AN APPLICATION OF WAGE BARGAINING MODELS TO NORWEGIAN PANEL DATA

By FREDRIK WULFSBERG

Department of Economics, University of Oslo, P.O. Box 1095, Blindern, N-0317 Oslo, Norway

Wage determination in Norwegian industry is investigated using a panel of over 7,000 firms over 17 years. A standard right-to-manage bargaining model serves as theoretical background for the empirical specification. Wage equations focusing on the relative importance of insider vs outsider variables are estimated, controlling for firm-specific fixed effects. Although outsider variables are more important, insider variables have a significant effect on wages. Estimated insider weight is less than for countries with less-centralised wage determination. Using regional unemployment data we estimate unemployment elasticity between -0.04 and -0.10. There is no strong evidence for real wage resistance in the long run.

1. Introduction

THE WAGE setting system is often seen as a key feature in determining macro-economic performance in general, and the different unemployment histories between countries, in particular. During the last decade a substantial empirical literature on wage formation has emerged using bargaining theory as a theoretical framework (see Layard et al., 1991, and Calmfors, 1990). A major issue in the literature has been whether the differences in labour market institutions can explain why the EFTA countries have experienced lower unemployment rates than the EC countries during the previous two decades.

Most investigations on wage formation in Scandinavia have been based on aggregated time series data (for a survey see Calmfors, 1990) with Holden (1989), Holmlund and Zetterberg (1991), Forslund (1994), and Johansen (1996) as notable exceptions. This paper studies wage determination in Norway using annual firm level panel data from 1972–88. Panel data offers possibilities to investigate a wide range of issues related to labour market behaviour at the micro level over time, e.g. the effects of regional differences in unemployment history on the wage determination system. The present data set has the advantage that it has a considerably longer time span than previous studies of similar issues.

The basic theoretical framework of the paper is the one suggested by Nickell and Wadhwani (1990) in their analysis of UK micro data. Nickell and Wadhwani (1990) focus on the relative importance of insider vs outsider forces in wage determination. Insider forces are represented by firm-related variables, typically producer prices and productivity, while outsider forces include, for example, wages in other firms (alternative wages) and unemployment. Nickell and Wadhwani (1990) identify the ‘insider weight’ as the
long-run elasticity of the wage with respect to the average productivity of labour, and find it to be in the range of 8–15%. Nickell and Kong (1992) report an average insider weight across 11 UK industries of 13% and find little variation between industries. Lee and Pesaran (1992) estimate an average insider weight of 20% (my computations) also on UK industry data.

In contrast to the UK economy, the Norwegian economy is usually characterised as highly centralised (Bruno and Sachs, 1985; Calmfors and Driffill, 1988; and Layard et al., 1991). There are wage negotiations at national level, unionisation is high and the authorities have on several occasions intervened in the negotiations between the unions’ and the employers’ confederations, either as a third party or as a mediator. Direct action was taken when a wage and price freeze was adopted from September 1978 until December 1979. Major disputes have also been settled by arbitration in 1978 and 1981.

Centralised wage setting is often claimed to be one of the causes of the lower unemployment in the Scandinavian countries compared to EC countries (cf. Layard et al., 1991). The intuition is that with centralised wage setting, the negative external effects of higher wages are internalised and consequently firms and unions are prevented from trying to increase their relative wage. This implies a steep wage curve; wages are highly responsive to changes in unemployment. Centralised wage setting also reduces the insider weight compared to a decentralised system. In a completely centralised system the firm specific (insider) variables play no role in wage setting, hence there would be no additional information in firm level data compared to aggregate time series data. The little empirical evidence there is on the insider weight in Scandinavian economies is rather mixed. Holmlund and Zetterberg (1991) report an insignificant insider weight close to zero in the Scandinavian countries using industry level data. In contrast Johansen (1996), also using Norwegian industry data, estimates a significant insider weight of about 20%.

Although wage formation in Norway is thought of as highly centralised, wage bargaining takes place at both national and local levels. Following the nationwide negotiations, there are annual firm level negotiations over local adjustments to the central agreements (tariff wage), resulting in wage drift. However, since national bargaining seems more important in Norway than, for example, the UK, we should expect a smaller insider weight.

Because the tariff wage is pre-determined in the local wage bargain, it is possible to argue that the local bargaining system fully offsets the national agreements. Holden (1989) finds evidence against this hypothesis, which can be explained by the fact that the tariff wage affects the pay-offs of both union and firm during a dispute without a strike. Rødseth and Holden (1990) suggest that the central unions can predict the size of wage drift and thus set or negotiate a tariff wage that will meet their desired total wage increase.

