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Janos Köllő† and Tomasz Mickiewicz‡
† Hungarian Academy of Science, IZA and WDI
‡ SSEES at University College London.
Senate House, Malet Street, London WC1E 7HU *T.Mickiewicz@ssees.ucl.ac.uk*

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University College London
Centre for the Study of Economic and Social Change in Europe
Senate House, Malet Street, London, WC1E 7HU
Tel: 44(020) 7863 8517
Fax :44(020) 7862 8641
Email: csesce@ssees.ac.uk

Wage Bargaining, Privatisation, Ability to Pay, and Outside Options – Evidence from Hungary[♣]

Janos Köllő† and Tomasz Mickiewicz‡

† Hungarian Academy of Science, IZA and WDI
‡ SSEES at University College London. Senate House, Malet Street, London WC1E 7HU
T.Mickiewicz@ssees.ucl.ac.uk

ABSTRACT
This paper examines the determinants of short-term wage dynamics, using a sample of large Hungarian companies for the period of 1996-1999. We test the basic implications of an efficient contract model of bargaining between the incumbent employees and the managers, which we are unable to reject. In particular, there are structural differences between the ownership sectors consistent with our prior knowledge on relative bargaining strength and unionisation measures. Stronger bargaining position of workers leads to higher ability to pay elasticity of wages, and lower outside option elasticity. Our results indicate that while bargaining position of workers in domestic privatised firms may be weaker than in the state sector, the more robust difference relate to state sector workers versus the privatised firms with the majority foreign ownership.
We examine several extensions. We augment the bargaining specification by controls related to workers' skills and find that the basic findings are robust to that. We take a closer look at the outside options of the workers. We find some interactive effects, where unemployment modify the impact of availability of rents on wages. We interpret our results as an indication that bargaining power of workers may be affected by changes in their outside options. We also experiment with one concise indicator of reservation wage which is closest to the theoretical model specification and combines sectoral wages, unemployment benefits and regional unemployment levels,. We found that measure performing well.
Finally, we found that while responsiveness of wages towards ability to pay is higher in the state sector, variation in wage dynamics is lower. This may indicate some wage smoothing in the state sector, consistent with the preferences of employees.

Keywords: wages, bargaining, unemployment, privatisation, foreign ownership, Hungary

JEL Classification: D21, J30, L32, P31

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Introduction

Wage inflexibility and rent sharing are potentially important explanations of why employment levels fail to recover in post-communist economies, and why unemployment rates stabilised at double-digit levels in many of them. Wage rigidity may have also contributed to the survival of regional inequalities paramount by western standards.

However, researcher's knowledge of how wages are actually set in these countries is rather poor. The characterisations of the institutional setup as 'centralised' versus 'decentralised' or 'coordinated' versus 'uncoordinated' are inevitably arbitrary since it is difficult to assess the practical importance of the existing institutions. In the country under examination, for instance, Labour Force Survey data from 2001 suggested that 22% of the employees were union members, 24% earned a wage influenced by collective agreements, 41% was employed at unionised firm, over 95% worked in a two-digit industry where at least one firm (potential wage leader) concluded collective wage agreement, and 100% was subject to minimum wage regulation and addressee of a national tripartite agreement on the 'desirable' rate of wage growth. Which of these figures bear relevance for wage determination is an open empirical question that can be best understood by studying actual wage evolutions on the micro level.[1]

In this paper we analyse a panel of Hungarian firms applying a bargaining framework where wages respond to changes in both ability to pay, outside options and, potentially, bargaining power. We take sales per employee as a proxy of ability to pay. The influence of regional and industrial factors are observed through the responses of wages to worker's outside options. Bargaining power is related to the distinction between firms with majority ownership of the state, private domestic owners and foreign owners.

The main contribution of our paper is the following. First, we test the implications of the bargaining framework controlling for skills and experience characteristics, and demonstrate that the basic implications are robust to the augmentation. Second, the results shed some light on the implications of the

[1] Authors' calculations using the 2001 April-June wave of the LFS. It might worth mentioning that the data are themselves of questionable precision. Union coverage in the mid-1990s was estimated to be in

privatisation process for rent sharing. In particular, privatisation to foreign strategic investors induce a stronger attenuating effect on rent sharing than privatisation to domestic owners. Third, while rent sharing is more pronounced in the state sector, we also find indication of some wage smoothing there. Forth, we use a measure of regional unemployment, which takes into account that firm' employment may be split between several regions. Arguably, the measure applied has lower measurement error an that is why we are able to detect a strong wage-curve type effect. Fifth, we discover that regional unemployment may also modify the ability to pay elasticity of wages, that is the inside and outside variables interact.

Section 2 motivates the basic model of wage bargaining we refer to. Next, in section 3 we discuss proxies and indicators for variables. Section 4 describes data. Subsequently we focus on interference, presenting our specifications and results in section 5. Section 6 concludes.

2. Model of wage bargaining

We start with a brief presentation of a theoretical model that may motivate the intuition of the empirical specification we wish to test. The two main categories of bargaining models relate to (i) 'right to manage' and (ii) 'efficient contract' frameworks.[2] The difference relates to the fact that in the latter case the bargaining process may include employment while in the first case, the managers determine employment unilaterally after wage decisions are taken. Therefore, in the first case, the resulting wage and employment combination is always placed on the labour demand (marginal revenue product of labour) curve. In the second case, they may be off labour demand curve, as simultaneous bargaining over two variables extends the possible range of solutions.

However, the empirical difference between the predictions of the 'right to manage' and 'efficient contract' models do not relate to interference on wage levels, but rather on employment. Many models in the "efficient contract" category assume high weight attached to employment in the objective function of the risk-averse

the range of 20-30% on the basis of firm and worker surveys (Neumann 1997) while some often-quoted sources like ILO (1997) and OECD (1999) suggested 60% level, see Cazes and Nesporova (2003).

