentation of establishments employing large fractions of such employees in the final sample will bias the extent of rent-sharing upwards. However, it should be noted that this does not affect the robustness of our results concerning the overall pattern of wage responses, since for uncovered establishments the rent-coefficient is estimated to be significantly positive, whereas wages in covered establishments are still found to be completely invariant against establishment-specific profitability conditions.

5.2 Alternative Interpretations of the Correlation between Wages and Rents

Several authors have emphasised that a positive correlation between quasi-rents and wages need not necessarily imply rent-sharing, but may simply reflect movements of labour demand along an upward sloping labour supply curve (see e.g. Blanchflower et al. 1996, van Reenen 1996, Hildreth and Oswald 1997). If this were the case, the inclusion of the employment level should render the coefficient on quasi-rents insignificant. For uncovered establishments we are able to rule out such an alternative interpretation, since the positive coefficient on quasi-rents is robust to the inclusion of establishment size as an explanatory variable in all regressions. As a further robustness check we have also included employment growth a proxy for demand shocks in the differenced GMM regressions, which left the coefficients on quasi-rents also largely unchanged.[19]

5.3 Endogeneity of the Bargaining Regime

Thus far, we have considered the collective bargaining regime as exogenous. However, in Germany firms may leave their employers' associations and may, thus, to some extent influence the choice of the bargaining regime. This shows up in our data, where the share of establishments subject to a centralised contract declined from 82 per cent in 1995 to 75 per cent in 2001, and the fraction of establishments with a firm-specific agreement decreased from 8.5 to 6 per cent. What is relevant for our estimates is that a non-random selection into the regimes might bias our rent-sharing coefficients, particularly if a firm's choice is correlated with its profitability conditions. If, for example, centralised contracts shelter firms against excessive rent-sharing at the firm-level, highly profitable firms might systematically select themselves into the centralised regime. To check the robustness of our findings to the endogeneity of the collective bargaining regime, we first estimate eq. (1) separately by bargaining coverage using a selection model which accounts for a potential non-random selection into the three wage determination regimes. To assess the importance of a potential endogeneity bias, we subsequently compare the estimates with the corresponding POLS regression results. Defining regimes R_1, R_2 and R_3 as wage determination under no-coverage, centralised contracts and firm-specific contracts, respectively, the wage equations for each regime R_j, $j = 1, 2, 3$, become

$$w_j = \alpha_j + \beta_{j\pi} \cdot \pi_j + \gamma_j \cdot \mathbf{x}'_j + \delta_j \cdot \mathbf{s}'_j + u_j \quad \text{if } R_j = 1, \tag{5}$$

where the variance of u_j is given by σ^2 and the indices i, t are suppressed for expositional convenience. To account for $E(u_j|R_j = 1)$, we adopt an extension of

[19] The results are not reported here, but are available on request.

the two-step selection-bias correction method developed by Heckman (1979), which has been proposed by Lee (1983). Assuming that selectivity into the regimes can be modelled as a multinomial logit with a vector of explanatory variables $\mathbf{z}'$ and a parameter vector θ_j, Lee (1983) shows that

$$E(w|R_j = 1) = \alpha_j + \beta_{j\pi} \cdot \pi_j + \gamma_j \cdot \mathbf{x}'_j + \delta_j \cdot \mathbf{s}'_j - \sigma\rho_j\lambda_j(\theta_j \cdot \mathbf{z}'). \qquad (6)$$

where

$$\lambda_j(\ \theta_j \cdot \mathbf{z}') \ = \frac{\phi(\Phi^{-1}(P_j))}{P_j},\ j = 1, 2, 3 \text{ and } P_j = \frac{\exp(\theta_j \cdot \mathbf{z}')}{\sum_k \exp(\theta_k \cdot \mathbf{z}')},\ j, k = 1, 2, 3, \qquad (7)$$

and with ρ_j denoting the correlation-coefficient between u_j and the unobservables in the selection equation (eq. (8) in the Appendix). A more detailed exposition of this selectivity correction can be found in Appendix A1. From eq. (6) it can be seen that OLS estimates of eq. (5) are biased if $\lambda_j(\theta_j \cdot \mathbf{z}')$ is correlated with the observables in eq. (5) and u_j and the error term in the selection equation are correlated. Consistent estimates of all parameters of interest in eq. (6) may be obtained by a two-step procedure, where the first step involves the generation of predicted values of $\lambda_j(\theta_j \cdot \mathbf{z}')$ by estimating the selection equation using a multinomial logit approach. In the second step, the predicted values are added to eq. (5), which is estimated by OLS. In the selection equation, the vector of observables $\mathbf{z}'$ includes the observables in eq. (5) and further identifying covariates which are excluded from eq. (5). For the excluded observables, we choose (1) a dummy taking on the value of unity if an establishment belongs to a publicly listed company and (2) an establishment-age dummy indicating whether an establishment has been founded after 1990 or earlier. We believe that those variables provide appropriate identifying exclusion restrictions for several reasons. First, when testing the significance of those variables in eq. (5) estimated separately by bargaining coverage, the corresponding F-tests indicate that both variables have no direct significant impact on wages (with p-values of 0.18, 0.24 and 0.44). Second, the identifying variables all appear to be significant predictors of the bargaining regime, since the corresponding F-statistic is highly significant in the selection equation (with a p-value close to zero). Third, we argue that is reasonable to assume that the identifying variables are exogenous in the selection equation, since they are unlikely to be influenced by unobservables affecting the bargaining regime.

Table A5 in the Appendix shows the POLS estimates and the selectivity-corrected estimates using the original sample of 3,515 establishments. The negative coefficients on $\lambda_j(\theta_j \cdot \mathbf{z}')$, which are estimates of $-\sigma\rho_j$, indicate that the choice of collective contracts is endogenous, with the error term in the selection equation and u_j being positively correlated. In the uncovered regime, the coefficient on $\lambda_j(\theta_j \cdot \mathbf{z}')$ is positive, but not significantly different from zero. The direction of bias under collective contracts depends on the correlation between $\lambda_j(\theta_j \cdot \mathbf{z}')$ and the covariates in eq. (5). Given that $\lambda_j(\theta_j \cdot \mathbf{z}')$ is decreasing in P_j and in all covariates that have a positive impact on P_j, the negative estimates of $-\sigma\rho_j$ suggest that the OLS-coefficients on covariates that are positively correlated with the choice of either centralised or firm-specific contracts should be upward biased. The multinomial logit estimates in Table A6 show that the log-odds ratio of choosing centralised contracts as compared to no-coverage increases significantly with quasi-rents. By contrast, the effect of quasi-rents on the log-odds ratio of choosing firm-specific contracts as compared to no-coverage is found to be insignificant. The resulting marginal effects of quasi-rents on centralised contracts, firm-specific contracts and no-coverage are 0.0002, -0.00003 and -0.0002. Given the estimated coefficients on $\lambda_j(\theta_j \cdot \mathbf{z}')$, we expect that correcting for selectivity should make little difference under no-coverage and firm-specific contracts, and should lead to a decline in the coefficient under centralised contracts. Indeed, the selectivity-corrected rent-sharing coefficient is found to be slightly lower for centralised contracts than the corresponding OLS coefficient. As selectivity plays no major role for the uncovered regime, most of the selectivity-corrected estimates do not substantially differ from the OLS-estimates. If a selectivity-correction changes anything at all, it results in even smaller rent-coeffcients under centralised contracts. In sum, this leads us to conclude that the overall pattern of rent-sharing across the three wage determination regimes appears to be quite robust to the endogeneity of the bargaining regime.

6 Summary and Conclusions

The aim of this paper was twofold: first, we have addressed the question of whether German wages respond to firm-specific profitability conditions and second, we have been interested in whether the sensitivity of wages to firm profits depends on collective bargaining coverage. The institutional discussion has shown that firm-specific contracts and flexibility provisions under centralised contracts provide a means to adjust wages to local conditions at the firm-level and that such adjustments are generally influenced by industry-wide unions which may retain control over centralised union objectives. Provided those flexibility provisions are used, bargaining power considerations lead us to expect wages to react stronger to local conditions in firms that are covered by a collective contract than in uncovered firms. However, there are various reasons for why unions might not want to adjust wages to local firm performance, such as high transaction costs or workers' demand for income insurance. We therefore take our empirical findings as a test of whether flexibility provisions are really exploited or whether unions suppress inter-firm wage dispersion due to heterogeneous firm performance.

Using data from the IAB Establishment-Panel, the results of our empirical analysis offer a remarkably consistent picture: in general, rent-sharing is found to be an empirically relevant phenomenon in Germany. However, the extent of rent-sharing seems to be significantly lower in establishments that are subject to a collective wage agreement - irrespective of whether the agreement is industry- or firm-specific. While POLS-estimates still yield positive estimates of the rent-sharing coefficient in covered establishments, GMM-results accounting for unobserved heterogeneity and the endogeneity of rents point to a rent-sharing coefficient of zero. This finding is indicative of the presence of unobserved factors in covered establishments which are positively correlated with profits and impact positively upon wages. One such factor may be that a compressed intra-firm wage structure under collective wage contracts causes establishments to upgrade the quality of their workforce. This might lead to higher unobserved worker productivity in such establishments and therefore to upward-biased estimates in the simple POLS-specification. Finally, we find the pattern of rent-sharing to be robust to sample selection and the endogeneity of the bargaining regime.

For centralised wage-agreements, the invariance of wages against local profits suggests that the use of flexibility provisions in central wage agreements appears to be empirically negligible. Even though firms may pay wages above the going rate and may make use of opt-out clauses, the potential for adjustments to local profitability conditions appears to be largely unused. A similar result holds for wage determination under firm-specific wage contracts. As such contracts are generally concluded by industry-specific unions, one possible explanation might be that a considerable fraction of firm-specific contracts simply adopts wage bargains negotiated in the corresponding industry agreement. Taken together, our results seem to support the notion that unions favour a compressed intra-industry wage structure and suppress firm-level rent-sharing, either due to workers' demand for income insurance or due to high transaction costs.

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A Appendix

A.1 Selectivity correction

Defining regimes R_1, R_2 and R_3 as wage determination under no-coverage, centralised contracts and firm-specific contracts, respectively, the wage equations for each regime R_j, $j = 1, 2, 3$, become

$$w_j = \alpha_j + \beta_{j\pi} \cdot \pi_j + \gamma_j \cdot \mathbf{x}_j' + \delta_j \cdot \mathbf{s}_j' + u_j \quad \text{if } R_j = 1, \tag{8}$$

where the variance of u_j is given by σ^2 and the indices i, t are suppressed for expositional convenience. Selection into the three regimes is modelled by the following equation

$$R_j^* = \theta_j \cdot \mathbf{z}' + \eta_j \ , \ j = 1, 2, 3, \tag{9}$$

where R_j^* denotes the unobserved utility from choosing bargaining regime R_j, which is determined by a vector of observable covariates $\mathbf{z}'$ with a parameter vector θ_j. η_j denotes an error term which is possibly correlated with u_j. Regime R_j is being chosen when its utility R_j^* exceeds the utility levels of the other regimes, i.e. if $R_j^* > \max\limits_{k \neq j} \{R_k^*\}$, $j = 1, 2, 3$, which is equivalent to

$$\theta_j \cdot \mathbf{z}' > \ \varepsilon_j \text{ with } \varepsilon_j = \max_{k \neq j} \{\theta_k \cdot \mathbf{z}' + \eta_k - \eta_j\} \, , \ j = 1, 2, 3. \tag{10}$$

It is clear from eqs. (8) and (10) that unbiased estimates of all parameters of interest will be obtained only if $E(u_j | \theta_j \cdot \mathbf{z}' > \varepsilon_j) = 0$. To account for $E(u_j | \theta_j \cdot \mathbf{z}' > \varepsilon_j)$, Lee (1983) proposes an extension of the two-step selection-bias correction method developed by Heckman (1979). Assuming that the error term u is normally distributed and selectivity into the regimes can be modelled as a multinomial logit, Lee (1983) shows that

$$E(u_j | \theta_j \cdot \mathbf{z}' > \varepsilon_j \ ) \ = -\sigma \rho_j \lambda_j(\ \theta_j \cdot \mathbf{z}') \tag{11}$$

where

$$\lambda_j(\ \theta_j \cdot \mathbf{z}') \ = \frac{\phi(\Phi^{-1}(P_j))}{P_j}, \ j = 1, 2, 3 \text{ and } P_j = \frac{\exp(\theta_j \cdot \mathbf{z}')}{\sum\limits_k \exp(\theta_k \cdot \mathbf{z}')}, \ j, k = 1, 2, 3, \tag{12}$$

and ρ_j is the correlation-coefficient between u_j and (a transformation of) ε_j. The conditional expectation of wages, given that bargaining regime R_j is being chosen, then becomes eq. (6) in the main text.

