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Intra-firm wage inequality and firm performance
– First evidence from German linked
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Intra-firm wage inequality and firm performance – First evidence from German linked employer-employee-data

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Abstract

Economic theory suggests both positive and negative relationships between intra-firm wage inequality and productivity. This paper contributes to the growing empirical literature on this subject. We combine German employer-employee-data for the years 1995-2005 with inequality measures using the whole wage distribution of a firm and rely on dynamic panel-data estimators to control for unobserved heterogeneity, simultaneity problems and possible state dependence. Our results indicate a relative minor influence of intra-firm wage inequality on firm productivity. If anything, they provide some support for a view suggesting that some inequality may be beneficial, while too much leads to a detrimental effect on productivity.

Keywords: Wage dispersion, labor productivity

JEL Classification: J31, M52

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1 Introduction

Economic theory suggests several possible relationships between intra-firm wage inequality and labor productivity. On the one hand, tournament theory and related approaches argue that a high wage spread implies high gains associated with promotions for highly productive workers. On the other hand, several theories point towards a negative effect of a wage spread that becomes too large.

The first group of arguments is related to efficiency wage theory. If a firm rewards relative performance, workers are rewarded based on a comparison with their peers. Since wage increases or promotions are awarded only to the very best, workers have an incentive to work as hard as possible. High wage dispersion in that scenario implies high wage gains for the top-performers and should consequently be associated with higher productivity.

A positive relationship between wage inequality and labor productivity may also be expected if a firm chooses a performance related pay scheme that rewards absolute performance, e.g. piece-rates, and workers are heterogeneous with respect to their ability (see Lazear 1996, 2000 for a case study and theoretical discussion). Performance related pay should create an incentive for workers to work harder thus leading to a rise in average productivity. These incentives are stronger for high-ability workers who find it easier to raise their individual output levels than their low-ability counterparts. Since these differences in individual output directly map into differences in pay under a performance pay scheme, wage dispersion may be expected to rise.

While these theoretical ideas suggest a positive relationship between wage inequality and worker performance, others point at a different direction. Lazear (1989) argues that high wage gains in tournaments not only create incentives to work harder, but might also

induce workers to sabotage the work of other members of their respective comparison group. Greater pay equality within these reference groups makes this deviant behavior less worthwhile thus reducing the efficiency losses.

A similar argument relates to rent-seeking by workers (see Milgrom 1988, Milgrom and Roberts 1990). Instead of conducting productive work, workers may choose to allocate more time to unproductive rent-seeking, e.g. trying to convince their supervisors to change the wage structure in their favor. The incentive to engage in these redistributive activities is clearly larger if the wage spread is large. A simple remedy for a firm faced with this problem is to compress the wage structure thus reducing the potential gains from rent-seeking.

Frey (1997) argues that monetary rewards may crowd out intrinsic motivation of workers. If a firm raises wage inequality to create incentives for workers to become more productive, this crowding-out might work in the opposite direction thus leading to a smaller, if not negative effect on labor productivity.

Finally, another strand of the literature focuses on aspects of fairness or cohesion between workers. Akerlof and Yellen (1988, 1990) derive a variation of efficiency wage theory based on “fair” wages. In their model, workers do not only care about their absolute wage but base their effort on a comparison between their actual wage and a wage they consider fair. This fair wage is in turn influenced by the wages of a comparison group, e.g. similar workers in the same firm. In this model, greater wage compression may lead to higher productivity by means of a wage structure that places more workers close to their fair wage.

A similar argument is used by Levine (1991) who considers the importance of cohesion between workers in certain work environments. His arguments suggest that when cohesion

is important, e.g. in firms that rely heavily on group work, a certain amount of wage compression will be beneficial as it raises overall productivity.

Empirically, the direction of the relationship between wage dispersion and labor productivity remains unanswered. A number of studies, reviewed in greater detail in section 2, have dealt with this question without reaching a definite consensus. Studies using evidence from professional sports usually find either a negative or insignificant relationship between wage inequality and productivity, while studies focusing on executive compensation or firm level inequality reach somewhat different conclusions.