[2] One may also notice that the insider control model can be interpreted as a special case of the efficient contract model.

employees and therefore predict contract curves, on which employment levels are higher than those resulting from the competitive equilibrium. In contrast, in the 'right to manage' situation, the bargaining solution is always to the left of the maximum profit (competitive) solution, while still on the marginal revenue product curve. Therefore in the latter case the employment is lower.

As we do wish to focus on wages without making predictions about the employment outcomes, the model of 'efficient contract' with risk neutral incumbent workers bargaining with managers of the firm appeals to us. This is also motivated by the patterns of unionisation in the country under examination. Union density is relatively low with 8.2 per cent of the workers being union member in small firms (less than 50 workers) and 23.8 per cent in large firms. Even in unionised large firms members account for only half of the employees.[3] Practices characteristic of unions that maximise the welfare of a *fixed* membership, such as restrictions put on hiring workers other than those previously laid off from the firm, are largely missing. Hungarian unions apparently seem uninterested in several issues relevant for employment-aware bargaining such as import policies, customs duties, or immigration legislation. Given these features, the assumption of bargaining between the firm and a small group of insiders (Carruth and Oswald 1985) seems to fit better than the presumption of employment-aware unions (McDonald and Solow 1982).

In view of this, we adopt the model, in which the contract curve is vertical, thus employment remains equivalent to the competitive solution. The empirical appeal of such a model results from the fact, that – unlike models assuming solutions along the demand for labour curve – the increase in bargaining position of the incumbent workers leads to higher wages, but does not affect the employment level negatively. The highest attainable wage corresponds to zero profits. Lowest wage is equivalent to alternative wage and corresponds to the profits, which would result from the (short term) competitive equilibrium. The bargaining is depicted by Figure 1 below.

(Figure 1)

Following Svejnar (1986) we write the generalised Nash bargaining as:

[3] Hungarian Labour Force Survey April-June 2002.

$$U_L^{\beta} U_L^{1-\beta} = \left[\frac{L}{\delta \overline{L}}(w^{\delta} - w_a^{\delta})\right]^{\beta} \pi^{1-\beta} = \left[\frac{L}{\delta \overline{L}}(w^{\delta} - w_a^{\delta})\right]^{\beta} [R - wL - H]^{1-\beta} \tag{1}$$

where:

w is total wage (labour cost) derived by the worker from his/her employment,

w_a corresponds to the alternative wage,

$\overline{L}$ represents incumbent employment,

L relates to employment secured as a result of the (implicit) bargain,

δ is a measure of risk aversion (with higher value implying more risk loving),

π represents profits,

R is total revenue,

H relates to non-labour costs of production, and

γ represents the bargaining power of the incumbent workers, where $0 \leq \gamma \leq 1$.

Assuming risk neutrality ($\delta = 1$), the contract curve corresponding to this problem on the employment-wage plane is vertical, that is employment level is equivalent to the profit maximising (competitive) equilibrium ($L=L^*$) while wages exceed the opportunity cost level in proportion to the incumbent workers bargaining power. The solution reduces to the following condition:

$$w = w_a + \gamma \frac{\pi^*}{L^*} = (1-\gamma) w_a + \gamma \left(\frac{R}{L^*} - \frac{H}{L^*} \right) \tag{2}$$

where π^* represents non-zero profits evaluated at the (short term) competitive solution.

Thus, we can immediately derive the following implications from the model:

/i/ higher bargaining strength is associated with higher responsiveness of wages to the firm's ability to pay and lower responsiveness to outside options;

/ii/ increase in the reservation wage of the employees will lead to the increase in wages ($\frac{\partial w^*}{\partial w_a} = 1 - \gamma + \frac{R}{L^*} - \frac{H}{L^*} > 0$), as long as the firm produces positive value added;

/iii/ with stable employment, positive external shocks to profits (positive to revenues, negative to non-wage costs) will lead to the increase in wages, as long as the workers have some bargaining strength ($\frac{\partial w}{\partial \left(\frac{R}{L^*} \right)} = \gamma \geq 0$).

We will focus on those three implications in the empirical section, yet without attempting to estimate the exact structural form.[4]

3. Proxies

3.1. Ability to pay

As discussed, following equation (2), we expect wages to respond positively to alternative wages, profits and bargaining power of incumbent workers. Yet the issue of empirical proxies is not trivial.

A number of studies use profits per employee as a proxy for quasi rent (Fakhfakh and Fitzroy 2002, Hildreth and Oswald 1997). Yet, there are problems since profits are clearly endogeneus.[5] In particular, in line with the bargaining model, wage is a function of profit estimated at the competitive equilibrum solution (π^*), not of realised profit after wage cost is paid (π). As value added is distributed between profit and wages, that may lead to negative correlation and create an attenuation bias when profit is taken as determinant of wages with assumed positive sign. The problem could be alleviated by use of instruments, and this is the approach followed by the authors quoted above, but that brings in different estimation problems – as always, reliable exogeneous variables which may affect profits on individual firm level are difficult to find.[6] Third, profits are volatile – they vary significantly from one year to another. The current ability to pay is in practice determined by retained earnings accumulated over several years. The above authors were able to use data sets, which span over long time dimension and could control for several lagged values of profits. Such dataset are rarely available for transition countries.

For that reason, revenue per employee may be used as a proxy for ability to pay, which is still consistent with equation (2). This variable was utilised in seminal paper by Nickell and Wadhwani (1990) and applied in the transition economies context by Grosfeld and Nivet (1999), Basu *et al.* (2000), Christev and Fitzroy (2002) and Mickiewicz and Bishop (2003). In particular, when the specification is augmented by sectoral wages, the difference between revenue per employee and prevailing sector

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[4] In particular we will follow typical design and approximate the model by log-linear specification.
[5] See Van Reenen (1996).
[6] One possibility is to use sectoral level data on profits and demand shocks, provided the relevant data is available. This is the approach taken by Abowd and Lemieux (1993) and Christofides and Oswald (1992).

wage may be treated as a very good indicator of available quasi-rent (Van Reenen 1996). Revenue per head might be also interpreted as labour productivity. Yet, unlike productivity measures based on some production function specifications, this is a very imperfect indicator of labour productivity. Even if we label the variable as "productivity", it is still a very good indicator of availability of quasi rents, as argued above. In the context of productivity it is important however to control for the skill composition of workers. This will be addressed in the empirical section.