A.2 Descriptive Statistics by Bargaining Coverage

	CENT		DECENT		NO-COVERAGE	
Variables	Mean	Std.-Dev.	Mean	Std.-Dev.	Mean	Std.-Dev.
w	54.64	17.66	48.37	16.04	38.78	17.95
π	85.14	102.28	68.40	90.40	37.15	64.13
$\overline{w}$	53.42	10.87	49.63	10.93	46.61	11.97
HIGHSHARE	0.27	0.20	0.26	0.20	0.19	0.19
BLUESHARE	0.39	0.22	0.47	0.23	0.48	0.25
APPSHARE	0.04	0.06	0.04	0.05	0.05	0.07
FEMSHARE	0.26	0.19	0.24	0.19	0.29	0.24
SIZE	865.33	3,123.63	417.65	870.11	75.81	161.27
WCOUNCIL	0.80	0.40	0.75	0.44	0.24	0.43
K/L	204.85	400.02	721.65	3,780.18	150.59	385.72
EAST	0.29	0.45	0.56	0.50	0.68	0.47
OWN	0.20	0.40	0.12	0.32	0.28	0.45
Obs.	2,120		392		899	

Table A2a: Final sample

	CENT		DECENT		NO-COVERAGE	
Variables	Mean	Std.-Dev.	Mean	Std.-Dev.	Mean	Std.-Dev.
w	53.80	19.02	47.65	18.53	38.68	18.24
π	81.95	103.10	71.42	106.92	40.26	74.40
$\overline{w}$	53.74	10.95	50.06	11.29	48.46	12.15
HIGHSHARE	0.26	0.20	0.24	0.19	0.20	0.21
BLUESHARE	0.39	0.22	0.46	0.24	0.45	0.26
APPSHARE	0.05	0.06	0.04	0.06	0.05	0.08
FEMSHARE	0.25	0.20	0.27	0.21	0.31	0.25
SIZE	650.80	2,214.64	366.84	1,019.92	72.09	161.91
WCOUNCIL	0.75	0.43	0.69	0.46	0.21	0.41
K/L	177.55	381.70	405.84	2,537.03	140.07	362.74
EAST	0.27	0.45	0.52	0.50	0.66	0.47
OWN	0.21	0.41	0.18	0.38	0.33	0.47
Obs.	4,751		892		2,974	

Source: IAB Establishment Panel 1995-2002.

Note: All monetary values are measured in 1,000 DM. 1 € corresponds to 1.95583 DM.

Table A2b: Original sample

A.3 Regression Results

Model	*1*	*2*	*3*	*4*
$\overline{w}$	0.403**	0.360*	0.322	0.330
	(0.175)	(0.190)	(0.198)	(0.211)
w(t-1)		0.064	0.066	-0.002
		(0.042)	(0.043)	(0.056)
π	0.077***	0.064***	0.057**	0.047*
	(0.029)	(0.025)	(0.024)	(0.028)
$\pi*$CENT	-0.079**	-0.073**	-0.074***	-0.061**
	(0.034)	(0.030)	(0.028)	(0.030)
$\pi*$DECENT	-0.092**	-0.084**	-0.082**	-0.065*
	(0.042)	(0.041)	(0.040)	(0.038)
π(t-1)			0.023	0.034
			(0.025)	(0.035)
$\pi*$CENT(t-1)			-0.011	-0.028
			(0.021)	(0.028)
$\pi*$DECENT(t-1)			-0.016	-0.018
			(0.022)	(0.035)
π(t-2)				0.043*
				(0.024)
$\pi*$CENT(t-2)				-0.047*
				(0.024)
$\pi*$DECENT(t-2)				-0.036
				(0.028)
$\pi = -\pi*$CENT ($p-$value)	0.908	0.586	0.321	0.429
$\pi = -\pi*$DECENT ($p-$value)	0.586	0.504	0.415	0.481

*Significant at 10%-level.
**Significant at 5%-level.
***Significant at 1%-level.

Table A3: Differenced GMM regression results

... to be continued on next page

... continue Table A3

Model	*1*	*2*	*3*	*4*
HIGHSHARE	-12.835**	-9.038*	-8.310	-11.834
	(5.133)	(5.346)	(5.140)	(6.927)
BLUESHARE	-1.280	2.578	2.379	5.706
	(2.568)	(3.098)	(3.003)	(3.549)
APPSHARE	-1.818	8.088	6.258	11.176
	(11.414)	(13.274)	(13.294)	(16.319)
FEMSHARE	-1.834	-5.925	-5.832	-6.503
	(8.722)	(9.083)	(8.933)	(10.561)
SIZE	-0.001	-3.0e^{-04}	-7.0e^{-04}	0.002
	(0.003)	(0.003)	(0.003)	(0.004)
SIZE^2	-1.92e^{-08}	-1.61e^{-08}	-6.34e^{-09}	-1.30e^{-08}
	(2.98e^{-08})	(2.96e^{-08})	(3.20e^{-08})	(4.55e^{-08})
CENT	3.652**	3.694**	3.863**	3.252*
	(1.809)	(1.723)	(1.672)	(1.665)
DECENT	4.179**	3.981*	3.783*	2.794
	(2.020)	(2.074)	(2.041)	(1.994)
WCOUNCIL	-1.702	-2.243	-2.490	-1.043
	(2.080)	(2.125)	(2.108)	(2.618)
K/L	-0.001***	-0.002***	-0.002***	-0.002***
	(4.0e^{-04})	(3.0e^{-04})	(3.0e^{-04})	(3.0e^{-04})
Sargan/Hansen ($p-$value)	0.288	0.463	0.604	0.443
AR(2) ($p-$value)	0.932	0.460	0.688	0.401
Establishments	661	661	661	661
Observations	2,750	2,089	2,089	1,428

Note: The dependent variable is the aggregate per-capita wagebill. All variables are first-differenced. Results are reported for one-step differenced GMM-estimators. All specifications include time dummies. Heteroscedasticity-robust standard errors are in parentheses.
*Significant at 10%-level.
** Significant at 5%-level.
***Significant at 1%-level.

Table A3: Differenced GMM regression results

Model	*1*	*2*	*3*	*4*
$\overline{w}$	0.386**	0.374**	0.323**	0.367*
	(0.156)	(0.149)	(0.161)	(0.197)
w(t-1)		0.221***	0.228***	0.227***
		(0.046)	(0.046)	(0.051)
π	0.084***	0.072***	0.059**	0.044
	(0.024)	(0.025)	(0.026)	(0.029)
$\pi*$CENT	-0.070**	-0.067**	-0.070**	-0.053*
	(0.029)	(0.030)	(0.028)	(0.030)
$\pi*$DECENT	-0.062**	-0.061**	-0.057**	-0.053
	(0.031)	(0.030)	(0.028)	(0.035)
π(t-1)			0.007	0.006
			(0.016)	(0.019)
$\pi*$CENT(t-1)			0.014	0.011
			(0.013)	(0.016)
$\pi*$DECENT(t-1)			0.006	0.024
			(0.015)	(0.020)
π(t-2)				0.019
				(0.014)
$\pi*$CENT(t-2)				-0.023
				(0.016)
$\pi*$DECENT(t-2)				-0.004
				(0.019)
π = -$\pi*$CENT ($p-$value)	0.267	0.739	0.389	0.535
π = -$\pi*DE$CENT ($p-$value)	0.204	0.483	0.906	0.689

* Significant at 10%-level.
** Significant at 5%-level.
***Significant at 1%-level.

Table A4: SYS-GMM regression results

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... continue Table A4

Model	*1*	*2*	*3*	*4*
HIGHSHARE	-7.306	-5.846	-4.884	-3.073
	(4.756)	(4.976)	(4.867)	(5.935)
BLUESHARE	3.618	6.267*	6.140*	11.751**
	(2.990)	(3.606)	(3.537)	(4.647)
APPSHARE	-10.914	-7.663	-8.394	3.351
	(11.145)	(11.573)	(11.883)	(19.267)
FEMSHARE	-4.603	-5.326	-4.706	-4.775
	(4.929)	(4.720)	(4.688)	(5.291)
SIZE	0.002***	0.002**	0.002**	0.002**
	($7.0e^{-04}$)	($7.0e^{-04}$)	($7.0e^{-04}$)	($8.0e^{-04}$)
SIZE^2	-2.69e^{-08*}	-2.22e^{-08}	-2.28e^{-08}	-2.80e^{-08}
	($1.46e^{-08}$)	$1.44e^{-08}$	($1.49e^{-08}$)	($1.76e^{-08}$)
CENT	4.850***	4.809***	5.279***	3.710*
	(1.861)	(1.856)	(1.818)	(2.066)
DECENT	3.258**	3.613**	3.498**	4.276*
	(1.636)	(1.627)	(1.690)	(2.479)
WCOUNCIL	6.210**	3.882	3.841	4.250
	(2.633)	(2.392)	(2.395)	(2.796)
K/L	$3.67e^{-06}$	-2.0e^{-04}	-1.0e^{-04}	-3.0e^{-04}
	($1.0e^{-04}$)	($2.0e^{-04}$)	($2.0e^{-04}$)	($2.0e^{-04}$)
Sargan/Hansen ($p-$value)	0.312	0.346	0.312	0.318
Diff. Test comp. to Table A2 ($p-$value)	0.484	0.220	0.056	0.184
AR(2) ($p-$value)	0.713	0.128	0.204	0.118
Establishments	661	661	661	661
Observations	2,750	2,750	2,750	2,089

Note: The dependent variable is the aggregate per-capita wagebill. Results are reported for one-step SYS-GMM-estimators. All specifications include time dummies, 15 industry dummies as well as an east-west and an ownership dummy. All endogenous and predetermined variables are assumed to be correlated with the establishment-specific effect.

Heteroscedasticity-robust standard errors are in parentheses.

*Significant at 10%-level.