This paper contributes to the literature by combining for the first time firm-level panel data and inequality measures built on the entire wage distribution within firms with an estimation strategy that accounts for endogeneity, unobserved heterogeneity and possible dynamics in the adjustment process of firm productivity to changes in the wage structure. More specifically, we use linked-employer-employee data from Germany for the years 1995 to 2005 to estimate the relationship between several measures of wage inequality and average productivity. To account for endogeneity and time-constant unobserved heterogeneity, we use dynamic Arellano-Bond panel-data estimators that also allow for possible state dependence in productivity.

The rest of this paper is organized as follows. Section 2 gives an overview of previous evidence from empirical studies. Section 3 describes the data used in this study while section 4 outlines our estimation strategy along with the estimator used. Results are presented in section 5. Section 6 concludes.

2 Previous evidence

Much of the previous econometric evidence relating wage inequality to the economic performance of organizations comes from either professional sports (e.g. Bloom 1999, Depken 2000, DeBrock et al. 2004, Frick et al. 2003, Harder 1992), executive compensation (e.g. Eriksson 1999, Leonhard 1990, Main et al. 1993) or from other special environments (see e.g. Pfeffer, Langton 1993 for college and university faculty). The results are generally mixed between as well as within these strands of the literature. A slight majority of the “sports papers” reports a detrimental effect of wage inequality, while the papers concerned with executives are more in favor of a positive relationship between performance and wage dispersion. One should note, however, that none of these results generalizes easily to the whole population of workers in a firm and the question what effects inequality in their wages might have on a firm’s performance.

In recent years a small, but growing numbers of papers, summarized in table 1, has used either firm or linked employer-employee-data to investigate the relationship between pay inequality among workers and firm performance. Cowherd and Levine (1992) use data on 102 business units from the UK and the USA. They use two measures of wage inequality, specifically lower management wages compared to those of hourly paid workers and lower compared to upper management wages, and relate those to a measure of product quality. Results from standard linear regressions reveal a negative relationship between these variables.

(TABLE 1 ABOUT HERE.)

In a study for Austria, Winter-Ebmer and Zweimüller use data from Social Security. The data encompasses all workers from 130 firms between 1975 and 1991. They construct a

measure of conditional wage inequality that takes into account the wage differences justified by different compositions of the workforce. More specifically, they estimate firm- and year-specific standard wage regressions and use the variance of the residuals as a measure of wage inequality. Since their data does not contain a measure of firm performance, they rely on the results from the wage regression to calculate standardized wages for typical workers as a proxy for productivity. Their results from group-mean regressions show a non-monotonic relationship for white-collar-workers where productivity increases with rising wage inequality at low levels and decreases if inequality becomes too large. For blue-collar workers the same overall relationship emerges. However, here most observations are to the left of the turning point, thus implying a positive, though decreasing relationship between wage inequality and productivity.

Using aggregate data from Sweden for the years 1964-1993 for industries and 1972-1993 for plants to estimate an augmented Cobb-Douglas-function, Hibbs and Locking (2000) find a positive association between wage inequality and firm performance, measured as value added and value added per worker. They also find a negative effect of between-industry wage inequality on aggregate output and productivity growth.

In a paper for Denmark, Bingley and Eriksson (2001) use linked employer-employee data from integrated database for labour market research (IDA) and the business statistics database (BSD). Their data contains the population of Danish firms that have at least five managers, five other white-collar workers and five blue-collar workers, as well as information on all their employees for the years 1992 to 1995, resulting in 22,665 firm-year observations. They measure productivity and worker effort on the firm level by using total factor productivity and sickness absence respectively. These are regressed on a measure of wage dispersion similar to that used by Winter-Ebmer and Zweimüller (1999), that is

residuals from a wage regressions, here calculated using all workers in the sample. Regarding the effects of pay inequality on productivity, their results for white-collar workers show the same non-monotonic relationship as found by Winter-Ebmer and Zweimüller (1999), while no effect can be found for blue-collar worker wage inequality. For both blue- and white-collar workers a higher pay spread is associated with reduced sickness absence and thus higher effort.