3.2. Outside options

The indicators of outside options, which appear often in the estimation of wage equations, are regional unemployment and outside wages.

The negative relationship between regional unemployment rates and wages is often interpreted in terms of the wage curve, which relates to the cross-sectional relationship between the level of wages and the level of unemployment (Blanchflower and Oswald 1995). The empirical results confirming microeconomic wage curve are common in studies of European transition economies, but empirical specifications differ. We may notice first that the standard interpretation of the basic wage bargaining model implies regressing change in wages against change in unemployment, if the second is taken as an indicator of outside options (typically in natural logarithms). This amounts to first-differencing from wage curve, thus the derived specifications are parallel. Thus, bargaining model is one possible way to provide theoretical justification for the wage curve.

However, the research on the cross-sectional link between unemployment and wages is driven by empirics and specifications differ. In particular, Blanchflower (1990) found that four alternative measures of unemployment and employment – some in levels, and some in first differences are negatively and significantly related to annual earnings. Grosfeld and Nivet (1999) regress first difference of wages against – alternatively – both the level of regional unemployment and the first difference in regional unemployment (for Poland), while Mickiewicz and Bishop (2003) use the latter specification. Duffy and Walsh (2001) and Kertesi and Köllő (1999) (for Hungary) apply directly the wage curve (both variables in levels), while Christev and FitxRoy (2002) regress first difference in wages against unemployment level (both papers for Poland) similar to earlier papers by Christofides and Oswald (1992) (for the

UK) and Holmund and Zettenberg (1991) (five OECD countries). Holmund and Zetterberg (1991) also hypothesise that unemployment (they use an economy wide unemployment rate) is likely to slow down wage growth,[7] in their study of the determinants of industry wages in five countries. Yet, their results show that the effects of aggregate unemployment vary across countries: negative as expected for Sweden, Finland and Germany, yet positive for Norway and the USA. The result is interesting as it possibly reflects differences in institutional labour market characteristics implying that wage curve may be specific to the labour market institutions. In particular, the positive coefficient for the USA may imply a more competitive labour market, which can be interpreted along the lines of the 'first generation' models, where wages may compensate for higher unemployment risk. Following this line of argument, one may notice that evolving institutional frameworks in transition countries make testing wage curves for transitional economies a non-trivial task.

One should also note that the link between wages and unemployment can be interpreted not only in terms of bargaining theory but also in terms of efficiency wage theory, where wages do not result from bargaining process but from optimising decisions by the firms (Shapiro and Stiglitz, 1984).

In addition, as just mentioned, there is a dissenting tradition of the neo-classical or "first generation" of papers by researchers such as Harris- Todaro (1970), Hall (1970) and Rosen (1986) predicting that unemployment and wages would move in the same direction. This relies on the perfectly competitive theory and compensating differentials. Wages may have to compensate for job characteristics, location, flexibility, risk to health etc. Duffy and Walsh (2001) provide a brief survey of the "first generation of papers" written in the 1970's and 1980's, which all found a positive relationship between wages and unemployment. However, they criticise this line of research for failing to control for regional fixed effects. They argue that after including regional dummies, the relationship between regional pay and unemployment are in fact negatively correlated (*Ibid*., p.25). The evidence is still not conclusive. A recent study by Cahuc *et al*. (2002) on a panel of French firms finds some new evidence that confirms the predictions of equalising differences, as unions accept

[7] However, they note that the depressing effect of higher unemployment is likely to be bigger in an aggregate wage equation.

lower aggregate wages when workers benefit from lower unemployment risk. As argued above, the parameters of the wage curve are conditional on labour market institutions (competitiveness in particular), and may be therefore neither cross-country nor time invariant, thus still worth further testing.

Less controversy relates to the use of alternative wage as an indicator of outside options. One may notice that the latter link should be more relevant, if the likelihood of re-entering employment for those workers, who may loose their job is high. Thus, outside wages and regional unemployment may interact in their effect on wage dynamics. High unemployment/vacancy ratio, low turnover in the job market and low outflow into jobs from unemployment, would diminish the importance of alternative wages, and the level of employment benefits would count more. For studies based on one country, unemployment benefits are typically uniform. But the likelihood of entering a new job may be negatively correlated with regional unemployment. If so, one would expect the latter to be significant component of outside opportunities.

3.3. Bargaining strength

Finally, we consider proxies of bargaining strength. In line with the empirical papers exemplified above (in particular: Grosefeld and Nivet (1999), Christev and FitzRoy (2002) and Mickiewicz and Bishop (2003)), we link bargaining to ownership characteristics. In particular we hypothesise the stronger position of incumbent workers in the state sector versus the private sector. This intuition is supported by results of earlier research on Hungarian labour market by one of the authors. The results are based on data from the 1998 Wage Survey augmented with data on collective agreements. In that year the probability that a worker was covered by a collective wage agreement was lower by 8.2% in case of mixed ownership, 18% in case of private domestic owners, and 34.5% in case of foreign ownership compared to state-owned firms, after controlling for firm size. (based on logit model, with firms in the budgetary public sector excluded, marginal effects, N=103,561, pseudo r2=0.356).[8] This allows direct interpretation based on institutional characteristics of bargaining (not necessary explicit). Interestingly, it is supported by another dimension, which is directly linked to outside options. This second piece of information comes from the single data set on severance pay – 1994, workers losing jobs and becoming unemployment insurance recipients in April 1994. The probability that a state sector worker received severance pay after controlling for tenure and a manual/non-manual dummy was 11% higher in fully state-owned firms and 16% higher in partly state-owned firms compared to private firms including foreign ones (logit, marginal effects, N=5075, pseudo r2=0.2).