** Significant at 5%-level.

***Significant at 1%-level.

Table A4: SYS-GMM regression results

	CENT		DECENT		NO-COVERAGE	
	OLS	Selectivity corrected	OLS	Selectivity corrected	OLS	Selectivity corrected
π	0.020***	0.019***	0.024***	0.025***	0.045***	0.045***
	(0.003)	(0.003)	(0.009)	(0.010)	(0.006)	(0.005)
$\overline{w}$	0.547***	0.619***	0.658***	0.671***	0.259***	0.268***
	(0.063)	(0.072)	(0.157)	(0.157)	(0.066)	(0.083)
HIGHSHARE	17.996***	15.715***	14.622***	16.117***	14.544***	14.284***
	(1.923)	(1.953)	(4.487)	(4.442)	(2.064)	(2.482)
BLUESHARE	4.944***	4.485***	5.259**	3.696	6.433***	6.306***
	(1.174)	(1.247)	(2.183)	(2.524)	(1.076)	(1.261)
APPSHARE	1.562	-1.451	8.141	11.216	-6.885*	-7.050**
	(4.007)	(4.541)	(9.067)	(11.931)	(3.638)	(3.447)
FEMSHARE	-19.017***	-17.510***	-14.397***	-15.159***	-12.565***	-12.291***
	(1.327)	(1.447)	(2.924)	(2.995)	(1.184)	(1.670)
SIZE	0.002***	0.001***	0.001	0.002	0.023***	0.022***
	($3.00e^{-04}$)	($3.00e^{-04}$)	(0.001)	(0.002)	(0.004)	(0.005)
SIZE^2	-2.60e^{-08}**	-1.92e^{-08}	-4.10e^{-08}	-6.90e^{-08}	-8.81e^{-06}***	-8.79e^{-06}***
	($1.07e^{-08}$)	($1.29e^{-08}$)	($8.74e^{-08}$)	($2.24e^{-07}$)	($2.06e^{-06}$)	($2.38e^{-06}$)
WCOUNCIL	8.491***	5.455***	7.470***	4.141**	3.256***	2.763
	(0.595)	(1.075)	(1.092)	(2.017)	(0.672)	(2.344)
K/L	0.002***	0.002***	$6.28e^{-05}$	-0.0004**	$2.00e^{-04}$	$2.00e^{-04}$
	($7.00e^{-04}$)	($7.00e^{-04}$)	($1.10e^{-04}$)	($3.00e^{-04}$)	($8.00e^{-04}$)	($8.00e^{-04}$)
OWN	-4.442***	-4.825***	-4.720***	-4.064**	-6.524***	-6.598***
	(0.556)	(0.622)	(1.334)	(1.345)	(0.511)	(0.698)
EAST	-3.194***	0.206	-3.588	-5.487**	-8.713***	-8.308***
	(1.009)	(1.431)	(2.240)	(2.253)	(1.020)	(2.165)
Intercept	8.036***	8.752*	5.050	21.610*	24.197***	22.453**
	(3.746)	(4.784)	(9.511)	(11.367)	(4.844)	(10.394)
$\lambda_j(\theta_j \cdot z')$		-6.124***		-9.569***		0.781
		(1.814)		(3.388)		(3.531)
Ind.-/Time	Yes	Yes	Yes	Yes	Yes	Yes
Observations	4,725	4,725	892	892	2,970	2,970

Note: The dependent variable is the aggregate per-capita wagebill. Heteroscedasticity-robust (POLS) and bootstrapped standard errors (selectivity corrected results - 100 repetitions) are in parentheses. All models include time dummies and 15 industry dummies.
*Significant at 10%-level ** Significant at 5%-level ***Significant at 1%-level

Table A5: Selectivity corrected regression results

	$\log\frac{P(\text{DECENT})}{P(\text{NO COVERAGE})}$	$\log\frac{P(\text{CENT})}{P(\text{NO COVERAGE})}$
π	$5.82e^{-04}$	0.001^{***}
	$(5.06e^{-04})$	$(3.79e^{-04})$
$\overline{w}$	-0.033^{***}	-0.043^{***}
	(0.011)	(0.007)
HIGHSHARE	0.438	1.189^{***}
	(0.345)	(0.240)
BLUESHARE	0.594^{***}	0.431^{***}
	(0.220)	(0.155)
APPSHARE	-0.608	0.910^{*}
	(0.845)	(0.534)
FEMSHARE	-0.696^{***}	-1.424^{***}
	(0.249)	(0.179)
SIZE	0.001^{***}	0.002^{***}
	$(1.80e^{-04})$	$(2.00e^{-04})$
WCOUNCIL	1.621^{***}	1.587^{***}
	(0.108)	(0.759)
K/L	$1.12e^{-04}$	$-8.92e^{-06}$
	$(7.18e^{-05})$	$(7.10e^{-05})$
OWN	0.009	0.266^{***}
	(0.112)	(0.073)
EAST	-0.719^{***}	-1.714^{***}
	(0.181)	(0.122)
Publicly listed company	0.529^{**}	0.637^{***}
	(0.243)	(0.211)
Founded in 1990 or later	0.076	-0.531^{***}
	(0.097)	(0.071)
Observations	8,587	
Pseudo R^2	0.254	

Note: The dependent variable is the bargaining regime.
The specification includes time dummies and 15 industry dummies.
*Significant at 10%-level
** Significant at 5%-level
***Significant at 1%-level

Table A6: Multinomial logit estimates

Rent-Sharing and Collective Wage Contracts - Evidence from German Establishment-Level Data

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Abstract

Using German establishment-level data, this paper analyses whether wages respond to firm-specific profitability conditions. Particular emphasis is given to the question of whether the extent of rent-sharing varies with collective bargaining coverage. In this context, two conflicting hypotheses are tested. The first one asserts that unions exploit their bargaining power at the firm-level and appropriate a larger share of rents than the bargaining parties in uncovered firms. The second one states that unions favour a compressed intra-industry wage structure and suppress the responsiveness of wages to firm-specific profitability conditions. The empirical analysis provides strong support for the second hypothesis. While pooled OLS estimates yield positive estimates of the rent-sharing coefficient in covered establishments, dynamic panel data estimates accounting for unobserved heterogeneity and the endogeneity of rents point to a rent-sharing coefficient of zero.

Keywords: Rent-Sharing, Wage-Setting Structure, Unions, Panel Data
JEL Code: C23, J31, J51

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1 Introduction

The question of whether wages vary systematically with firms' ability-to-pay has long been of considerable interest to labour economists. In the literature, various theoretical explanations have been advanced for a positive relationship between wages and profits (see e.g. Blanchflower et al. 1996, Hildreth and Oswald 1997). Apart from temporary frictions and efficiency wages, a frequently invoked explanation refers to union power. Under collective bargaining, workers' remuneration may be expected to increase with profits, as unions will be able to appropriate part of the industry or firm-specific rents. Whether wages react to industry or firm-specific conditions should naturally depend on the level of bargaining. Intuitively, wages ought to be most responsive to firm-specific profitability conditions if wage determination allows for some adjustment to local conditions at the firm level.

Although the bargaining structure appears to be an important determinant for the degree of rent-sharing at the firm or industry level, there is surprisingly little empirical evidence on this topic. While the question of whether wages vary systematically with profits has spawned a vast empirical literature (see e.g. Abowd and Lemieux 1993, van Reenen 1996, Arai 2003, Budd et al. 2005)[1], few studies explicitly address the role of the bargaining structure for rent-sharing. One exception is the study by Holmlund and Zetterberg (1991), which draws on a cross-country comparison to analyse this question. The authors find that countries with highly centralised and coordinated bargaining institutions exhibit less industry level rent-sharing than countries with relatively decentralised bargaining systems. In this paper, we draw on establishment-level data from Germany and present new evidence on rent-sharing and collective bargaining by exploiting intra-national variations in the bargaining structure. Clearly, such variations offer the advantage of controlling for a large part of the unobserved heterogeneity in institutional conditions characterising cross-country comparisons.

The German institutional environment provides a useful example for the coexistence of different bargaining structures. Until the early 1990s, wage determination was dominated by centralised wage bargaining between industry-specific unions and employers' associ-

[1]Further studies include Christofides and Oswald (1992), Blanchflower et al. (1996), Hildreth and Oswald (1997), Abowd et al. (1999), Margolis and Salvanes (2001), Kramarz (2003), Dobbelaere (2004) and Martins (2006) amongst others. There are only few previous studies on the relationship between wages and profits in Germany: Hübler and König (1998) and Klodt (2000) use data from the 'Hanover establishment panel' and report a significant positive impact of profits on average firm wages.

ations. Industry agreements were embedded in a corporatist environment characterised by a high degree of coordination (Soskice 1990). However, in the last decade, there has been a strong tendency towards decentralisation of wage determination, as wage determination without any bargaining coverage has increasingly grown in importance (Hassel 1999, Ochel 2005). Even within centralised industry agreements, there have been numerous attempts to allow for more (downward) flexibility of wages by introducing opt-out and hardship clauses. Moreover, since bargained wages in centralised agreements merely represent a lower bound for wages, there is also sufficient room for upward flexibility.

Given this intra-national variation in German wage determination regimes, the principal aim of this paper is to shed light on the following questions: Do firm-specific contracts and flexibility provisions in centralised industry agreements allow for rent-sharing at the firm level? If so, does the extent of rent-sharing differ from that in firms without any bargaining coverage? A striking feature of the German wage determination process is that decentralisation in collective wage determination merely refers to the *level* of bargaining and not to the *degree of coordination*. The reason is that - as will be discussed below - collective wage determination at the firm level is generally influenced by industry-wide unions which may retain control over centralised union objectives. Bargaining power considerations lead us to expect the extent of rent-sharing to be larger the more coordinated the wage-setting process and the more decentralised the level of wage determination. The first hypothesis to be tested, therefore, is that unions are able to skim off an even larger part of rents than the bargaining parties in uncovered firms. The question of whether this notion may be confirmed empirically is of considerable interest in an institutional environment such as the German one, which has long been regarded as corporatist and has been thought of to offer little scope for excessive rent-sharing at the firm level. On the other hand, there are various reasons why unions might favour a compressed intra-industry wage structure, such as high transaction costs or workers' demand for income insurance (see Agell and Lommerud 1992, Burda 1995, Agell 2002). A countervailing hypothesis to be tested, therefore, is that unions suppress any inter-firm wage dispersion due to heterogeneous firm performance.

We investigate the relationship between wages and profitability using the *IAB Establishment Panel*. This data set is particularly useful for our purposes as it provides detailed information on whether an establishment is subject to an industry-wide wage agreement, a firm-specific wage agreement or to no wage agreement at all. In our estimation strategy, we first focus on simple static pooled Ordinary Least Squares (POLS) estimates. The OLS estimations serve as a benchmark case and will be modified by using dynamic panel data methods. First, we address the possibility of unobserved firm-specific time invariant factors. A second problem concerns the endogeneity of our profitability measure, since wages and profits are simultaneously determined. Third, we consider dynamic specifications to allow for possible dynamics in the response of wages to profitability conditions. Finally, we investigate whether our results are robust to sample selection and the endogeneity of the bargaining structure.

The remainder of the paper is organised as follows: the institutional background of German wage determination is presented in Section 2. Section 3 provides a theoretical discussion to derive testable hypotheses about the extent of rent-sharing under the different wage-setting regimes. These hypotheses are tested in Section 4. While Section 4.1. presents the general empirical model, Section 4.2. describes the data set and the main variables used in the empirical analysis. Section 4.3. reports the estimation results. Section 5 discusses the robustness of the results. Finally, Section 6 provides a discussion and some conclusions.

2 Institutional Background

In Germany, basically three forms of wage determination may be distinguished: central collective wage agreements, firm-specific collective wage agreements as well as wage determination without any collective bargaining coverage. Until the early 1990s, wages determination was dominated by central regional and industry-wide collective wage agreements (*Flächentarifverträge*). Such central wage agreements are negotiated between an industry-specific trade union and an employers' association. They are legally binding on all member firms of the respective employers' association and on all employees who are members of the trade union. Although, strictly speaking, the negotiated wage only applies to union members, member firms generally extend the wage settlement to non-member employees as well.[2]

Since the early 1990s, centralised collective wage agreements have substantially declined in importance in the German system of wage determination (see e.g. Hassel 1999,

[2]The reason is that non-unionised employees who would receive a lower wage may be expected to join the union anyway in order to benefit from the higher union wage. In addition, central wage agreements may also apply to non-member firms and their employees if the agreement is declared to be generally binding by the Federal Ministry of Labour.

Kohaut and Schnabel 2003). The tendency towards more decentralisation is the result of three major developments. First, the number of firm-specific collective wage agreements has grown in absolute terms.[3] These agreements are negotiated between an individual firm and a union. A noteworthy feature of these agreements is that they are concluded by industry-specific unions and do not involve uncoordinated wage bargaining of independent firm-specific unions. That is, decentralisation here merely refers to the *level* of bargaining and not to the *degree of coordination*. Second, wages are now more often determined without any bargaining coverage at all. In firms that are not covered by a collective agreement wages are either determined in individual wage contracts or in plant-specific agreements (*Betriebsvereinbarungen*) between works councils and the management.[4] In contrast to firm-specific collective wage agreements, this kind of wage determination can be characterised as decentralised and uncoordinated. Third, there is a tendency even within centralised wage agreements to allow for more flexibility at the firm level. In recent years, contractual opt-out clauses or hardship clauses have become a widespread element of central agreements. While opt-out clauses delegate 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. In general, the adoption of such clauses requires the approval of the collective bargaining parties (Hassel 1999, Ochel 2005). Moreover, bargained wages in centralised agreements merely represent a lower bound for wages, so that there is always sufficient room for upward flexibility. Even though the wage drift is part of the local negotiations between works councils and the firm, it is also likely to be coordinated by the centralised bargaining parties. The reason is that union density among works councils members is very high (Hassel 1999), and this is particularly relevant for covered firms. Thus, similar to firm-specific collective contracts, the adoption of flexibility provisions in centralised wage agreements is still coordinated by the centralised bargaining parties and involves merely a decentralisation of the level of bargaining.