Using firm-level data from the UK, Beaumont and Harris (2003) look at the relationship between wage dispersion, measured as the ratio of manual to non-manual labor, and gross value added per worker. They estimate Cobb-Douglas-type production functions for several industries using Arellano-Bond dynamic panel estimators to account for the endogeneity of labor productivity, employment, capital stock and relative wages. Their results show positive effects of their measure of wage dispersion for most industries, more specifically electronic data processing, motor vehicles and engines, aerospace and miscellaneous foods, and a negative effect in one industry (pharmaceutics).

In another paper trying to account for the endogeneity of wage dispersion, Lallemand et al. (2004) use cross-sectional linked-employer-employee data from the Belgian 1995 Structure of Earnings Survey and the 1995 Structure of Business Survey. Their sample encompasses 17,490 individuals working for 397 firms from the private sector with at least 200 employees. They construct a conditional wage dispersion measure by using the same approach as Winter-Ebmer and Zweimüller (1999) and use it alongside several unconditional dispersion indicators in regressions on gross operating surplus per worker. As wage dispersion measures may be rendered endogenous by bonus payments, they instrument overall wage dispersion by the dispersion of wages excluding bonus payments in a 2SLS-equation. Their results show a positive and significant relationship between wage inequality and firm

performance which is larger for blue-collar workers.

Grund and Westergaard-Nielsen (2004) look at both the dispersion of wages and wage increases using Danish linked employer-employee data for firms with at least 20 employees for the years 1992 to 1997. Their findings indicate a hump-shaped relationship between wage dispersion and log value added that disappears when controlling for unobserved heterogeneity. The dispersion of wage growth is found to have a U-shaped influence, with most of the firms in the sample being situated on the decreasing part of that curve.

Also using linked employer-employee data for 1991 and 1995 from several Swedish data sources, Heyman (2005) estimates the relationship between several measures of wage dispersion among managers and white-collar employees and a variety of performance measures. He accounts for unobserved heterogeneity and the potential endogeneity of wage dispersion by estimating first differences from 1991 to 1995 and instrumenting the 1995 wage dispersion with their 1991 counterpart. His findings indicate a positive impact of both managerial and white-collar wage inequality on profits and average pay and a positive relationship between managerial wage inequality and variation in sales.

Finally, Jirjahn and Kraft (2007) use firm-level data from Germany to look into the interaction between (blue-collar) wage inequality, measured as the difference between the highest wage for a skilled worker and the lowest wage of an unskilled worker in the same establishment, pay and promotions schemes, collective bargaining, work councils and their impact on firm performance. While they cannot control for either unobserved heterogeneity or the potential endogeneity of wage dispersion, their results indicate that the effect of wage dispersion differs greatly when taking the aforementioned interactions into account.

Taken together, the evidence suggests an either positive or hump-shaped relationship

between wage inequality and firm performance. This relationship seems to be relatively robust to the exact definition of wage inequality and the performance measure used. It also seems somewhat stable over countries and time periods with the insignificant German results from Jirjahn and Kraft (2007) being an exception.

3 Data

This paper uses the linked employer-employee data of the Institute for Employment Research in Nuremberg, the so-called LIAB.¹ The LIAB is created by merging establishment information from the IAB Establishment Panel, a representative survey conducted annually since 1993, with employee information from notifications to German social security. This paper uses the cross-sectional version of the LIAB, currently available for the years 1993 to 2005, where the panel data on establishments is merged with cross-sectional information on all employees working in the respective establishment on July 30th of each year. Note that this results in an annual panel at the firm level.

The employee data originates from Social security information and is collected in the so called *employee history* by the Federal Employment Agency.² Employers are obliged to deliver annual information on their employees to social security by German law which is used for pension, health and unemployment insurance. The resulting data contains information on the begin and end of employment, daily wages, a person's age and sex, as well as several variables collected for statistical purposes, e.g. education.

This data is combined with firm level data from the IAB Establishment Panel by a

¹For a short introduction see Alda et al. (2005), detailed information (in German) is available in Jacobebbinghaus and Alda (2007) and on the website of the Research Data Center of the Federal Employment Agency in the Institute for Employment Research <http://fdz.iab.de>.