4. Data

4.1 Data description

As argued by Hamermesh (1993) firm level data may be superior to household data for studying the firm-specific issues such as rent sharing. Our sample of large firms is drawn from the National Labour Centre's Wage Survey (WS), which is a matched employer-employee database. The surveys were carried out in May 1986 and 1989 and have been conducted each May since 1992. It contains data of about 150,000 workers employed in 6,000 to 12,000 firms, depending on year.

The sampling procedure is two-step. At the first step *firms* are selected, while at the second, a random sample of full-time *employees* is drawn within each firm. The table below summarises the variations in the sampling procedures by sector and firm size.

Table 1. Sampling procedures for the National Labour Centre Wage Survey

Category	Selection of firms	Selection of employees	Notes
Budget institutions	100%	100%	Armed forces excluded
Firm > 20 employees	100%	about 10%	1986-
Firms with 11-20 employees	about 12%	100%	1995-
Firms 5-10 employees	about 12%	100%	1999-

The sub-sample for this panel was drawn by selecting firms, which reported at least 30 individual observations (representing roughly 300 employees) in all years between 1996 and 1999. The firm-level data on the level of employment and nominal

[8] Turnover statistics broken down by ownership are not available.

variables were taken from the firms' Financial Reports delivered to the tax authorities. All data refer to annual flows or annual average stocks. The PPI indexes were drawn from National Accounts on the 4, 3, 2 or 1 digit level whichever was available (32 distinct values). The unemployment rate attached to the firm is the weighted mean of the micro-region-level unemployment rates given the location of the firm's branches. The ownership variable is based on shares in equity as reported in the Financial Reports. The industry-level variables are calculated using the data of all firms observed in the WS. Since the firm-level response rate is close to 100 per cent in the size category considered in this paper we did not use weights to correct for occasional non-response.

4.2 Descriptive statistics

Table 2 below presents a description of all available variables.

Table 2. Description of variables

l_rwage: logarithm of real wage (deflated using a corresponding sectoral ppi)
ebt_rev: earnings before taxes (but net of financial costs) divided by sale revenue
rtrev: real total revenue from sales
rtremp: real total revenue from sales per employee
rswage: real sectoral wage
unsu_n: ILO-methodology survey based regional unemployment (constructed as a weighted average of unemployment rates is case of few places of operation)
reswage: reservation wage, constructed as described in section 5.1 below,
Experience and education:
exp_old_young = No of old educated / No of young educated,
ed_low_high = No of low educated / No of high educated
exp_old_tot = No of old educated / (No of old educated + young educated))
lowedu = No of low educated / (No of low educated + high educated),
where 'old' means age above population mean, and 'young' below mean, and 'high education' stands for secondary and higher education.
Dummies:
small = lowest one third of firms ranked by employment at the beginning of the period,
large= highest one third of firms ranked by employment at the beginning of the period,
state = majority state ownership,
mixed = state, private and foreign shares are all<50 per cent
dom_maj = majority private domestic ownership,
for_maj = majority foreign ownership,
year97, year98, year 99 = annual dummies,
Sectoral controls constructed as a set of orthogonal contrasts between:
serv_ind = services versus industry,
trade_ser = trade versus other services,
minh_ind = mining & heavy industry versus other industry,
util_ind = utilities versus other industry,
cons_ind = construction versus other industry,
eng_man = engineering versus other manufacturing,
chem_man = chemical industry versus other manufacturing.

Note: The following prefixes will be used: *d_* denotes annual change, *l_* denotes logarithm, *dl_* denotes logarithmic difference, *dp_* relates to percentage change (applied where logarithms cannot be directly applied due to variables with some negative values, like profit)

Median values for selected variables and for basic ownership and sectoral cross-sections of firms in the data set are presented in Table 3 below. Reported significance levels relate to non-parametric tests on the equality of medians.

Table 3. Median values for selected variables over 1996-1999

Category	log change in real wage *dl_rwage*	earnings before taxes/sales *ebt_rev*	change in (EBT / sales) *d_ebt_rev*	log change in real sales *dl_rtrev*	log change in (real sale/empl) *dl_rtremp*	% change in (low ed/ total) *dp_lowedu*	% change in (old ed/ total edu) *dp_exp_old*
All firms	.060 (*1279*)	.134 (*1796*)	.001 (*1279*)	.036 (*1279*)	.055 (*1279*)	-.011 *(1033)*	-.000 *(1035)*
State majority ownership	.049 (*306*)	.216*** (*452*)	.002 (*306*)	.020** (*306*)	.037** (*306*)	-.009 (*269*)	.006 (*269*)
Mixed	.010** *(38)*	.126 (*63*)	-.002 (*38*)	.029 (*38*)	.055 (*38*)	-.026 (*34*)	.035 (*34*)
Private majority ownership	.062 (*508*)	.135 (*687*)	.002 (*508*)	.016* (*508*)	.060 (*508*)	-.005* (*392*)	.001 (*392*)
Foreign majority ownership	.070† (*427*)	.095*** (*594*)	.000* (*427*)	.086*** (*427*)	.076† (*.427*)	-.026** (*338*)	-.023† (*340*)
Industry	.057 (*946*)	.128*** (*1327*)	.001 (*946*)	.024*** (*946*)	.047 (*946*)	-.013 (*759*)	.000 (*759*)
Services	.067 (*329*)	.164*** (*463*)	.000 (*329*)	.073*** (*329*)	.070 (*329*)	-.003 (*270*)	-.007 (*274*)

Notes:
(i) Number of observations in each category is given in brackets. The growth rates of the variables were trimmed so that outlier observations in the tails of each variable were removed (0.5% on both ends, i.e. 1% in case of each variable). That relates to all subsequent estimations. Results on data with outliers (N=1323 in case of first differences, as compared with N=1279 here) are available on request.
(ii) *** Significant at .001; ** Significant at .01; **** Significant at .05; † Significant at .1
(iii) Significance levels relate to Pearson χ^2 (continuity corrected) based on the non-parametric test on the equality of medians. Fisher's exact test (two-sided) produces very similar significance results (not reported).