1 Holmlund and Zetterberg (1991) also find an insider weight of 10% for West Germany and 30% for the US. Brunello and Wadhwani (1989) obtain an insider weight of 33% for Japan.
The wage drift’s share of the total wage increase per year varied from 30% to 80% during the 70s and 80s and has increased during the period. This may indicate that the system has become less centralised. A move towards a decentralised wage setting system would imply that the insider variables have become relatively more important than the outsider variables and as a consequence the insider weight has increased.

The paper is organised as follows. In Section 2 we outline the theoretical framework in the spirit of Nickell and Wadhwani (1990). The empirical evidence is presented in Section 3. First, we apply the same specification that Nickell and Wadhwani (1990) and Lee and Pesaran (1992) use on their UK data. The main purpose of this section is to provide a cross-country comparison. Using the same specification for different countries seems the natural way to illuminate differences in wage setting between countries. However, this specification turns out to be unsatisfactory from an econometric point of view. We therefore proceed by estimating an equilibrium correction specification which provides the basis for the exploration of various empirical issues. Section 4 concludes.

2. Theoretical background

There are several possible ways of modelling the bargaining interactions between a union and a firm (see Manning, 1987). The most popular one in applied work is the right-to-manage model where a union and a firm first bargain over wages and the firm then sets employment. One reason for adopting this model is that we observe wage bargaining taking place at fixed intervals while employment may be adjusted more periodically. Secondly, unions and firms, when asked, in general say that they do not bargain over employment (see Moene and Seierstad, 1990, for Norway and Oswald, 1987, for the UK). Due to the empirical nature of this article we outline in Appendix 1 a standard theoretical model that captures the main characteristics of this wave bargaining model.

The outcome of the wage bargaining model is an implicit wage function which presides the basis for estimation

\[ W_i = W_i(\beta_i, \bar{W}_i, U, B, Q_i, \Phi_i) \]  

Subscript \( i \) is an indicator of firm, \( W \) is the wage cost including payroll taxes, \( \beta \) is the relative bargaining power of the union, \( \bar{W} \) is the alternative wage, \( U \) is unemployment rate, \( B \) is the unemployment benefit, \( Q \) is productivity, and \( \Phi \) is the wedge between real labour cost and real consumption wages due to prices and taxes.

Nickell and Wadhwani (1990) specify a Cobb–Douglas production function, so that the logarithm of the wage cost equals a logarithmic sum of the insider variables, outsider variables, and the wedge. Using lowercase letters to denote logarithms, we then obtain
where $p$ is the product price, $y$ is output, and $n$ is employment. Following Nickell and Wadhwani (1990), in the empirical analyses we impose the long-run homogeneity restriction; $\mu_1 + \mu_2 = 1$, implying that wages increase proportionally to an equal increase in revenue per head, $(p + y - n)$, and alternative wages. Given this homogeneity, $\mu_1$ is an estimate of the relative importance of insider variables and is named the insider weight.

The term $\Delta n$ is included as a test for insider hysteresis effects. The motivation for this is that the union cares about the insiders, assumed to be equal to last period’s employment. Then for a given wage the probability of being laid off is lower the smaller the last period’s employment (Blanchard and Summers, 1986).

If the wedge affects the wage costs, this is usually referred to as real wage resistance. Many empirical studies assume a priori that $\mu_3 = 0$. The motivation for this is usually empirical since there does not seem to be any long-run relationship between employment and taxes or prices. Given that employment is independent of the wedge, the real labour cost is also independent of the wedge, so that any change in the wedge must shift the real consumption wage. A theoretical argument against the existence of a correlation between wage costs and the wedge applies when capital is highly mobile between economies, since this implies a given rate of return on capital in equilibrium. In this case an increase in the wedge shifts the real consumption wage only. Nickell and Wadhwnani (1990) assume that the union utility function is either isoelastic or risk neutral, which implies that $\mu_3 = 0$.

Equation (2) can be interpreted as representing a steady state equilibrium due to its static nature. However, in an empirical specification of the model we need to allow for dynamic adjustments. Furthermore, a dynamic specification is also necessary to study some of the much debated issues concerning wage formation, such as hysteresis or inertia. In addition to the insider effect discussed above, hysteresis may arise as a consequence of the unemployed becoming less efficient job seekers the longer they stay unemployed. Nickel and Wadhwnani (1990) find some evidence consistent with this, with an increase in the proportion of long term unemployment increasing wages. However, reliable data on the proportion of unemployed who have been unemployed for more than one year in Norway do not exist for the time period of our analysis. On the other hand, since long-term unemployment is negatively correlated with the change in unemployment, an increase in unemployment may reduce the wage level.