²More information on person-level data from German social security records can be found in Bender et al. (2000).

unique identification number used by the Federal Employment Agency. The IAB Establishment Panel is conducted annually since 1993 in West Germany, with East Germany joining in 1996.³ It is representative for all German firms with at least one worker covered by social security. Case numbers vary between approximately 4.000 and 16.000 cases per year. The resulting panel is unbalanced, though response rates are usually around 80% for plants that are interviewed repeatedly (Alda et al. 2005, p. 331). The data contains detailed annual information on the employee structure, industrial relations, economic conditions, establishment size, etc., as well as special surveys varying by year.⁴

To construct the estimation sample used in this paper, we take the following steps. In a first step, we restrict the sample to firms situated in West Germany. First, East German workers born in the Socialist German Democratic Republic might have perceptions of fairness or attitudes towards wage inequality that are fundamentally different from West German workers. Consequently we cannot expect the relationship between wage dispersion and labor productivity to be same in these two regions. Secondly, during the time period considered in this paper, the transformation process after the German reunification might still be in progress in East Germany, making the regions not fully comparable.

Since any measure of wage inequality is not particularly meaningful in very small firms, we restrict the sample to firms with five or more workers in at least one year. Additionally, as information on collective bargaining is only available from 1995 onwards, we restrict the sample to the years 1995 to 2005. Finally, to avoid problems with outliers, establishments with outcome measures in the top and bottom 1% are dropped, leaving us with ca. 21,000

³Detailed information on the IAB Establishment Panel (up to the year 2000) can be found in Kölling (2000).

⁴See Alda et al. 2005 for a short overview on the annual topics, a complete list of all items as well as complete questionnaires and additional information can be found on the website of the on the website of the Research Data Center of the Federal Employment Agency in the Institute for Employment Research <http://fdz.iab.de>.

firm-year-observation from 6,000 firms, the exact number depending on the specification and the estimator used.

When measuring wage inequality, it is useful to distinguish between two different concepts: overall wage inequality without taking differences in the composition of the workforce into account and inequality in comparable subgroups, e.g. between workers with similar characteristics. Depending on the exact nature of the data-generating process, both concepts might be relevant for labor productivity.

Consider first the case of tournament theories where a greater wage inequality may be taken as a sign for larger incentives associated with high performance. In that scenario, overall wage inequality might be relevant if workers believe that each wage level in a firm can actually be reached by high performance. Wage inequality within certain groups would better reflect the fact that most often workers will compete against similar peers for a promotion.

A similar argument holds for the fairness theories. Overall wage inequality would reflect workers sentiments towards the overall wage spread in a firm, e.g. the belief that upper management should only earn a certain multiple of the average blue-collar worker wage. Wage inequality between similar workers comes closer to the intuitive notion of fairness, specifically that an allocation is more unfair if equal workers are treated unequally.

In this paper, we use variants of both concepts of wage inequality. In a first step, wages above the social security contribution threshold are imputed by means of a Tobit-based imputation procedure (Gartner 2005) using standard wage regression variables as sex, age, education, occupational position, regions and industry. In a second step, the overall wage inequality is measured by the coefficient of variation using all wages from a

specific firm. Within group inequality is measured by a weighted average of coefficients of variation calculated for subgroups defined by age, occupational positions and education (five categories each), where the weights are the number of people in each subgroup.

To create a third measure of inequality, we rely on an approach similar to the one used by Winter-Ebmer and Zweimüller (1999): We estimate year-specific wage equations by Tobit-regressions to account for the top-censoring of the wage variable using the same variables as in the imputation described above. Inequality is then measured by the standard deviation of residuals for all persons in a given firm.

Labor productivity is approximated using (log) sales per worker (in Euro). Furthermore, we control for the structure of the workforce using the share of qualified workers (both blue and white collar) and the share of women, information on collective bargaining agreements and firm size. Unfortunately, the data does not contain information on a firm's capital stock thus preventing the use of production functions.

(TABLE 2 ABOUT HERE.)

Table 2 displays descriptive information on some key variables. Note that the high share of qualified workers is not uncommon in Germany where most workers have completed vocational training. The high number of firms covered by collective bargaining agreements at the industry level and the much lower number of firms with individual agreements is also typical for industrial relations in Germany.

4 Econometric modeling

Consider the following data-generating process

$$y_{it} = \gamma y_{it-1} + \beta X_{it} + \tau D_{it} + \eta_i + \epsilon_{it}, \quad (1)$$

where y_{it} is log sales per worker, γ captures state dependence of the dependent variable, X_{it} contains time-varying control variables, D_{it} is a matrix containing measures of wage inequality, η_i is a firm-specific fixed effect and ϵ_{it} an error term.