Several conclusions follow immediately from Table 2. Wage growth seems to be very similar across the ownership and sectoral cross-sections, apart from mixed ownership, where it is lower, but this category contain a very small number of observations. More importantly, the wages are growing faster in the foreign sector as compared with the rest of the sample. The result can be linked to better performance of foreign firms in terms of growth of both sale revenues and sales per employee (but not in terms of profits). Interestingly, the foreign controlled companies are also changing the skill composition of their workforce reducing the share of low-educated workers; the effect is highly significant as compared with other ownership sectors. Obviously this is an important complementary explanation of stronger wage growth, which should be controlled for in multivariate settings.

State firms are characterised by significantly higher profitability, but on the other hand, the performance of state sector is worse if measured by both the growth of sales and the dynamics of the sales per employee, albeit the last difference is not significant. The combination of those two characteristics may suggest some static rents resulting from market power.

Also, the comparison between the industry and service sector is showing no differences in wage increase, but clear differences in performance indicators. The service sector is performing better in terms of both profitability and dynamics of revenues (but the difference in the dynamics of revenue per head is not significant). The underlying tertiarisation process and the initial underdevelopment of services inherited from the command economy period may suggest higher growth opportunities in the latter period.

Last but not least, we looked into the distribution pattern of the key variables. The most interesting case relates to the pattern of wage dynamics, once it is split between the state and the private sector. Namely, standard deviation of wage dynamics is far lower in the state sector than in the private sector. The corresponding histograms are presented as Figure 5 below. We will return to this result later.

(Figure 2)

5. Interference

5.1 Methods and specifications

The panel we have at our disposal has a very short time dimension. That renders any attempt at dynamic specification difficult to justify. For that reason our estimation strategy relies on transforming all variables into natural logarithms and applying 'within' panel estimation (fixed effects model). The model seems to have most natural interpretation in terms of the comparative statics of equation (2), which were discussed above. In addition, the Hausman test rejected a potentially more efficient GLS random effects estimator as inconsistent (for specification (1) in Table 4 below, the test renders $\chi^2(6)=189.18$, which is highly significant). A possible criticism of the use of the fixed effects estimator for wage equations based on regional data is that it does not take into account the potential endogeneity of unemployment. While the issue has been raised in the literature, Bell *et al.* (2000) argue that the problem is unlikely to be serious due to “the high degree of persistence in labour demand and the notoriously sluggish response of unemployment to shocks of any kind” (*Ibid.* p.9). Moreover, in case of estimation based on individual company level wages, where regional unemployment is included on the right hand size, the problem is alleviated even further, even if we allow for some impact of large companies on regional labour markets.

Following earlier discussion, in the benchmark specification we regress wages on revenue per employee, sectoral wages and regional unemployment (all controlling for time effects and individual fixed effects). Subsequently, we attempt to see if the model is robust to alternative specifications.

First, we modify the proxies for alternative options, combining unemployment and outside wages in one indicator of reservation wage:

$W_r = (1\text{-}U)*W_s + Ub$,

where *b* relates to replacement ratio, *Ws* is sectoral wage, and *U* is the local unemployment rate.[9] Finally, we introduce interactive effects of unemployment with the ability to pay measure (revenue per employee).

The next set of tests relates to interactive effects based on ownership sectors. We look at the differences to see if they are consistent with our prior expectations related to bargaining strength.

Finally, we apply tree quantile regressions, for medians, first and third quartile of wage growth. While the fixed effect model coefficients may be interpreted as short-term effects around the individual means, the median regression offer an additional test of robustness of our results. Here, to account for individual effects, we first difference all the variables. In addition, the two quartile regressions offer an opportunity to test directly if the characteristics of response differ along the distribution of the dependent variable. That is interesting to investigate, given the fact that the distribution of wage dynamics variable in the state sector is more compressed than in the private sector (see Figure 2 in the descriptive section above). More specifically, we are able to test if estimated coefficients of explanatory variables for wage dynamics are different for firms characterised by high wage growth from those

[9] In time of our work of this paper we became familiar with Dobbelaere (2004) who applied similar approach independently (her sample being drawn from Bulgaria 1997-1998).

for firms with low or negative wage growth and if the ownership differences matter in this context.

5.2 Results

All estimation results are presented in Tables 4-6.

Table 4. Fixed effects (within) models. Dependent variable: natural logarithm of real wage (l_rwage). Benchmark specifications: (1)-(3); controlling for education and experience: (4)-(5).

	(1)	(2)	(3)	(4)	(5)
ln (real sales/employment) (l_rtremp)	.217*** (.013)	.220*** (.013)	.276*** (.048)	.205*** (.013)	.204*** (.013)
ln real sectoral wage (l_rswage)	.148*** (.020)	-	.149*** (.020)	.147*** (.020)	.144*** (.020)
ln regional unemployment (l_unsu_n)	-.120*** (.033)	-	-.156*** (.044)	-.136*** (.036)	-.134*** (.036)
"reservation wage" construct (l_reswage)	-	.151*** (.020)	-	-	-
ln (unemployment times sales/employment) (lrem_x_lun)	-	-	.021 (.017)	-	-
ratio of low educated to high educated workers (ed_low_high)	-	-	-	-.002 (.002)	-
ratio of experienced high educated to unexperienced high educated workers (exp_old_young)	-	-	-	-.002 (.002)	-
share of low educated in total no of workers (lowedu)	-	-	-	-	-.001† (001)
share of experienced in total no of workers (exp_old_tot)	-	-	-	-	-.044 (.040)
year97	-.002 (.007)	.001 (.007)	-.002 (.007)	-.009 (.007)	-.010 (.007)
year98	-.021 (.010)	.007 (.008)	-.013 (.010)	-.020 (.011)	-.019 (.011)
year99	.072*** (.011)	.090*** (010)	.071*** (.011)	.062*** (.012)	.061*** (.012)
constant	-.647*** (.085)	-.335*** (.020)	.746*** (.115)	-.652*** (.093)	-.566*** (.101)
Null: all individual firm effects = zero (F statistics for joint significance)	24.14***	25.90***	24.11***	22.93***	19.11***
R^2	.49	.49	.49	.50	.50
No of firms	492	492	492	492	492
No of observations	1796	1796	1796	1608	1611

Notes:

(i) Standard errors in parentheses.