[3]According to the German Federal Ministry of Labour the number of firms with a firm-specific wage agreement increased economy-wide from about 2,500 in 1990 to 8,000 in 2004. However, evidence from the *IAB Establishment Panel* indicates that, in relative terms, firm-specific contracts have again declined in importance since the mid 1990s: Between 1996 and 2004 the fraction of establishments with a firm-specific contract decreased from 10 to 2 per cent in western Germany, while the decline was from 15 to 4 per cent in eastern Germany.

[4]According to the German Works Constitution Act, works councils are not allowed to negotiate about issues that are normally dealt with in collective agreements, even in firms that are not party to collective agreements. In practice, however, works councils may be expected to play a crucial role in wage determination (see e.g. Hassel 1999, Hübler and Jirjahn 2003).

3 Theoretical Considerations

The purpose of the present section is to derive testable hypotheses about the degree of rent-sharing under the different wage determination regimes. The institutional discussion in Section 2 has yielded two important insights. First, collective contracts are by no means an obstacle to the adjustment of wages to local conditions at the firm level, since recent decentralisation tendencies in Germany have introduced - at least formally - the option of making such adjustments. Second, even if wages are collectively determined at the firm level, they are still influenced by industry-wide unions which may retain control over centralised union objectives.

Thus far, the theoretical literature has mainly focused on the effects of different bargaining regimes on overall wage levels (see e.g. Calmfors and Driffill 1988, Soskice 1990, Dowrick 1993). There is little theory to guide us on the expected effects on the returns to firm-specific attributes such as profits. The rent-sharing literature generally predicts a pay-performance link that depends on the relative bargaining strength of the bargaining parties (e.g. Abowd and Lemieux 1993, Blanchflower et al. 1996, van Reenen 1996). Such considerations lead us to expect the sensitivity of wages to firm-specific profits to be larger under firm-specific contracts than in uncovered firms. An important argument is that firm-specific contracts in Germany are concluded by industry-specific unions, whose bargaining power is likely to be considerably greater than that of works councils determining wages in uncovered firms. This argument is reinforced by the fact that the wage bargaining process under firm-specific contracts is highly coordinated by an industry-wide union, whereas it is completely uncoordinated in uncovered firms. The bargaining parties in uncovered firms therefore have an incentive to cut wages in order to gain a larger share of industry demand, and this restricts their ability to raise wages in response to more favourable profitability conditions. With an industry union, this competitive mechanism completely disappears, since a central union may coordinate wage determination at the industry level.[5] For this reason, one might expect an industry union to capture a larger share of rents under firm-specific contracts than works councils or individuals in uncovered firms.

The extent of rent-sharing under centralised contracts ultimately depends on whether the bargaining parties make use of flexibility provisions. If such provisions are exploited, the extent of rent-sharing should be larger under industry-contracts than in uncovered

[5]We have formalised this argument elsewhere (Guertzgen 2005).

firms. The argument here is similar to the reasoning for firm-specific contracts, since the institutional discussion has shown that any adjustment to local conditions at the firm level is still highly coordinated by the centralised bargaining parties. At this point, it is worth noting that the question of whether flexibility provisions are used to adjust wages to local profitability conditions still remains to be answered empirically. For example, even though contractual opt-out and hardship clauses have become an important (formal) element of centralised agreements, empirical evidence on the use of such clauses is rather scarce.

Note, in this context, that there are various reasons why unions (and possibly employers) might favour a compressed intra-industry wage structure. First, the transaction costs incurred when adjusting wages to firm-level profitability may be high and outweigh any gain involved with wage differentiation. Second, a further rationale for unions to maintain a compressed wage structure might be workers' demand for income insurance (see Agell and Lommerud 1992, Burda 1995, Agell 2002). In our context, intra-industry wage compression provides insurance against two dimensions of uncertainties. First, wage compression between firms at a given point in time may reduce income risk if workers face uncertainties over the allocation to more or less profitable firms. Second, with a compressed intra-industry wage structure wage growth is likely to depend on average sector performance, so that workers' wages at a given employer should also be sheltered against fluctuations in firm-level profitability over time.[6] Thus, the countervailing hypothesis to be tested is that unions suppress inter-firm wage dispersion due to heterogeneous firm performance.

4 Empirical Analysis

4.1 Empirical Model and Testable Hypotheses

In order to quantify the relationship between firm-specific profitability and wages across different wage-setting regimes, we impose a wage equation taking the basic form

$$w_{it} = \alpha + (\beta_0 + \beta_{\pi\_CENT} \cdot CENT_{it} + \beta_{\pi\_DECENT} \cdot DECENT_{it}) \cdot \pi_{it} + \gamma \cdot \mathbf{x}'_{it} + \delta \cdot \mathbf{s}'_{it} + f_i + u_{it}. \quad (1)$$

Since we will use establishment level panel data, all variables are subscripted by a establishment-index i and a time index t. The dependent variable, w, is the establishment-

[6] The issue of wage insurance at the firm level has been taken up recently by Guiso et al. (2005) and Cardoso and Portela (2005). However, these empirical studies do not distinguish different bargaining regimes.

specific average wage per worker. The explanatory variable of main interest is π, measuring establishment-specific per-capita profitability.[7] Following most of the rent-sharing literature (see e.g. see Abowd and Lemieux 1993, van Reenen 1996), profitability, π, is measured by per-capita quasi-rents. We choose quasi-rents - defined as value-added minus the opportunity cost of labour - for two reasons. First, from a theoretical perspective quasi-rents may be interpreted as representing the 'pie' to be divided between the bargaining parties. Second, from an econometric perspective, the use of quasi-rents instead of profits enables us to circumvent the endogeneity problem induced by the accounting relationship between wages and profits.

In eq. (1), $\mathbf{x}'$ represents a (column) vector of further establishment characteristics with a coefficient vector γ, while $\mathbf{s}'$ denotes a vector of industry characteristics with a coefficient vector δ. For $\mathbf{s}'$ we include the average sectoral wage as well as industry dummies. The latter are intended to capture industry-specific factors, such as the overall level of industry demand and the degree of competition. The vector of establishment-specific characteristics, $\mathbf{x}'$, includes among other variables dummies for the three wage-setting regimes since the bargaining regime is likely to affect not only the extent of rent-sharing but also the overall wage level. Moreover, $\mathbf{x}'$ contains shares of different skill groups and shares of female workers to control for establishment-specific compositions of the workforce. To account for unobserved differences in worker quality and differences in technologies, further explanatory variables include firm size and the capital-labour ratio. Establishment-specific fixed effects f_i are added to eq. (1) in order to capture unobserved time-invariant factors. Finally, time dummies are included to capture common macroeconomic shocks, and u_{it} is a serially uncorrelated white-noise error term.

Since the emphasis of our analysis is on the impact of different wage-setting regimes on the sensitivity of wages to local profitability conditions, we specify interaction terms $CENT_{it} \cdot \pi_{it}$ as well as $DECENT_{it} \cdot \pi_{it}$,where $CENT$ is a dummy taking the value of unity if an establishment adopts a centralised collective wage agreement and $DECENT$ takes on the value of unity if a firm is party to a firm-specific collective wage contract. Recall that according to our first hypothesis, β_{π_DECENT} and β_{π_CENT} should be positive, if firm-specific contracts and flexibility provisions are used to adjust wages to local firm performance. Conversely, testing $\beta_{\pi\_CENT} = -\beta_0$ (and $\beta_{\pi\_DECENT} = -\beta_0$) provides a

[7]Particularly in case of multi-plant firms, one might object that firm-level profitability provides a more appropriate measure than establishment-level profitability. However, since we only have access to the establishment-level measures, these are taken as a proxy for firm-level profitability.

direct test of the second hypothesis, according to which unions enforce a compressed intra-industry wage structure.

4.2 Data and Variable Description

The empirical analysis uses data from the *IAB-Establishment Panel*. This data set is based on an annual survey of establishments in western Germany administered since 1993. Establishments in eastern Germany entered the panel in 1996. The data base is a representative sample of German establishments employing at least one employee paying social security contributions. The survey data provide information on establishment structure and performance, such as for example the aggregate wage bill, sales, size and composition of the workforce (see e.g. Bellmann et al. 2002). Moreover, the data contain information on whether an establishment is covered by an industry-wide collective wage agreement, a firm-specific wage agreement or by no collective agreement at all.

In our analysis we use data for the years 1995 to 2002, since detailed information on bargaining coverage is available only from 1995 onwards. Because information on a number of variables, such as sales and the share of materials in total sales are gathered retrospectively for the preceding year, we lose information on the last year. Moreover, we restrict our sample to mining and manufacturing establishments with at least two employees. We focus on these sectors, since the introduction of opt-out and hardship clauses has been particularly relevant in central collective wage agreements in these industries. These sectors therefore provide a particularly interesting case for testing the empirical relevance of the use of such clauses. As we apply dynamic panel data methods, only establishments with consistent information on the variables of interest and at least four consecutive time series observations are included in our sample. This results in a sample of 661 establishments with 3,411 observations, yielding an unbalanced panel containing establishment observations with, on average, 5.16 years of data.[8]

The variables used in the subsequent empirical analysis are defined as follows. The dependent variable, w, is defined as the annual aggregate wage bill divided by the number of employees. The number of employees and the wage bill are reported for the month

[8]Originally, the sample included 3,546 establishments with consistent information on all the variables of interest. 21 observations were dropped due to suspected errors in the establishment size variable. These observations featured per-capita values of rents of above DM 1 million. For the same reason, 81 observations with a per-capita wage bill of less than DM 8,000 were discarded from the sample. This results in a sample of 3,515 establishments with a total of 8,617 observations. Only 661 establishments feature at least four consecutive time-series observations.

June, where the wage bill is defined exclusive of employers' mandatory social security contributions as well as fringe benefits. Per capita quasi-rents are constructed as the difference between annual sales, material costs and the alternative annual wage bill divided by establishment size, so that

$$\pi = \frac{SALES - MATERIALCOST - \overline{w} \cdot SIZE}{SIZE}. \tag{2}$$

Establishment size ($SIZE$) is calculated as the reported number of employees averaged over the present and preceding year. The alternative wage bill, $\overline{w} \cdot SIZE$, is defined as the annual wage bill which each establishment would incur if it had to pay the average industrial wage. Thus, we approximate $\overline{w}$ by the weighted average of industry-specific annual wages (separately for eastern and western Germany) for blue- and white-collar workers with the weights being the establishment-specific shares of those worker groups in the total work force. All monetary values are expressed as real values by deflating them with a sector-specific producer price index normalised to 1 in 2000. Industry-specific price indices and wages are obtained from the Federal Statistical Office Germany and are matched to the establishment data on the basis of a two-digit sector classification.

Further variables include the share of high-skilled workers (defined as skilled white-collar workers), the share of skilled blue-collar workers, the share of female workers and the share of apprentices in the total work force. Because we do not directly observe the capital stock, we need to construct a proxy. We measure capital by using the perpetual inventory method starting from the capital value in the first observation year and using the information on expansion investment in the following years. 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, δ.[9] Capital-stocks in subsequent periods are calculated by adding real expansion investment expenditures.[10] To obtain real values, nominal investment expenditures are deflated by the producer price index of investment goods of the Federal Statistical Office Germany. The capital-labour ratio, K/L, is constructed by dividing the resulting capital proxy by establishment size. An ownership dummy variable indicates whether the establishment is

[9] This involves the assumption that investment expenditures on capital have grown at a constant average rate, g, so that the capital stock in the base year is $K_1 = I_0 + (1-\delta)I_{-1} + (1-\delta)^2 I_{-2} + ... = I_1 \sum_{s=0}^{\infty} [\frac{1-\delta}{1+g}]^s = I_1/(\delta + g)$. In particular, to calculate K_1, we set $\delta = 0.1$ and $g = 0.05$ (see Hempell 2005).

[10] More specifically, $K_t = K_{t-1}(1-\delta) + I_{t-1} = K_{t-1} + EI_{t-1}$,where K_t is the capital stock at the beginning of period t, i.e. at the end of period $t-1$, and EI_t are expansion investment expenditures in period t.

part of a company owned by persons with unlimited liabilities. Table A1 in the appendix contains a summary of all establishment covariates.