Equation (1) is used to estimate four models, each using different measures of inequality: Model I uses both the overall and the average within-group coefficient of variation. τ in this model measures the impact of changing one type of inequality while holding the other constant. Models II and III use only the overall or the average within-group coefficient of variation respectively, giving rise to unconditional estimates for each inequality measure. Finally, model IV uses the residual variation measure that controls for differences in workforce composition.

The firm-specific fixed effect η_i captures the fact that both wage dispersion as well as labor productivity may be influenced by time-invariant unobserved factors, e.g. firm “culture”. It also captures the impact of all time invariant variables, like industry affiliation, that might also influence productivity. The presence of these fixed effects implies that identification of all parameters uses variation over time within firms.

The state dependence parameter γ allows for the influence of past on present productivity. Many factors that influenced productivity in the last period, e.g. motivation of the workforce, can be expected to adjust only slowly, implying state dependence of the

productivity measure.

Finally, as already noted by Lallemand et al. (2004) there might be a simultaneity problem between productivity, especially when measured as sales, and wage dispersion: Consider a case where a firm experiences an unobserved productivity shock leading to higher profits. These in turn might influence intra-firm wage inequality through bonus payments, ultimately leading to contemporaneous correlation between wage dispersion and the error term from equation (1). This correlation in turn prevents identification of the causal impact of wage dispersion on productivity.

To address these three points equation (1) is estimated using the Arellano-Bond dynamic panel estimator (Arellano and Bond 1991).⁵ In a first step the firm-specific fixed effects η_i are eliminated by applying a forward orthogonal deviations transformation (Arellano and Bover 1995). Note that the usual within-transformation that subtracts the within-subject average of a variable from its current realization leads to inconsistent estimates in the presence of endogeneity because the transformed variable is related to all its current, future and past realizations – including those endogenous. The forward orthogonal deviations transformation is similar to the usual within-transformation but uses only the available future realizations of each variable for the transformation, thus leaving past realizations as valid instruments.

In a second step the simultaneity problem is addressed by using second and third lags of the right hand side variables as instruments. These are used alongside the strictly exogenous variables (basically a series of time dummies) to form a set of moment conditions that are used for GMM-estimation. As this yields more instruments than endogenous regressors, the equation is overidentified and difference-in-Sargan-tests can be used to test

⁵Using Stata SE 9.2 this involves the user-written command `xtabond2` (Roodman 2007), see Roodman 2006 for a description.

for the exogeneity of instruments. These cannot reject the Null hypothesis of exogeneity for almost all instruments with the lowest assorted p-values being 0.099 (Model I for *size*) 0.082 (Model II for *size*), 0.101 (Model III for *size*) and 0.062 (Model IV for *collective bargaining at the industry level*). One should note at this point that the estimation results indicate that size has only a minimal influence on productivity and collective bargaining agreements on the branch level can only be marginally influenced by a specific firm which should minimize potential problems with these instruments. Potentially harmful second-order autocorrelation does not seem to be an issue as the Arellano-Bond-Test for (no) autocorrelation (Arellano and Bond 1991) yields p-values of 0.309 (Model I), 0.542 (Model II), 0.325 (Model III) and 0.286 (Model IV). For benchmark purposes we also report estimates from standard within-estimators that assume strict exogeneity of all variables conditional on the firm-specific fixed effect η_i .

5 Results

Consider first the results for the inequality measures. In the standard fixed effects estimation we obtain significant results for overall wage inequality that suggest a negative (Model I) or U-shaped relationship (Model II) with firm returns. For average within group inequality the point estimates also suggest a U-shaped relationship in models I and III that is never significant on any conventional level. These results are basically reversed if we consider the (theoretically superior) Arellano-Bond-estimates: Here, we obtain point estimates that suggest the same hump-shaped relationship found by Winter-Ebmer and Zweimüller (1999) and Bingley and Eriksson (2001). Again, these results are not significant on any conventional level. For residual variation (Model IV) the results seem relatively clear: Regardless of the specific estimation procedure, all coefficients are close to zero and

consequently insignificant.