(ii) *** Significant at .001; ** Significant at .01; **** Significant at .05; † Significant at .1

Table 5. Fixed effects (within) models. Dependent variable: dl_rwage. Interactive ownership effects.

	(1)	(2)	(3)	(4)
ln (real sales/employment) (l_rtremp)	.245*** (.019)	.248*** (.019)	.250*** (.019)	.252*** (.019)
ln (real sales/employment) x priv. domestic dummy (l_rtremp_do)	-.010 (.014)	-.011 (.015)	-.014 (.015)	-.018 (.015)
ln (real sales/employment) x foreign dummy (l_rtremp_fo)	-.051*** (.016)	-.055*** (.016)	-.057*** (.016)	-.053*** (.016)
ln real sectoral wage (l_rswage)	.141*** (.020)	.140*** (.020)	.093** (.035)	.116** (.037)
ln real sectoral wage x priv. domestic dummy (l_rswage_do)			.040 (.034)	.017 (.035)
ln real sectoral wage x foreign dummy (l_rswage_fo)			.068† (.036)	.041 (.039)
ln regional unemployment (l_unsu_n)	-.136*** (.036)	-.076 (.046)	-.087† (.047)	-.093* (.047)
ln regional unemployment x priv. domestic dummy (l_unsu_n_do)		-.036 (.030)	-.026 (.031)	-.016 (.032)
ln regional unemployment x foreign dummy (l_unsu_n_fo)		-.083* (.039)	-.057 (.041)	-.046 (.042)
share of low educated in total no of workers (lowedu)	-.001† (.001)	-.001† (.001)	-.001† (.001)	.000 (.001)
share of low educated x priv. domestic dummy (lowedu_do)				-.001† (.001)
share of low educated x foreign dummy (lowedu_fo)				-.002† (.001)
private domestic majority ownership dummy (dom_maj)	.027 (.025)	-.063 (.081)	-.034 (.085)	.080 .108
foreign majority ownership dummy (for_maj)	.128*** (.038)	-.078 (.104)	-.018 (.109)	.083 (.122)
year97	-.011 (.007)	-.011 (.007)	-.012 (.007)	-.012 (.007)
year98	-.023* (.011)	-.020† (.011)	-.019† (.011)	-.020† (.011)
year99	.057*** (.012)	.060*** (012)	.061*** (.012)	.060*** (.012)
constant	-.666*** (.102)	-.523*** (.125)	-.553*** (.126)	-.635*** (.133)
Null: all individual firm effects = zero (F statistics for joint significance)	18.67***	18.69***	18.71***	18.25***
R^2	.51	.51	.51	.51
No of firms	492	492	492	492
No of observations	1612	1612	1612	1612

Notes:
(i) Standard errors in parantheses.
(ii) *** Significant at .001; ** Significant at .01; **** Significant at .05; † Significant at .1

Table 6. Quantile regressions. Dependent variable: dl_rwage

	(1) q25	(2) q50	(3) q75	(4) q75-q25	(5) q50-q25	(6) q75-q50
dl_rtremp	.194*** (.023)	.245*** (.030)	.264*** (.037)	.069† (.038)	.051* (.024)	.019* (.032)
dl_rtremp_st	.158** (.058)	.108* (.043)	.069 (.064)	-.090 (.064)	-.050 (.041)	-.039 (.045)
dl_rswage	.087** (.027)	.081*** (.023)	.113*** (.021)	.026 (.027)	-.006 (.022)	.032 (.020)
dl_unsu_n	-.130** (.046)	-.040 (.033)	-.026 (.044)	.105* (.047)	.091 (.037)	.014 (.039)
dlrev_x_dlun	.384** (.139)	.467* (.184)	.593* (.232)	.209 (.198)	.083 (.142)	.126 (.126)
small	.003 (.007)	.006 (.007)	.013† (.007)	.010 (.008)	.004 (.006)	.007 (.006)
large	.005 (.008)	.006 (.006)	.012 (.007)	.006 (.008)	.001 (.006)	.005 (.006)
state	.018* (.008)	.003 (.006)	-.016† (.008)	-.034** (.011)	-.015† (.008)	-.019* (.008)
for_maj	-.002 (.008)	.004 (.007)	.003 (.009)	.005 (011)	.006 (.007)	-.001 (.007)
year98	.007 (.011)	.010 (.010)	-.007 (.009)	-.014 (.011)	.003 (.010)	-.017 (.011)
year99	.078*** (.010)	.070*** (.008)	.069*** (.011)	-.009 (.012)	-.009 (.008)	-.001 (.009)
serv_ind	-.001 (.002)	-.002 (.002)	-.005† (.003)	-.004 (.003)	-.002 (.002)	-.002 (.003)
trade_ser	-.014* (.006)	-.012* (.005)	-.000* (.008)	.013 (.010)	.002 (.006)	.011 (.007)
minh_ind	.004 (.004)	.004 (.004)	.002 (.004)	-.002 (.005)	.001 (.003)	-.003 (.004)
util_ind	-.004* (.002)	-.005*** (.001)	-.007*** (.002)	-.003 (.002)	-.001 (.002)	-.002 (.002)
cons_ind	-.000 (.003)	.002 (.003)	.003 (.006)	.003 (.005)	.002 (.003)	.001 (.005)
eng_man	.012** (.004)	.007* (.003)	.005 (.003)	-.007† (.004)	-.005 (.003)	-.002 (.004)
chem_man	.006 (.005)	.001 (.003)	-.000 (.004)	.006 (.005)	-.005 (.004)	-.001 (.004)
constant	-.052*** (.009)	.001 (.007)	.060*** (.009)	.122*** (.010)	.054*** (.007)	.058*** (.008)
Pseudo R^2	.20	.22	.21	-	-	-
N of observ.	1275	1275	1275	1275	1275	1275

Notes:
(i) Number of bootstrap replications: 100.
(ii) Bootstrap standard errors in parantheses.
(iii) *** Significant at .001; ** Significant at .01; **** Significant at .05; † Significant at .1
'Smaller firms' (column 4) refer to the bottom 1/3 of the sample when ordered by size at the beginning of the sample period. Correspondingly, 'larger firms' (column 5) refer to the top 1/3 of the sample.