Table 1 presents sample statistics for the main variables used in the subsequent analysis. The figures disclose that quasi-rents vary considerably more than average wages. With respect to collective bargaining coverage, the fraction of observations covered by an industry-wide wage agreement amounts to about 62 per cent, while the fraction of observations with a firm-specific agreement is only 11 per cent. 27 per cent of all observations are subject to no agreement at all. Breaking down the sample into those establishments adopting an industry-wide agreement, a firm-specific agreement and into those without any bargaining coverage reveals that average wages are highest under industry-wide agreements and lowest without bargaining coverage (see Table A2a in the Appendix). The variability in wages is higher in uncovered establishments with a coefficient of variation of 0.46 as compared to 0.32 and 0.33 in covered establishments. Moreover, establishments under centralised agreements outperform those under firm-specific contracts and those without bargaining coverage in terms of per-capita quasi-rents. Establishments adopting industry-wide agreements also have more employees and exhibit the largest fraction of high-skilled workers, while establishments without bargaining coverage employ on average more women than those covered by a collective wage agreement. Finally, establishments with firm-specific contracts feature the largest capital-labour ratio.

4.3 Results

4.3.1 Estimation Strategy

We first focus on a simple static pooled Ordinary Least Squares (POLS) specification of eq. (1). The POLS estimations serve as a benchmark case and will be modified in various respects: first, we address the possibility of unobserved establishment-specific time invariant factors. In our context, the presence of unobserved heterogeneity may result from neglected capital costs in the rent measure as well as from differences in technological conditions[11] and worker quality that are not captured by our control variables. As such unobserved factors are likely to be correlated with our profitability measure, simple POLS estimates may be expected to yield biased estimates of β_π. A second problem concerns

[11]With respect to differences in technologies, establishment-specific fixed effects capture e.g. production processes that provide firms with higher rents and which may require compensating wage differentials (e.g. processes involving dangerous work). Such differences might lead to a positive wage-rent correlation which would not be due to rent-sharing (see e.g. Margolis and Salvanes 2001).

Variable	*Definition*	*Mean*	*Std.-Dev.*	*Obs.*
w	Per-capita wage bill	49.74	18.84	3,411
π	Per-capita quasi-rents	70.57	94.59	3,411
$\overline{w}$	Alternative wage	51.19	11.57	3,411
HIGHSHARE	Share of skilled white-collar workers	0.25	0.20	3,411
BLUESHARE	Share of skilled blue-collar workers	0.42	0.23	3,411
APPSHARE	Share of apprentices	0.05	0.06	3,411
FEMSHARE	Share of female workers	0.27	0.21	3,411
SIZE	Establishment size	605.80	2505.35	3,411
CENT	Centralised collective agreement	0.62	0.49	3,411
DECENT	Firm-specific collective agreement	0.11	0.32	3,411
WCOUNCIL	Works council	0.64	0.48	3,411
K/L	Capital-labour ratio	249.94	1344.08	3,411
EAST	Eastern Germany	0.42	0.49	3,411
OWN	Private ownership	0.21	0.41	3,411

Source: IAB-Establishment Panel 1995-2002. Entries are unweighted.
Note: All monetary values are measured in DM 1,000 whereby 1 € corresponds to DM 1.95583.

Table 1: Descriptive statistics

the endogeneity of per-capita rents. A first source of bias is a standard simultaneity bias which occurs if wages, output and quasi-rents are jointly determined. In general, the direction of bias can go either way and largely depends on the underlying relationship between output and employment (see Abowd and Lemieux 1993). In addition, because alternative wages and establishment wages are likely to be positively correlated, there will always be some source of downward bias. Third, we consider more dynamic specifications and include lagged wages and rents as explaining variables in our wage regression. The inclusion of lagged rent measures and lagged wages is meant to allow for possible dynamics in the reaction of wages to profitability conditions and sluggish adjustment of wages.

4.3.2 Pooled OLS-Results

Table 2 reports results from POLS estimations of the impact of quasi-rents per worker on wages. The variables are specified in levels rather than logs, since the use of logs would have required discarding all observations with negative quasi-rents. The estimate of quasi-rents per employee on the average wage is 0.042 when including only the alternative wage in the regression. Adding worker characteristics reduces the coefficient to 0.036, suggesting that around 14 per cent of the correlation between rents and wages is due to systematic sorting of workers across establishments (Model (2)). In particular, high-qualified workers appear to be associated with more profitable establishments. The effects of rents on wages are further reduced if other establishment characteristics, such as establishment size, bargaining coverage, the existence of a works council and ownership status are included (Model (3)). Apart from $APPSHARE$ (fraction of apprentices), the capital-labour ratio K/L and $DECENT$, all control variables enter the regression with their expected sign and are all significant at conventional levels. Establishment size is found to have a significant positive effect on average wages, a result which is consistent with earlier evidence.[12] In the literature, various explanations have been advanced for a positive relationship between firm size and wages, such as differences in profits, capital equipment, worker quality and monitoring costs among others (e.g. Oi and Idson 1999). As we control explicitly for differences in the work force composition, the capital-labour ratio and quasi-rents, the establishment size variable may be interpreted as capturing some part of unobserved worker quality and technology differences.

In Model (4), the inclusion of industry and time dummies leaves the coefficient on rents largely unchanged. Adding industry and time dummies changes the coefficient on the capital-labour ratio to its expected sign, indicating some systematic differences in capital-intensities across industries. Including an east-west dummy does not change the coefficient on rents either (Model (5)). As far as the bargaining coverage effects are concerned, the coefficients on centralised contracts are always significantly positive, whereas decentralised contracts seem to have no significant impact on wages. In addition to the collective bargaining regime, we control for the existence of a works council, which, in line with earlier studies (see e.g. Addison et al. 2001, Hübler and Jirjahn 2003), are found to exert a positive impact on averages wages.

Finally, our main interest concerns the question whether the rent-coefficient differs systematically across the three wage-setting structures. Model (6) includes interactions between collective bargaining coverage and quasi-rents. The inclusion of interactions leads to a larger and more precise estimate of the coefficient on the dummy for firm-level agreements ($DECENT$). In sum, the results indicate that the extent to which wages react to local profitability conditions is significantly lower in establishments that are covered by a collective wage agreement. Even in establishments covered by a firm-specific contract wages appear to be less sensitive to rents. Moreover, the adoption of a centralised wage agreement seems to reduce the magnitude of rent-sharing to a slightly larger extent, as a Wald-Test of $\beta_{\pi\_CENT} = \beta_{\pi\_DECENT}$ can be rejected at the 10 per cent

[12]For German evidence on employer size effects see e.g. Schmidt and Zimmermann (1991) and Gerlach and Hübler (1998).

Model	*1*	*2*	*3*	*4*	*5*	*6*
π	0.042***	0.036***	0.024***	0.025***	0.025***	0.061***
	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.010)
$\pi *$ CENT						-0.042***
						(0.010)
$\pi *$ DECENT						-0.030***
						(0.012)
$\overline{w}$	0.992***	0.815***	0.718***	0.797***	0.650***	0.654***
	(0.024)	(0.031)	(0.031)	(0.038)	(0.065)	(0.064)
HIGHSHARE		11.484***	7.420***	6.727***	10.179***	9.952***
		(1.721)	(1.572)	(1.644)	(1.953)	(1.923)
BLUESHARE		1.433	2.206**	4.308***	4.862***	4.830***
		(1.207)	(1.098)	(1.117)	(1.146)	(1.148)
APPSHARE		-3.183	3.392	7.628**	1.841	1.980
		(3.626)	(3.215)	(3.525)	(3.922)	(3.901)
FEMSHARE		-17.374***	-14.990***	-15.339***	-15.484***	-15.370***
		(1.225)	(1.107)	(1.250)	(1.245)	(1.233)
SIZE			0.001***	0.002***	0.002***	0.002***
			(0.0003)	(0.0003)	(0.0003)	(0.0003)
SIZE^2			-2.17e^{-08}*	-3.03e^{-08}***	-2.94e^{-08}**	-2.91e^{-08}**
			(1.19e^{-08})	(1.16e^{-08})	(1.17e^{-08})	(1.18e^{-08})
CENT			3.772***	4.045***	3.692***	5.619***
			(0.613)	(0.628)	(0.635)	(0.700)
DECENT			0.671	1.195	1.046	2.058**
			(0.754)	(0.745)	(0.744)	(0.886)
WCOUNCIL			7.149***	7.401***	7.116***	7.033***
			(0.632)	(0.631)	(0.632)	(0.627)
K/L			-6.57e^{-06}	0.0002*	0.0002	0.0002
			(-9.55e^{-05})	(0.0001)	(0.0001)	(0.0001)
OWN			-4.108***	-4.267***	-4.225***	-4.189***
			(0.535)	(0.539)	(0.534)	(0.534)
EAST					-2.849***	-2.803***
					(0.995)	(0.988)
Ind.-/Time	No	No	No	Yes	Yes	Yes
Adj. R^2	0.482	0.522	0.597	0.610	0.610	0.615
Observations	3,411	3,411	3,411	3,411	3,411	3,411
Establishments	661	661	661	661	661	661

Note: The dependent variable is the aggregate per-capita wage bill. Heteroscedasticity-robust standard errors are in parentheses. Models (4) - (6) include time dummies and 15 industry dummies.
*Significant at 10%-level ** Significant at 5%-level ***Significant at 1%-level.

Table 2: Pooled OLS regression results

level (with a $p-$value of 0.078). However, the null hypotheses of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$ are also rejected (with $p-$values close to zero), suggesting that the overall impact of rents on wages is still positive under both regimes.

4.3.3 Dynamic Specifications

This section addresses potential econometric problems, such as the possibility of unobserved establishment-specific time invariant factors as well as the endogeneity of rents. A further possible endogenous regressor is firm size, as higher wages are likely to induce firms to reduce their labour force. To allow for sluggish adjustment of wages and time lags in the response of wages to profitability conditions, we add lagged wages and quasi-rents as explanatory variables to our regression. The wage equation then takes the following form

$$w_{it} = \alpha + \beta_w w_{it-1} + \sum_{k=0}^{K} \beta_{\pi t-k} \cdot \pi_{it-k} + \gamma \cdot \mathbf{x}'_{it} + \delta \cdot \mathbf{s}'_{it} + f_i + u_{it}, \tag{3}$$

where the coefficients $\beta_{\pi t-k}$ are specified as in eq. (1). First differencing eq. (3) eliminates time-invariant establishment-specific effects.[13] In eq. (3), first differencing causes the lagged dependent variable Δw_{it-1} to become correlated with the error term Δu_{it}, so that it is necessary to instrument lagged wages. In the absence of second-order correlation in the error term, w_{it-2} and earlier lags provide suitable instruments, since they do not correlate with Δu_{it}. Because rents, their interactions with the wage-setting regimes and establishment size are likely to be endogenous, they are to be instrumented as well. As with the lagged dependent variable, suitable candidates are lagged rents and establishment size in $t-2$ and earlier provided they do not enter eq. (3) as explanatory variables. Since this might be particularly relevant for lagged rents, we test for the significance of rents up to $t-2$.

To estimate eq. (3), we first apply the differenced Generalized 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

[13]First-differenced estimates of specification (6) in Table 2 yield rent-coefficients of 0.025, -0.034 and -0.018 with standard errors of 0.013, 0.015 and 0.014 (for no-coverage, interactions with centralised contracts and firm-specific contracts, respectively), suggesting a considerable upward bias of the POLS estimates. For the sake of expositional brevity, we do not report the full first-differenced specifications. The estimates are available on request.

are assumed to be uncorrelated with u_{it} and u_{it+1}, but are correlated with u_{it-1}. As first differencing causes such variables to become correlated with the error term Δu_{it}, they are instrumented by lagged values in $t-1$ and earlier. In particular, we allow all human capital variables, the capital-labour ratio and the alternative wage to be predetermined in order to capture potential feedback effects from wages in period t on those covariates in subsequent periods. To test the validity of the moment conditions, we present the 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-order serial correlation of the error terms (testing the null of no second-order serial correlation).