(TABLE 3 ABOUT HERE.)

While the preceding discussion focused on statistical significance, it is important to consider the question whether the estimated effects are also economically important, especially when some point estimates suggest huge changes in the dependent variable. This consideration is also important in the case of insignificant results where one has to distinguish insignificance due to large standard errors, that is imprecise estimates, from insignificance due to small coefficients, that is economic irrelevance. To put the effects into perspective, we consider the change in returns caused by an increase by one (within firm) standard deviation of the respective inequality measure.

In the case of overall wage inequality, an increase by 0.0688 causes only moderate changes in log sales. In the fixed effects regressions, sales drop by 0.016 (Model I) and 0.019 (Model II), while they increase by 0.024 (Model I) or remain practically unchanged (Model II) when looking at the Arellano-Bond-results. Similar modest changes are obtained when varying average group inequality. A rise by 0.0760 causes returns to drop by 0.003 (Model I) and 0.008 (Model III) using the fixed effects estimates and to rise by 0.017 (Model I) and 0.028 (Model III) using the Arellano-Bond-estimates. As already noted above, the effects using residual variation are also very small, with a one standard deviation increase leading to a rise in log returns of 0.002 (fixed effects) and 0.010 (Arellano-Bond). Taken together this evidence suggests that wage inequality does not seem to play a particularly prominent role in explaining differences in firm performance.

Now, looking at the parameters for the lagged dependent variable, we find our initial presumption that current productivity depends on its past values confirmed. The esti-

mated state dependence parameters are highly significant in Models I, III and IV and at least weakly significant in Model II. Looking at the marginal effects reveals that past productivity explains between roughly 13% (Model II) and 22% (Model III) of current productivity. Note that this may explain the relatively small impact of changes in wage inequality: If firm performance depends strongly on its own past, we might expect it to adjust only slowly to changes in the right hand side variables.

Finally taking a quick look at the control variables, one finds large differences between the fixed effect and Arellano-Bond-results. Consider first the results for collective bargaining agreements. Here we obtain mixed results in our estimation: First, the fixed effects estimates indicate a weak positive relationship between coverage by a collective bargaining agreement and firm performance. Secondly, the Arellano-Bond-estimates show rather large though mostly insignificant results. Note, however, that considerable care should be taken in the interpretation of these results since they are identified by firms switching either into or out of a collective bargaining agreement. Given the relative stability of industry level collective bargaining agreements, the high degree of coverage of German industries by such agreements and the rare occurrence of firm level agreements, these effects are most likely estimated using only a small subset of the firms in the sample and their importance should consequently not be overstated.

Regarding the effect of employment structure, one should note that firm size has a negative but negligible impact on productivity over all models and estimation methods. The impact of the share of qualified personal varies between estimation methods: Fixed effect results indicate a small, insignificant effect while the Arellano-Bond results suggest a positive relationship. The latter results are also economically important with a 10 percentage point increase in the share of qualified workers leading to a 3.5 to 4.5 percent increase

in sales. A possible explanation for the observed differences between the fixed effects and Arellano-Bond results might be the reactions of firms that experienced a negative shock in sales and hired more qualified personal as a countermeasure thus leading to a simultaneity problem in the fixed effects estimation. A similar argument could be found for the negative relationship between the share of women and firm performance in the fixed effects estimation: If women are less mobile than men and thus less likely to leave a firm that just experienced a negative shock, we would expect to find the same relationship as in table 3.

6 Conclusion

While the relationship between pay inequality and firm performance has received considerable attention in theoretical economics, empirical relationships are still far from clear. In particular, there has been few research combining inequality measures related to the whole wage distribution in a firm with firm level panel data outside of special environments like professional sports. This paper contributes to this literature by using linked employer-employee-data from Germany. We construct several firm level inequality measures using social security data and regress these on returns per head as a measure for productivity. Our estimation approach using dynamic panel estimators allows us to address issues like unobserved heterogeneity, simultaneity of wage inequality and performance as well as possible dynamics in the adjustment process of firm performance to changes in the wages structure.