5.3. Discussion

Clearly, wages seem to respond to the measures of ability to pay, i.e. to sales per employee. The estimates of corresponding aggregate elasticities vary between 0.20 and 0.28 depending on specification (Table 4). They are similar to those found for Poland by other researchers. Comparing with previous results on Poland, we may see, that Grosfeld and Nivet (1999) reported sales per employee elasticity of wage at 0.14 for early transition period and Mickiewicz and Bishop (2003) at 0.23. However, Christov and FitzRoy (2002) fiund higher elasticities for Poland, at 0.60-0.62 for more recent period.

Turning to alternative wage, we may see that the sectoral wage is consistently significant (with elasticity estimates in a range of 0.14-0.15), and so is unemployment (elasticity between –0.12-0.14). Combining both variables into one proxy of reservation wage lead to estimate of elasticity (0.15), which is in a very similar to range.

We also investigated if the regional unemployment effect on wages may have a more composite way. In specification reported in column 3 of Table 4 we introduce an interactive effect between the unemployment rate and the sales revenue per employee (*lrem_x_lun*). The same effect is than reproduced in quantile regressions (Table 6), this time defined as interaction between logarithmic changes (*dlrem_x_dlun*). While the variable is insignificant in the first specification, it remains significant in the subsequent three. It suggest that in addition to the direct effect of the two variables, we have a situation, where increase in regional unemployment is associated with higher sensitivity of wages to ability to pay. For sake of illustration, see Figure 3 with the estimated effects from column 2 of Table 6 (a range of values for simulations is taken approximately within one standard deviation each way from the sample means). The curves depict the estimated change in wages as a function of change in company revenues, at different rates of change in regional unemployment. When unemployment is increasing (see the upper curve) changes in ability to pay have stronger effect on wages. Higher revenues are conducive to higher wages irrespective of whether unemployment falls or rises but the revenue-specific differentials widen as conditions on the local labour market are deteriorating.

Most of these implications are straightforward to interpret. When revenues do not change the firm simply takes no action. When revenues rise the workers acquire a part of the gain in the form of higher wages, and their ability to do so does not strongly vary with changes in outside opportunities. When revenues fall the firm cuts wages, particularly when unemployment is on the rise. This, we believe, is consistent with the assumption of revenue sharing with incumbents who are not highly exposed to fluctuations in the labour market.

We tested if the results are driven by changes in the firm's skill composition using data on the shares of (i) low-educated, (ii) young-educated and (iii) old-educated workers with 'high education' standing for secondary or higher education, and 'old' standing for experience longer than the median. Using this data we test two sets of variables. In Table 4, column 4 we report a specification, where we control for experience and education, defining the following variables:

dp_exp_old_you = percentage change (old educated / young educated),

dp_ed_low_high = percentage change (low educated / high educated)

and in column 5:

dp_exp_old_tot = percentage change (old educated / (old educated + young educated))

dp_exp_lowedu = percentage change (low educated / low educated + high educated).

The second specification detects an effect of change in skill heterogeneity on wage dynamics, even if the time period is very short to allow much variation in those variables, and the measures applied here are crude.

We do not intend to conclude that skills and experience dimension does not matter. Interestingly, we also explored interactive effects between ability to pay and the experience and education variables. It seems that both may have some positive modifying effect on ability to pay elasticity of wages. Thus, it may be that more experienced workforce and that with higher level of education has stronger bargaining position. Nevertheless, the obtained results were insignificant and not robust to specification. For that reason we leave it for further research.

Last but not least, we hypothesised that the bargaining position of incumbent workers is likely to be stronger in the state sector. To see if this is confirmed empirically, we estimated the basic model with two sets of interactive affects: for firms with majority domestic private ownership and those with majority foreign ownership, with the state sector firms taken as a benchmark (mixed with a small number of firms with mixed ownership). Table 5 presents the corresponding specifications. There are clear differences between the state sector and both private domestic and foreign sectors. Looking into details one may see, that the differences are most clear in case of the elasticity of ability to pay and again in case of the interactive effect with sectoral wage. Ability to pay elasticity of wages is insignificantly higher in the state firms as compared with the domestic private firms, and insignificantly higher as compared with the foreign firms. In the latter group, the modifying effect of local unemployment is also strongest.

The results are consistent with the theoretical model and our prior knowledge about the bargaining position of workers, which is the strongest in the state sector, followed by the domestic private firms and weakest in the foreign firms. Correspondingly, responsiveness to the ability to pay diminishes and responsiveness to outside options increases, as predicted.

Quantile regressions (Table 7) reveal another difference with respect to state sector behaviour. Where the econometric model predicts low (negative) real wage growth, the wage growth in the state firms is stronger. On the other hand, where the model predicts high growth, the state firms are characterised by weaker growth dynamics. The corresponding inter-quantile differences are significant (Table 7, columns 4-6) and may be taken as an indication that there is some wage smoothing in the state sector (which may be consistent with the workers preferences playing more important role there). On the other hand, in this respect the foreign companies are not different from the private domestic companies.

Again, In the quantile regressions, we measure differences in bargaining power of the state sector workers by introducing the interactive effect between the state sector dummy and ability to pay (*dl_rtremp_st*).[10] But this time we are able to detect if the effect vary for different positions of firms on wage distributions. We may see that it is significant and strongest in case of companies with lowest and average wage growth and weaker and insignificant where wage growth is high. A tentative conclusion is that the bargaining strenght matters most, where the wage growth is weak. In case of companies with strongest wage growth, the modifying impact of bargaining strength is weak.