It is important to 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 of major importance since the dependent variable, the average wage, and some of the explanatory variables, such as quasi-rents as well as establishment size, are constructed using some of the same quantities (in particular the employment level). As a result, measurement error in these variables can induce spurious correlations between these explanatory variables and the dependent variable.

Table A3 in the Appendix gives the results of the differenced GMM estimates.[14] While Model (1) contains the static specification, Model (2) contains the simplest dynamic specification adding solely the lagged wage to the explanatory variables. Model (3) additionally includes lagged rents, while Model (4) contains lags of rents up to $t-2$. Table A3 contains estimates for time-varying regressors only, since first-differencing eliminates all time-invariant explanatory variables.[15] Turning to the main variables of interest, the signs of the rent-coefficients exhibit the same pattern as the POLS estimates of Model (6) in Table 2. While the rent-coefficient is always significantly positive for uncovered establishments, wages appear to be less sensitive to rents in establishments that are covered by a collective wage agreement. Including the lagged wage as a further explanatory variable in Model (2) reduces the rent-coefficients somewhat. As mentioned earlier, using lagged rents in $t-2$ as instruments for contemporaneous rents requires that they do not enter eq. (3) as explanatory variables. To check the robustness of our findings, we therefore include lagged rents up to $t-2$ in Model (3) and (4). While lags of rents in $t-1$ are

[14] All estimations have been carried out using the "XTABOND2"-procedure in STATA 8.0 SE.

[15] In our sample, time-invariant variables are the ownership dummy, the east-west dummy and the industry dummies. The collective bargaining dummies and the works council dummy are time varying binary regressors.

found to be insignificant in Model (3), lagged rents in $t-2$ enter Model (4) significantly, indicating that wages do not only respond to contemporary establishment performance, but also to past profitability conditions. In specifications (3) and (4), the effects of (contemporaneous) rents on wages in uncovered establishments are reduced, but still remain significant, once lagged rents up to $t-2$ are controlled for. The last rows in the second part of Table A3 show that all specifications pass the test of overidentifying restrictions and the AR(2)-test. The last two rows in the first part of Table A3 report p-values of Wald-statistics testing the null of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$ for the contemporaneous rent coefficients. The values indicate that wages appear to be completely insensitive to profitability conditions in establishments that are covered by a collective agreement, irrespective of whether the agreement is industry or firm-specific.

With respect to the remaining covariates, the performance of the differenced GMM estimates turns out to be rather unsatisfactory: although the lagged wage enters specification (2) and (3) with its expected sign, it is not significant and its point estimates appear to be implausibly low. In Model (4), the estimate is even negative. In all specifications, establishment size and the works council dummy are always insignificant and for the most part incorrectly signed. The capital-labour ratio is found to be significant, but with a negative sign. As regards the workforce composition, the estimates of $HIGHSHARE$ and $APPSHARE$ also seem to be poorly determined, as they enter almost all regressions with an unexpected sign. The remaining controls for the workforce composition enter with their expected sign (except for $BLUESHARE$ in Model (1)), but are not statistically significant.

In light of the poor performance of the differenced GMM estimates, Table A4 reports results using the System-GMM (SYS-GMM) estimator as proposed by Arellano and Bover (1995). This estimator is motivated by the problem that lagged levels of a variable are likely to be weak instruments for the equation in first-differences if the individual time series exhibits near unit root properties. Closer inspection of the time-series properties of the explanatory variables reveals that particularly the size variable and the capital-labour ratio appear to be close to a random walk.[16] The SYS-GMM estimator exploits additional moment conditions for the equation in levels using lagged differences as instruments in the levels equation. In particular, predetermined variables are instrumented by contemporaneous first-differences in the levels equation, whereas endogenous and lagged

[16]SYS-GMM estimates of a simple AR(1)-process yield a coefficient of about 0.94 for establishment size and of 0.91 for the capital-labour ratio.

dependent variables are instrumented by lagged first-differences (Bond 2002). To test the additional moment conditions implied by the SYS-GMM estimator as compared to the differenced GMM estimates in Table A3, we present in each column difference tests which refer to the respective specifications in Table A3. The test statistics are calculated as the differences between the Sargan/Hansen statistics of the SYS-GMM and those of the differenced GMM estimates and have an asymptotic χ^2 distribution under the null that the additional moment restrictions are valid.

Overall, the SYS-GMM estimates appear to be more satisfactory than the differenced GMM results. The lagged wage enters all specifications with its expected sign and its estimates are considerably higher than the differenced GMM estimates, suggesting that the latter are severely downward biased. In all specifications, establishment size is found to have a significantly positive impact on average wages and is estimated much more precisely than in the differenced GMM specification. This is consistent with the random-walk property of this variable, indicating that the lagged level of establishment size is a weak instrument for first-differences.

From the human capital covariates, only $FEMSHARE$ and $BLUESHARE$ enter all regressions with their expected sign. The remaining worker controls are mostly incorrectly signed and not significant. Turning to the impact of rents on average wages, the estimates offer a similar picture as the differenced GMM results: in uncovered establishments, quasi-rents exert a positive impact on wages, while wages are generally found to be less sensitive to rents in establishments that are covered by a collective wage agreement. In all specifications, a Wald-Test fails to reject the null of $\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$. Similar to the differenced GMM estimates, the effects of contemporaneous quasi-rents on wages in uncovered establishments are further reduced but remain still significant, once lagged wages and lagged quasi-rents up to $t-1$ are controlled for. However, controlling for lagged rents up to $t-2$ leads to an insignificant rent-coefficient in uncovered establishments, which is slightly lower than that obtained by the differenced GMM estimates (Model (4)). All specifications pass the test of overidentifying restrictions and the AR(2)-test. Moreover, the difference Sargan/Hansen statistic testing the additional moment restrictions as compared to Table A3 confirms their validity in all specifications except for Model (3) (with a $p-$value of 0.056).

Finally, all specifications were re-run assuming that all predetermined explanatory variables are uncorrelated with the time-invariant establishment-specific effect. At least in specification (4), this causes all human capital and establishment covariates to enter with their expected sign. However, the remaining results from these regressions only differ slightly from those shown in Table A4, so that we do not report them here. Most importantly, the estimates of the rent-coefficients are very similar to Table A4. In specifications (2) to (4), a Wald-Test again fails to reject the null of a zero-coefficient on contemporaneous rents under centralised as well as firm-specific agreements ($\beta_0 = -\beta_{\pi\_CENT}$ and $\beta_0 = -\beta_{\pi\_DECENT}$). Only in Model (1), can the null of a zero-coefficient be rejected under centralised agreements at the 10 per cent level (with a $p-$value of 0.068). The coefficients on contemporaneous rents in uncovered establishments are slightly larger than those in Table A4, ranging between 0.095 in Model (1) and 0.056 in Model (4).

Comparing the GMM estimates of the rent-sharing coefficients to the POLS estimates reveals that the POLS estimates still yield positive estimates of the rent-sharing coefficient in covered establishments, whereas the SYS-GMM-results accounting for unobserved heterogeneity and endogeneity of rents point to a rent-sharing coefficient of zero. This finding is indicative of the presence of unobserved factors in covered establishments which are positively correlated with profits and impact positively upon wages. One such factor may be that a compressed wage structure under centralised wage contracts causes firms to upgrade the quality of their workforce. This might lead to higher unobserved worker quality in such firms and therefore to upward-biased estimates in the simple POLS-specification. Comparing the GMM estimates of the rent-sharing coefficients to the POLS estimates in uncovered establishments points to similar figures. While the POLS-coefficient in uncovered establishments amounts to 0.061, the SYS-GMM estimates range between 0.044 and 0.095. Given these coefficients and mean wages and quasi-rents per employee of 38.78 and 37.15 in uncovered establishments, the elasticity of the average wage with respect to contemporaneous quasi-rents is of the magnitude 0.042 to 0.091. How do these results compare to other estimates for Germany? Hübler and König (1998) use data from the Hanover establishment panel and report an elasticity of about 0.12, while Klodt (2000: pp.172-182) finds an elasticity of 0.14 using the same data set. Compared to these figures, our estimate of the contemporaneous rent-coefficient in uncovered establishments therefore appears to be rather low. However, it needs to be emphasised that these studies do not allow the rent-coefficient to vary with collective bargaining coverage. Given the variability of rents, our results suggest that the quantitative role of rent-sharing in wage determination is nevertheless substantial: calculating the share of variance in the distribution of wages due to the variability in rents, it can be shown that the variability in per-capita rents explains about 15.7 to 33.9 per cent of the variability in (average) establishment wages.[17]

For centralised wage-agreements, the invariance of wages against establishment-specific profitability indicates that the fraction of establishments making use of flexibility provisions seems to be rather negligible. Even though firms may pay wages above the going rate and may adopt opt-out clauses, this potential for adjustments to local profitability conditions appears to be largely unused.[18] Even more striking is the invariance of wages against local profits in establishments that are subject to a firm-specific wage contract. Although this result is to be interpreted with caution as the number of observations with a firm-specific wage contract is rather small, it does not seem to confirm our first hypothesis which led us to expect the sensitivity of wages to profits under firm-specific contracts to be larger than in uncovered establishments. In sum, these findings lend support to our second hypothesis that unions favour a compressed intra-industry wage structure and suppress inter-firm wage differentials.

5 Robustness Checks

5.1 Sample Selection

The use of dynamic panel data methods imposes strong restrictions on the size of our final sample, since we have to exclude all establishments featuring less than 4 consecutive time series observations. Tables A2a and A2b in the Appendix compare sample statistics for the original sample and the final sample used in the preceding analysis. The figures show that establishments subject to a collective contract are on average considerably larger and more capital-intensive in the final sample than in the original sample. The differences for uncovered establishments mainly concern the qualification structure, with a larger fraction of qualified blue-collar employees in the final sample as compared to the original statistics. It is clear that this sample selection might bias our estimates, although the direction of bias is not clear a-priori. For example, unions might want to suppress rent-sharing in

[17] This calculation is performed under the assumption that 95 per cent of the mass of a symmetric distribution is within plus or minus 2 standard deviations of the mean. The contribution of the variability of rents to the variability of wages can then be calculated as:
$\frac{\beta_\pi(\overline{\pi}+2\sigma_\pi)-\beta_\pi(\overline{\pi}-2\sigma_\pi)}{(\overline{w}+2\sigma_w)-(\overline{w}-2\sigma_w)} = \frac{\beta_\pi \cdot \sigma_\pi}{\sigma_w}$ (see e.g. Margolis and Salvanes 2001).

[18] This finding corroborates the results of Franz and Pfeiffer (2003), which are based on an employer survey of about 800 German firms. Their results indicate that only 18 per cent of those employers that covered by a collective contract allowing for hardship clauses make use of such provisions.

large establishments due to high transaction costs. Efficiency wage considerations, which might also play a role in explaining a positive profit effect, lead us to expect the wage-profit correlation to increase with establishment size and capital-intensity. To assess the importance and direction of bias involved with our sample selection, we re-ran the POLS regressions using the original sample of 3,515 establishments.

Re-running the POLS regressions separately by bargaining coverage on the original sample gives point estimates of the rent-coefficients of 0.045, 0.020 and 0.024 (for no-coverage, centralised contracts and firm-specific contracts, respectively) as compared to 0.071, 0.013 and 0.032 for the final sample. However, the differences in the rent-coefficient are statistically significant only for uncovered establishments. This leads us to conclude that for covered establishments our results are fairly robust to sample selection, whereas the selection appears to involve an upward bias of the rent-coefficient for uncovered establishments. Note that this might be caused by the differences in qualification structures. If rent-sharing is more relevant for qualified blue-collar workers, then the overrepresentation of establishments employing large fractions of such employees in the final sample will bias the extent of rent-sharing upwards. However, it should be noted that this does not affect the robustness of our results concerning the overall pattern of wage responses, since for uncovered establishments the rent-coefficient is estimated to be significantly positive, whereas wages in covered establishments are still found to be completely invariant against establishment-specific profitability conditions.