Our results indicate a relatively small impact of wage inequality on firm performance. Furthermore, we find that the relationship between these variables tends to depend on the estimation method used. While fixed effects estimators show a negative or U-shaped rela-

tionship between some inequality measures and firm performance, results from Arellano-Bond estimators indicate a hump-shaped relationship that is always insignificant. The latter results also show that firm performance tends to exhibit a certain degree of state dependence, indicating a delay in the adjustment process to changes in the right hand side variables, which might explain the relatively small reaction to changes in the wage structure.

Regarding the different theories relating wage inequality and firm performance our results show relatively few support for either fairness or incentive based theories. If anything, the results show some weak support for the idea that some inequality may be beneficial, while an increase beyond a certain point may harm performance.

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TABLE 1: OVERVIEW OF PREVIOUS STUDIES

Study	Country	Data	Inequality measure	Outcome	Method	Relationship between inequality and outcome
Cowherd, Levine (1992)	USA, UK	cross-section of 102 business units	(i) wage difference between management and hourly paid workers (ii) wage difference between upper and lower management	index of product quality	OLS	negative for both measures and outcome
Winter-Ebmer, Zweimüller (1999)	Austria	social security records aggregated at firm level, panel of 130 firms 1975-1991	variance of residuals from wage equations	standardized Wages	within- and between-estimators	(i) hump-shaped for white collar workers (ii) positive, decreasing for blue-collar workers
Hibbs, Locking (2000)	Sweden	(i) industry level 1964-1993 (ii) plant level 1972-1993	between- and within coefficients of variation	(log) value added per worker	OLS	positive for within-, negative for between-inequality
Bingley, Eriksson (2001)	Denmark	LEE-data 22,665 firm-year-obs. for 1992-1995	variance of residuals from wage equation	(i) total factor productivity (ii) sickness absence	IV-estimation	(i) hump-shaped for white-collar workers and productivity (ii) none for blue-collar workers and productivity (iii) positive for both and sickness absence
Beaumont, Harris (2003)	UK	Plant level 1978-1995 panel	ratio of wages of manual to non-manual labor	gross value added per worker	Arellano-Bond	positive for most industries
Lallemand et al. (2004)	Belgium	LEE, 1995 cross-section	(i) variance of residuals from wage equation (ii) several unconditional measures	gross operating surplus	2SLS	positive, larger for blue-collar workers
Grund, Westergaard-Nielsen (2004)	Denmark	LEE, panel 1992-1997	coefficients of variation for (i) wages and (ii) wage increases	log value added	(i) OLS (ii) within-estimators	(i) hump-shaped for wages (ii) U-shaped for wage increases
Heyman (2005)	Sweden	LEE, 1991 and 1995 cross-sections	(i) variance of residuals from wage equations (ii) coefficient of variation (iii) 90/10-percentile ratio	(i) profits per employee (ii) log average wage (iii) variance of sales	first differences, IV-estimation	(i) positive for profits and average pay (ii) positive for managers and variance of sales
Jirjahn, Kraft (2007)	Germany	firm-level 1997 cross-section	wage difference between unskilled blue collar and skilled white collar worker	log value added	OLS	varies with industrial relations and pay scheme, insignificant without interactions

See text for detailed discussion.

TABLE 2: DESCRIPTIVE STATISTICS, ESTIMATION SAMPLE

Variable	Mean	Std.dev.		Minimum	Maximum
		(overall)	(within)		
Sales per head (€, 2000 prices)	201,484.4	1,025,115	0	9.14*10 ⁷	
Share of qualified workers	.6329	.2547	.1178	0	1
Share of women	.3096	.2592	.0543	0	1
Size (no. of employees)	279.2	793.61	75.75	2	20,957
Collective bargaining agreement industry level	.6463	.4781	.2098	0	1
Collective bargaining agreement firm level	.0749	.2632	.1553	0	1
Coefficient of variation (overall)	.4063	.1562	.0688	0	1.7878
Coefficient of variation (average over subgroups)	.2206	.1336	.0760	0	4.5616
Residual variation	19.47	6.70	3.65	.0764	214.88

No. of observations 20,078. Only observations with both coefficients of variation and residual variation not missing.