6. Conclusions

The findings from the panel analysed in the paper seem to support the basic implications of a bargaining model with incumbent workers. In particular, the wages are responsive to alternative measures of firm's ability to pay and there are structural differences between the ownership sectors consistent with our prior knowledge on relative bargaining strength and unionisation measures.

However, we examined several extensions. We augmented the bargaining specification by controls related to evolution in workers' skills and find that the basic implication of the bargaining model are not affected, even if wage dynamics is influenced by the change in composition of workers skills as approximated by education. We took a closer look at the outside options of the workers. We found that while the effect of regional unemployment on wage dynamics is significant, when an appropriate measure is used. We also found an interactive effect, where unemployment dynamics modify the impact of availability of rents on wages. In case of rising unemployment, the effect of ability to pay appears to be amplified. Wages were most considerably cut in the case of falling revenues *and* fast-rising unemployment – a situation where worker's insistence on the prevailing wage may put even incumbent jobs at risk.

[10] Alternatively, we could drop *dl_rtremp* altogether, replacing it by a set of interactive effects, created by multiplying *dl_rtremp* by all ownership sectors dummies. However in that case, we would risk enforcing the significant results for those new variables. Our specification is more conservative, i.e. it is a better method of testing if state sector affiliation has modifying effect on elasticity. The coefficient on *dl_rtremp_st* may be interpreted as a differential effect of the state sector affiliation on elasticity.

We found significant link between firm-level and industry-level wage dynamics but the estimated elasticity of firm wages with respect to sector wages were fairly low (about 0.14), calling into question if industry wage agreements have strong impact on firm-level wage determination. We also experimented with one concise indicator of reservation wage, which combines sectoral wages, unemployment benefits and regional unemployment levels, which is closest to the theoretical model specification. We found that measure performing well.

Finally, we found that while responsiveness of wages towards ability to pay is higher in the state sector, variation in wage dynamics is lower. This may indicate some wage smoothing in the state sector, consistent with the preferences of employees.

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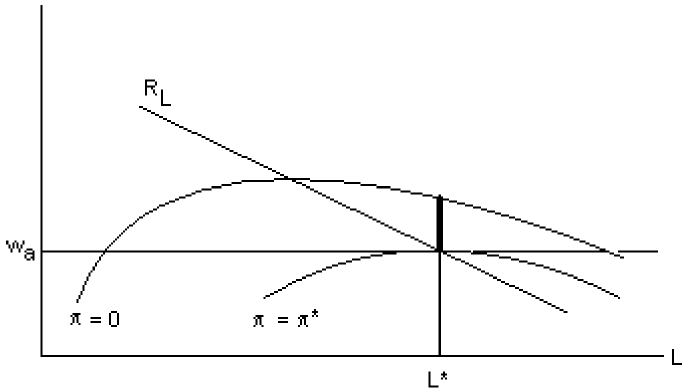
Figure 1. A model of wage bargaining



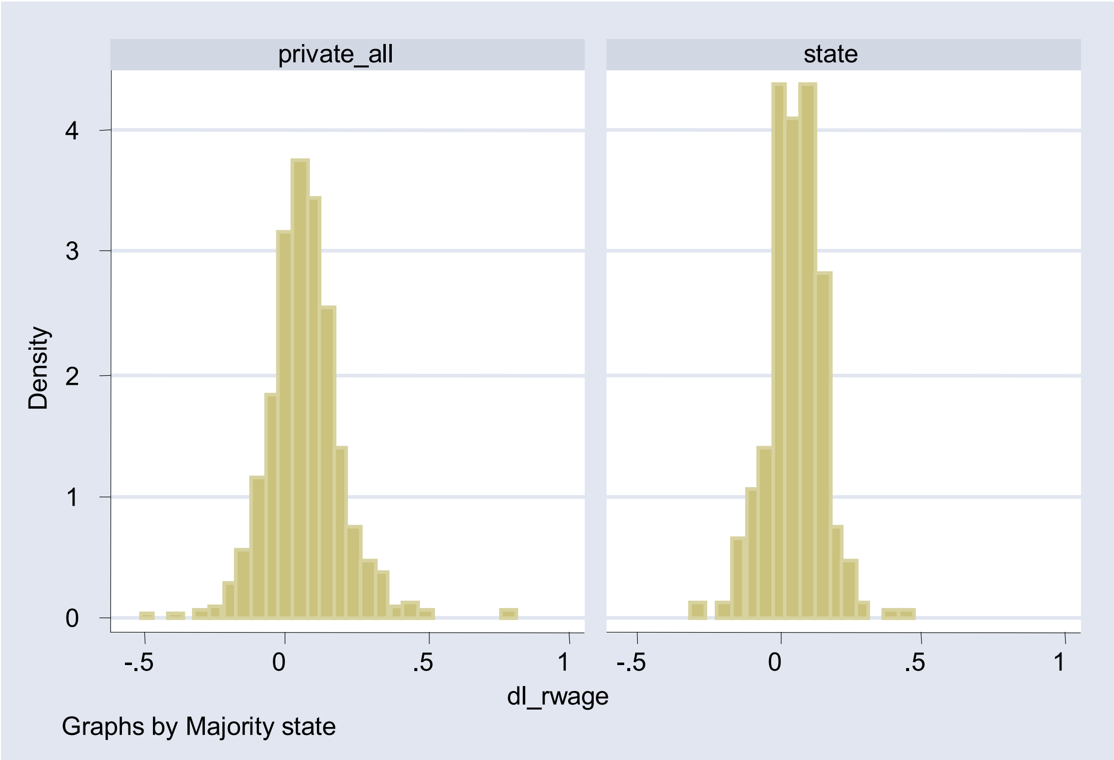


Figure 2. Histograms: wage dynamics in the state and private sector.

Figure 3. The effect of a change in (sales/worker) on the wage at various rates of change in local unemployment (simulations based on Table 6, column 2)