5.2 Alternative Interpretations of the Correlation between Wages and Rents

Several authors have emphasised that a positive correlation between quasi-rents and wages need not necessarily imply rent-sharing, but may simply reflect movements of labour demand along an upward sloping labour supply curve (see e.g. Blanchflower et al. 1996, van Reenen 1996, Hildreth and Oswald 1997). If this were the case, the inclusion of the employment level should render the coefficient on quasi-rents insignificant. For uncovered establishments we are able to rule out such an alternative interpretation, since the positive coefficient on quasi-rents is robust to the inclusion of establishment size as an explanatory variable in all regressions. As a further robustness check we have also included employment growth a proxy for demand shocks in the differenced GMM regressions, which left the coefficients on quasi-rents also largely unchanged.[19]

5.3 Endogeneity of the Bargaining Regime

Thus far, we have considered the collective bargaining regime as exogenous. However, in Germany firms may leave their employers' associations and may, thus, to some extent influence the choice of the bargaining regime. This shows up in our data, where the share of sample establishments subject to a centralised contract declined from 82 per cent in 1995 to 75 per cent in 2001, and the fraction of establishments with a firm-specific agreement decreased from 8.5 to 6 per cent. What is relevant for our estimates is that a non-random selection into the regimes might bias our rent-sharing coefficients, particularly if a firm's choice is correlated with its profitability conditions. If, for example, centralised contracts shelter firms against excessive rent-sharing at the firm level, highly profitable firms might systematically select themselves into the centralised regime. To check the robustness of our findings to the endogeneity of the collective bargaining regime, we first estimate eq. (1) separately by bargaining coverage using a selection model which accounts for a potential non-random selection into the three wage determination regimes. To assess the importance of a potential endogeneity bias, we subsequently compare the estimates with the corresponding POLS regression results. Defining regimes R_1, R_2 and R_3 as wage determination under no-coverage, centralised contracts and firm-specific contracts, respectively, the wage equations for each regime R_j, $j = 1, 2, 3$, become

$$w_j = \alpha_j + \beta_{j\pi} \cdot \pi_j + \gamma_j \cdot \mathbf{x}'_j + \delta_j \cdot \mathbf{s}'_j + u_j \quad \text{if } R_j = 1, \tag{4}$$

where the variance of u_j is given by σ^2 and the indices i, t are suppressed for expositional convenience. To account for $E(u_j|R_j = 1)$, we adopt an extension of the two-step selection-bias correction method developed by Heckman (1979), which has been proposed by Lee (1983). Assuming that selectivity into the regimes can be modelled as a multinomial logit with a vector of explanatory variables $\mathbf{z}'$ and a parameter vector θ_j, Lee (1983) shows that

$$E(w|R_j = 1) = \alpha_j + \beta_{j\pi} \cdot \pi_j + \gamma_j \cdot \mathbf{x}'_j + \delta_j \cdot \mathbf{s}'_j - \sigma\rho_j\lambda_j(\theta_j \cdot \mathbf{z}'). \tag{5}$$

where

$$\lambda_j(\ \theta_j \cdot \mathbf{z}')\ = \frac{\phi(\Phi^{-1}(P_j))}{P_j},\ j = 1, 2, 3 \text{ and } P_j = \frac{\exp(\theta_j \cdot \mathbf{z}')}{\sum\limits_k \exp(\theta_k \cdot \mathbf{z}')},\ j, k = 1, 2, 3, \tag{6}$$

[19]The results are not reported here, but are available on request.

and with ρ_j denoting the correlation-coefficient between u_j and the unobservables in the selection equation. From eq. (5) it can be seen that OLS estimates of eq. (4) are biased if $\lambda_j(\theta_j \cdot \mathbf{z}')$ is correlated with the observables in eq. (4) and u_j and the error term in the selection equation are correlated. Consistent estimates of all parameters of interest in eq. (5) may be obtained by a two-step procedure, where the first step involves the generation of predicted values of $\lambda_j(\theta_j \cdot \mathbf{z}')$ by estimating the selection equation using a multinomial logit approach. In the second step, the predicted values are added to eq. (4), which is estimated by OLS. In the selection equation, the vector of observables $\mathbf{z}'$ includes the observables in eq. (4) and further identifying covariates which are excluded from eq. (4). For the excluded observables, we choose (1) a dummy taking on the value of unity if an establishment belongs to a publicly listed company and (2) an establishment-age dummy indicating whether an establishment has been founded after 1990 or earlier. We believe that those variables provide appropriate identifying exclusion restrictions for several reasons. First, when testing the significance of those variables in eq. (4) estimated separately by bargaining coverage, the corresponding F-tests indicate that both variables have no direct significant impact on wages (with p-values of 0.18, 0.24 and 0.44). Second, the identifying variables all appear to be significant predictors of the bargaining regime, since the corresponding F-statistic is highly significant in the selection equation (with a p-value close to zero). Third, we argue that it is reasonable to assume that the identifying variables are exogenous in the selection equation, since they are unlikely to be influenced by unobservables affecting the bargaining regime.

Table A5 in the Appendix shows the POLS estimates and the selectivity-corrected estimates using the original sample of 3,515 establishments. The negative coefficients on $\lambda_j(\theta_j \cdot \mathbf{z}')$, which are estimates of $-\sigma\rho_j$, indicate that the choice of collective contracts is endogenous, with the error term in the selection equation and u_j being positively correlated. In the uncovered regime, the coefficient on $\lambda_j(\theta_j \cdot \mathbf{z}')$ is positive, but not significantly different from zero. The direction of bias under collective contracts depends on the correlation between $\lambda_j(\theta_j \cdot \mathbf{z}')$ and the covariates in eq. (4). Given that $\lambda_j(\theta_j \cdot \mathbf{z}')$ is decreasing in P_j and in all covariates that have a positive impact on P_j, the negative estimates of $-\sigma\rho_j$ suggest that the OLS-coefficients on covariates that are positively correlated with the choice of either centralised or firm-specific contracts should be upward biased. Multinomial logit estimates show that the log-odds ratio of choosing centralised contracts as compared to no-coverage increases significantly with quasi-rents.[20] By contrast, the ef-

[20]The results are not reported here, but are available on request.

fect of quasi-rents on the log-odds ratio of choosing firm-specific contracts as compared to no-coverage is found to be insignificant. The resulting marginal effects of quasi-rents on centralised contracts, firm-specific contracts and no-coverage are 0.0002, -0.00003 and -0.0002. Given the estimated coefficients on $\lambda_j(\theta_j \cdot \mathbf{z}')$, we expect that correcting for selectivity should make little difference under no-coverage and firm-specific contracts, and should lead to a decline in the coefficient under centralised contracts. Indeed, the selectivity-corrected rent-sharing coefficient is found to be slightly lower for centralised contracts than the corresponding OLS coefficient. As selectivity plays no major role for the uncovered regime, most of the selectivity-corrected estimates do not substantially differ from the OLS estimates. If a selectivity correction changes anything at all, it results in even smaller rent coefficients under centralised contracts. In sum, this leads us to conclude that the overall pattern of rent-sharing across the three wage determination regimes appears to be quite robust to the endogeneity of the bargaining regime.

6 Summary and Conclusions

The aim of this paper was twofold: first, we addressed the question of whether German wages respond to firm-specific profitability conditions and second, we explored the extent to which the sensitivity of wages to firm profits depends on collective bargaining coverage. The institutional discussion has shown that firm-specific contracts and flexibility provisions under centralised contracts provide a means to adjust wages to local conditions at the firm level and that such adjustments are generally influenced by industry-wide unions which may retain control over centralised union objectives. Provided those flexibility provisions are used, bargaining power considerations lead us to expect wages to react more strongly to local conditions in firms that are covered by a collective contract than in uncovered firms. However, there are various reasons why unions might not want to adjust wages to local firm performance, such as high transaction costs or workers' demand for income insurance. We therefore take our empirical findings as a test of whether flexibility provisions are really exploited or whether unions suppress inter-firm wage dispersion due to heterogeneous firm performance.

Using data from the *IAB-Establishment Panel*, the results of our empirical analysis offer a remarkably consistent picture: in general, rent-sharing is found to be an empirically relevant phenomenon in Germany. However, the extent of rent-sharing seems to be significantly lower in establishments that are subject to a collective wage agreement - irrespective of whether the agreement is industry or firm-specific. While POLS estimates still yield positive estimates of the rent-sharing coefficient in covered establishments, GMM-results accounting for unobserved heterogeneity and the endogeneity of rents point to a rent-sharing coefficient of zero. This finding is indicative of the presence of unobserved factors in covered establishments which are positively correlated with profits and impact positively upon wages. One such factor may be that a compressed intra-firm wage structure under collective wage contracts causes establishments to upgrade the quality of their workforce. This might lead to higher unobserved worker productivity in such establishments and therefore to upward-biased estimates in the simple POLS-specification. Finally, we find the pattern of rent-sharing to be robust to sample selection and the endogeneity of the bargaining regime.

For centralised wage agreements, the invariance of wages against local profits suggests that the use of flexibility provisions in central wage agreements appears to be empirically negligible. Even though firms may pay wages above the going rate and may make use of opt-out clauses, the potential for adjustments to local profitability conditions appears to be largely unused. A similar result holds for wage determination under firm-specific wage contracts. As such contracts are generally concluded by industry-specific unions, one possible explanation might be that a considerable fraction of firm-specific contracts simply adopts wage bargains negotiated in the corresponding industry agreement. Taken together, our results seem to support the notion that unions favour a compressed intra-industry wage structure and suppress firm-level rent-sharing, either due to workers' demand for income insurance or due to high transaction costs.

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A Appendix

A.1 Construction of Establishment Variables

Variable	Definition
w:	Annual aggregate wage bill*) (= 12*wage bill reported for the month June) divided by the number of employees reported for the month June.
SIZE	Number of employees reported for the month June averaged over the present and preceding year.
π:	Quasi-rents are constructed by subtracting material costs and the alternative wage bill from annual sales. Per capita values are obtained by dividing quasi-rents by establishment size. Nominal values are deflated by a sector-specific (two-digit) producer price index.
$\overline{w}$	The average annual sectoral wage per worker is approximated by the weighted average of industry-specific wages for blue and white-collar workers (separately for western and eastern Germany), with the weights being the establishment-specific shares of those worker groups in the total work force. Average hourly industrial wages of blue-collar workers are converted into monthly wages by multiplying them with establishment-specific average working time. Information on average sectoral wages of white-collar workers is available only on a monthly basis. Monthly values are converted into annual values by multiplying them with the factor 12.

Table A1: Description of establishment variables

... *to be continued on next page*

...continue Table A1

Variable	Definition
K/L	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.
WCOUNCIL	Dummy=1 if works council is present. In some waves (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.
DECENT	Dummy=1 if establishment is bound to a firm-specific agreement.
CENT	Dummy=1 if establishment is bound to a industry-specific agreement.
OWN	Dummy=1 if establishment is part of a company that is owned by a person with unlimited liabilities.

Note: *) Exclusive of employers' social security contributions as well as fringe benefits.
**) To calculate the capital stock in the first period, we set d=0.1 and g=0.05

Table A1: Description of establishment variables

A.2 Descriptive Statistics by Bargaining Coverage

Variables	CENT Mean	CENT Std.-Dev.	DECENT Mean	DECENT Std.-Dev.	NO-COVERAGE Mean	NO-COVERAGE Std.-Dev.
w	54.64	17.66	48.37	16.04	38.78	17.95
π	85.14	102.28	68.40	90.40	37.15	64.13
$\overline{w}$	53.42	10.87	49.63	10.93	46.61	11.97
HIGHSHARE	0.27	0.20	0.26	0.20	0.19	0.19
BLUESHARE	0.39	0.22	0.47	0.23	0.48	0.25
APPSHARE	0.04	0.06	0.04	0.05	0.05	0.07
FEMSHARE	0.26	0.19	0.24	0.19	0.29	0.24
SIZE	865.33	3,123.63	417.65	870.11	75.81	161.27
WCOUNCIL	0.80	0.40	0.75	0.44	0.24	0.43
K/L	204.85	400.02	721.65	3,780.18	150.59	385.72
EAST	0.29	0.45	0.56	0.50	0.68	0.47
OWN	0.20	0.40	0.12	0.32	0.28	0.45
Obs.	2,120		392		899	

Table A2a: Final sample

Variables	CENT Mean	CENT Std.-Dev.	DECENT Mean	DECENT Std.-Dev.	NO-COVERAGE Mean	NO-COVERAGE Std.-Dev.
w	53.80	19.02	47.65	18.53	38.68	18.24
π	81.95	103.10	71.42	106.92	40.26	74.40
$\overline{w}$	53.74	10.95	50.06	11.29	48.46	12.15
HIGHSHARE	0.26	0.20	0.24	0.19	0.20	0.21
BLUESHARE	0.39	0.22	0.46	0.24	0.45	0.26
APPSHARE	0.05	0.06	0.04	0.06	0.05	0.08
FEMSHARE	0.25	0.20	0.27	0.21	0.31	0.25
SIZE	650.80	2,214.64	366.84	1,019.92	72.09	161.91
WCOUNCIL	0.75	0.43	0.69	0.46	0.21	0.41
K/L	177.55	381.70	405.84	2,537.03	140.07	362.74
EAST	0.27	0.45	0.52	0.50	0.66	0.47
OWN	0.21	0.41	0.18	0.38	0.33	0.47
Obs.	4,751		892		2,974	

Source: IAB-Establishment Panel 1995-2002. Entries are unweighted.