TABLE 3: PRODUCTIVITY REGRESSIONS, DEPENDENT VARIABLE: LOG RETURNS PER HEAD, WITHIN- AND ARELLANO-BOND-ESTIMATORS

Variable	Model I		Model II		Model III		Model IV	
	Fixed Effects	Arellano-Bond	Fixed Effects	Arellano-Bond	Fixed Effects	Arellano-Bond	Fixed Effects	Arellano-Bond
Inequality measures								
Coefficient of Variation (overall)	-0.2334** (0.1146)	0.3782 (0.4386)	-0.2955*** (0.0910)	0.0216 (0.3541)				
Coefficient of Variation (overall, squared)	0.0952 (0.1143)	-0.4800 (0.3939)	0.1756** (0.0837)	-0.2249 (0.3160)				
Coefficient of Variation (averaged over subgroups)	-0.0455 (0.0722)	0.2408 (0.3522)			-0.1118 (0.0687)	0.4006 (0.3755)		
Coefficient of Variation (average, squared)	0.0795 (0.0778)	-0.2165 (0.3631)			0.1003 (0.0755)	-0.4751 (0.4003)		
Residual variation (squared)							0.0005 (0.0008)	0.0028 (0.0069)
Residual variation (squared)							0.0000 (0.0000)	0.0000 (0.0001)
State dependence								
Lagged dependent variable (one year lag)		0.1938*** (0.0595)		0.1215* (0.0685)		0.1991*** (0.0644)		0.1690** (0.0715)
Control variables								
Collective bargaining (industry)	0.0130* (0.0101)	0.1083 (0.0681)	0.0176* (0.0097)	0.1020 (0.0836)	0.0132 (0.0101)	0.1032 (0.0788)	0.0180* (0.0097)	0.1942** (0.0881)
Collective bargaining (firm)	0.0218 (0.0130)	0.1261* (0.0737)	0.0232* (0.0124)	0.0629 (0.0873)	0.0222* (0.0131)	0.1346* (0.0792)	0.0237* (0.0124)	0.1269 (0.0882)
Share of qualified workers	0.0182 (0.0209)	0.3223*** (0.0999)	0.0210 (0.0200)	0.3268*** (0.1114)	0.0246 (0.0210)	0.3039*** (0.1075)	0.0264 (0.0201)	0.3712*** (0.1192)
Share of women	-0.0688 (0.0454)	0.1874 (0.2479)	-0.0820** (0.0415)	0.1653 (0.3148)	-0.0797* (0.0455)	0.1620 (0.2727)	-0.0975** (0.0417)	0.4787 (0.3053)
Size (no. of employees)	-0.0001** (0.0001)	-0.0001* (0.0001)	-0.0001** (0.0001)	-0.0002* (0.0001)	-0.0001* (0.0001)	-0.0001* (0.0001)	-0.0001** (0.0001)	-0.0002* (0.0001)
Constant	11.6277*** (0.0393)		11.5987*** (0.0360)		11.5689*** (0.0319)		11.5109*** (0.0310)	
Time effects (years)	(included)	(included)	(included)	(included)	(included)	(included)	(included)	(included)
No. Obs.	19,773	8,799	21,042	9,434	19,773	8,799	21,042	9,434
No. Firms	5,801	2,718	6,018	2,887	5,801	2,718	6,018	2,887
Sig. (Model)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000

Coefficients, robust standard errors in parentheses. ***/**/* denote significance on the 1%, 5% and 10% level respectively. Fixed effects are eliminated by within-transformation (Fixed Effects-Model) and forward orthogonal deviations (Arellano-Bond, see Arellano and Bover 1995). Arellano-Bond estimators use second and third lags as instruments, resulting in 179 (Model I) and 145 (Model II, III and IV) instruments. The Arellano-Bond-Test for autocorrelation (Arellano and Bond 1991) indicates no second-order autocorrelation with assorted p-values of 0.309 (Model I), 0.542 (Model II), 0.325 (Model III) and 0.286 (Model IV). Difference- in-Sargan-tests do not reject the Null of exogeneity for almost all instruments (lowest assorted p-value is 0.099 (Model I) 0.082 (Model II), 0.101 (Model III) and 0.062 (Model IV)). Lowest p-values for instruments of inequality measures are 0.538 (Model I), 0.573 (Model II), 0.449 (Model III) and 0.945 (Model IV).

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