Note: All monetary values are measured in DM 1,000. 1 € corresponds to DM 1.95583.

Table A2b: Original sample

A.3 Regression Results

Model	*1*	*2*	*3*	*4*
$\overline{w}$	0.403**	0.360*	0.322	0.330
	(0.175)	(0.190)	(0.198)	(0.211)
w(t-1)		0.064	0.066	-0.002
		(0.042)	(0.043)	(0.056)
π	0.077***	0.064***	0.057**	0.047*
	(0.029)	(0.025)	(0.024)	(0.028)
$\pi*$CENT	-0.079**	-0.073**	-0.074***	-0.061**
	(0.034)	(0.030)	(0.028)	(0.030)
$\pi*$DECENT	-0.092**	-0.084**	-0.082**	-0.065*
	(0.042)	(0.041)	(0.040)	(0.038)
π(t-1)			0.023	0.034
			(0.025)	(0.035)
$\pi*$CENT(t-1)			-0.011	-0.028
			(0.021)	(0.028)
$\pi*$DECENT(t-1)			-0.016	-0.018
			(0.022)	(0.035)
π(t-2)				0.043*
				(0.024)
$\pi*$CENT(t-2)				-0.047*
				(0.024)
$\pi*$DECENT(t-2)				-0.036
				(0.028)
$\pi = -\pi*$CENT ($p-$value)	0.908	0.586	0.321	0.429
$\pi = -\pi*$DECENT ($p-$value)	0.586	0.504	0.415	0.481

*Significant at 10%-level,**significant at 5%-level, ***significant at 1%-level.

Table A3: Differenced GMM regression results

... to be continued on next page

... continue Table A3

Model	*1*	*2*	*3*	*4*
HIGHSHARE	-12.835**	-9.038*	-8.310	-11.834
	(5.133)	(5.346)	(5.140)	(6.927)
BLUESHARE	-1.280	2.578	2.379	5.706
	(2.568)	(3.098)	(3.003)	(3.549)
APPSHARE	-1.818	8.088	6.258	11.176
	(11.414)	(13.274)	(13.294)	(16.319)
FEMSHARE	-1.834	-5.925	-5.832	-6.503
	(8.722)	(9.083)	(8.933)	(10.561)
SIZE	-0.001	-3.0e^{-04}	-7.0e^{-04}	0.002
	(0.003)	(0.003)	(0.003)	(0.004)
SIZE^2	-1.92e^{-08}	-1.61e^{-08}	-6.34e^{-09}	-1.30e^{-08}
	(2.98e^{-08})	(2.96e^{-08})	(3.20e^{-08})	(4.55e^{-08})
CENT	3.652**	3.694**	3.863**	3.252*
	(1.809)	(1.723)	(1.672)	(1.665)
DECENT	4.179**	3.981*	3.783*	2.794
	(2.020)	(2.074)	(2.041)	(1.994)
WCOUNCIL	-1.702	-2.243	-2.490	-1.043
	(2.080)	(2.125)	(2.108)	(2.618)
K/L	-0.001***	-0.002***	-0.002***	-0.002***
	(4.0e^{-04})	(3.0e^{-04})	(3.0e^{-04})	(3.0e^{-04})
Sargan/Hansen ($p-$value)	0.288	0.463	0.604	0.443
AR(2) ($p-$value)	0.932	0.460	0.688	0.401
Establishments	661	661	661	661
Observations	2,750	2,089	2,089	1,428

Note: The dependent variable is the aggregate per-capita wage bill. All variables are first-differenced. Results are reported for one-step differenced GMM estimators. All specifications include time dummies. Heteroscedasticity-robust standard errors are in parentheses. *Significant at 10%-level, **significant at 5%-level, ***significant at 1%-level.

Table A3: Differenced GMM regression results

Model	*1*	*2*	*3*	*4*
$\overline{w}$	0.386**	0.374**	0.323**	0.367*
	(0.156)	(0.149)	(0.161)	(0.197)
w(t-1)		0.221***	0.228***	0.227***
		(0.046)	(0.046)	(0.051)
π	0.084***	0.072***	0.059**	0.044
	(0.024)	(0.025)	(0.026)	(0.029)
$\pi*$CENT	-0.070**	-0.067**	-0.070**	-0.053*
	(0.029)	(0.030)	(0.028)	(0.030)
$\pi*$DECENT	-0.062**	-0.061**	-0.057**	-0.053
	(0.031)	(0.030)	(0.028)	(0.035)
π(t-1)			0.007	0.006
			(0.016)	(0.019)
$\pi*$CENT(t-1)			0.014	0.011
			(0.013)	(0.016)
$\pi*$DECENT(t-1)			0.006	0.024
			(0.015)	(0.020)
π(t-2)				0.019
				(0.014)
$\pi*$CENT(t-2)				-0.023
				(0.016)
$\pi*$DECENT(t-2)				-0.004
				(0.019)
π = -$\pi*$CENT ($p-$value)	0.267	0.739	0.389	0.535
π = -$\pi*DE$CENT ($p-$value)	0.204	0.483	0.906	0.689

*Significant at 10%-level,**significant at 5%-level, ***significant at 1%-level.

Table A4: SYS-GMM regression results

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... continue Table A4

Model	*1*	*2*	*3*	*4*
HIGHSHARE	-7.306	-5.846	-4.884	-3.073
	(4.756)	(4.976)	(4.867)	(5.935)
BLUESHARE	3.618	6.267*	6.140*	11.751**
	(2.990)	(3.606)	(3.537)	(4.647)
APPSHARE	-10.914	-7.663	-8.394	3.351
	(11.145)	(11.573)	(11.883)	(19.267)
FEMSHARE	-4.603	-5.326	-4.706	-4.775
	(4.929)	(4.720)	(4.688)	(5.291)
SIZE	0.002***	0.002**	0.002**	0.002**
	(7.0e^{-04})	(7.0e^{-04})	(7.0e^{-04})	(8.0e^{-04})
SIZE^2	-2.69e^{-08*}	-2.22e^{-08}	-2.28e^{-08}	-2.80e^{-08}
	(1.46e^{-08})	1.44e^{-08}	(1.49e^{-08})	(1.76e^{-08})
CENT	4.850***	4.809***	5.279***	3.710*
	(1.861)	(1.856)	(1.818)	(2.066)
DECENT	3.258**	3.613**	3.498**	4.276*
	(1.636)	(1.627)	(1.690)	(2.479)
WCOUNCIL	6.210**	3.882	3.841	4.250
	(2.633)	(2.392)	(2.395)	(2.796)
K/L	3.67e^{-06}	-2.0e^{-04}	-1.0e^{-04}	-3.0e^{-04}
	(1.0e^{-04})	(2.0e^{-04})	(2.0e^{-04})	(2.0e^{-04})
Sargan/Hansen ($p-$value)	0.312	0.346	0.312	0.318
Diff. Test comp. to Table A2 ($p-$value)	0.484	0.220	0.056	0.184
AR(2) ($p-$value)	0.713	0.128	0.204	0.118
Establishments	661	661	661	661
Observations	2,750	2,750	2,750	2,089

Note: The dependent variable is the aggregate per-capita wage bill. Results are reported for one-step SYS-GMM estimators. All specifications include time dummies, 15 industry dummies as well as an east-west and an ownership dummy. All endogenous and predetermined variables are assumed to be correlated with the establishment-specific effect.
Heteroscedasticity-robust standard errors are in parentheses.
*Significant at 10%-level,**significant at 5%-level, ***significant at 1%-level.

Table A4: SYS-GMM regression results

	CENT		DECENT		NO-COVERAGE	
	OLS	Selectivity corrected	OLS	Selectivity corrected	OLS	Selectivity corrected
π	0.020***	0.019***	0.024***	0.025***	0.045***	0.045***
	(0.003)	(0.003)	(0.009)	(0.010)	(0.006)	(0.005)
$\overline{w}$	0.547***	0.619***	0.658***	0.671***	0.259***	0.268***
	(0.063)	(0.072)	(0.157)	(0.157)	(0.066)	(0.083)
HIGHSHARE	17.996***	15.715***	14.622***	16.117***	14.544***	14.284***
	(1.923)	(1.953)	(4.487)	(4.442)	(2.064)	(2.482)
BLUESHARE	4.944***	4.485***	5.259**	3.696	6.433***	6.306***
	(1.174)	(1.247)	(2.183)	(2.524)	(1.076)	(1.261)
APPSHARE	1.562	-1.451	8.141	11.216	-6.885*	-7.050**
	(4.007)	(4.541)	(9.067)	(11.931)	(3.638)	(3.447)
FEMSHARE	-19.017***	-17.510***	-14.397***	-15.159***	-12.565***	-12.291***
	(1.327)	(1.447)	(2.924)	(2.995)	(1.184)	(1.670)
SIZE	0.002***	0.001***	0.001	0.002	0.023***	0.022***
	($3.00e^{-04}$)	($3.00e^{-04}$)	(0.001)	(0.002)	(0.004)	(0.005)
SIZE^2	$-2.60e^{-08}$**	$-1.92e^{-08}$	$-4.10e^{-08}$	$-6.90e^{-08}$	$-8.81e^{-06}$***	$-8.79e^{-06}$***
	($1.07e^{-08}$)	($1.29e^{-08}$)	($8.74e^{-08}$)	($2.24e^{-07}$)	($2.06e^{-06}$)	($2.38e^{-06}$)
WCOUNCIL	8.491***	5.455***	7.470***	4.141**	3.256***	2.763
	(0.595)	(1.075)	(1.092)	(2.017)	(0.672)	(2.344)
K/L	0.002***	0.002***	$6.28e^{-05}$	-0.0004**	$2.00e^{-04}$	$2.00e^{-04}$
	($7.00e^{-04}$)	($7.00e^{-04}$)	($1.10e^{-04}$)	($3.00e^{-04}$)	($8.00e^{-04}$)	($8.00e^{-04}$)
OWN	-4.442***	-4.825***	-4.720***	-4.064**	-6.524***	-6.598***
	(0.556)	(0.622)	(1.334)	(1.345)	(0.511)	(0.698)
EAST	-3.194***	0.206	-3.588	-5.487**	-8.713***	-8.308***
	(1.009)	(1.431)	(2.240)	(2.253)	(1.020)	(2.165)
Intercept	8.036***	8.752*	5.050	21.610*	24.197***	22.453**
	(3.746)	(4.784)	(9.511)	(11.367)	(4.844)	(10.394)
$\lambda_j(\theta_j \cdot z')$		-6.124***		-9.569***		0.781
		(1.814)		(3.388)		(3.531)
Ind.-/Time	Yes	Yes	Yes	Yes	Yes	Yes
Observations	4,725	4,725	892	892	2,970	2,970

Note: The dependent variable is the aggregate per-capita wage bill. Heteroscedasticity-robust (POLS) and bootstrapped standard errors (selectivity corrected results - 100 repetitions) are in parentheses. All models include time dummies and 15 industry dummies.
*Significant at 10%-level ** Significant at 5%-level ***Significant at 1%-level

Table A5: Selectivity corrected regression results