The possible existence of nominal inertia suggests that changes in, as well as levels of, nominal variables may affect wages. A general dynamic and stochastic specification of (2) which also allows for firm specific fixed effects is

$$w_{it} = f_i + \delta_1(\Lambda)w_{i(t-1)} + \delta_2(\Lambda)[(p + y - n)]_t + \theta_1(\Lambda)\Delta n_t + \theta_2(\Lambda)b_t$$
$$+ \delta_3(\Lambda)[\bar{w}_t + \theta_3(\Lambda)u_t + \theta_4(\Lambda)b_t] + \delta_4(\Lambda)\phi_t + \varepsilon_{it}$$  

(3)
where $\varepsilon_{it}$ is assumed to be a stochastic error term $\sim iid(0, \sigma^2)$, $\Lambda$ is the lag-operator, and $\delta(\Lambda)$ and $\theta(\Lambda)$ are polynomials in the lag-operator. To ensure long-run homogeneity the condition $\delta_1(\Lambda) + \delta_2(\Lambda) + \delta_3(\Lambda) = 1$ is assumed. $f_i$ represents unobserved fixed effects, i.e. variation between firms which is constant over time. In panel data studies individual specific terms are usually thought of as firm specific, but they may also capture regional, industry, or any other time constant characteristic.

3. The Norwegian evidence

The present data set covers the entire population of Norwegian manufacturing firms over the period 1972–88. It is an unbalanced panel data set, but we keep only firms existing for more than nine subsequent years. This gives a sample of 7,323 firms covering on average 70% of the employees in the manufacturing sector. The firms are quite small with an average of 40 employees.

The variables are fully described in Appendix 2, but we will briefly comment on some of their main features. Employment is measured as average employment over the year, and wages are measured by the wage bill divided by employment. The revenue per head term $(p + y - n)$ is sales divided by average employment over the year, as is standard in the literature. Ideally we would prefer revenue per head net of input costs, but input costs are imperfectly observed, especially for energy costs.

The external variables $w$ and $u$ are region specific. The alternative wage refers to labour districts, of which there are approximately 100 of varying sizes. It is computed as the average wage per worker in all firms except firm $i$, within the same labour district. The rate of unemployment relates to counties, of which there are 19. We will also use the aggregate unemployment rate. The payroll tax rate varies between five zones, from 0% in the high north to 17.8% in the densely populated areas. The variation between the zones has also changed over time. The average payroll tax rate has varied much less over the same period.

Firm level variables on financial performance and union characteristics are not included in the study as such data are not available. However, union density has been fairly stable over the period of interest, with the majority of variation occurring between sectors. This will be controlled for in a fixed-effects model.

3.1. The Nickell–Wadhwani model

We estimate (3) using generalised method-of-moments estimation (GMM) as proposed by Arellano and Bond (1991). The GMM method yields unbiased

---

2 A referee has suggested keeping only the balanced part of the data set. This would probably make the remaining firms more homogeneous but at the cost of making the results less representative for the whole manufacturing sector.

3 Reliable labour force statistics are not available at the labour district level.
estimators in dynamic models when the unobserved fixed effects are correlated with the regressors. Furthermore, the Arellano and Bond (1991) procedure is a two-step method. Arellano and Bond (1991) report simulations which show that the two-step standard error estimates in short samples \((T = 10)\) are biased downwards, so that standard critical values of the \(t\)-statistic are no longer valid. This bias decreases rapidly with the sample size so we therefore report the two-step estimates (GMM2) in all models.

The GMM procedure controls for the fixed effects by first differencing (3), which results in

\[
\Delta w_{it} = \delta_1(\Lambda)\Delta w_{i(t-1)} + \delta_2(\Lambda)[\Delta(p + y - n)_{it} + \theta_1(\Lambda)\Delta^2 n_{it} + \theta_2(\Lambda)\Delta \beta_i] + \delta_3(\Lambda)\Delta \bar{w}_{it} + \theta_3(\Lambda)\Delta u_i + \theta_4(\Lambda)\Delta b_i] + \delta_4(\Lambda)\Delta \phi_{it} + \Delta \varepsilon_{it}
\]

Then the procedure uses lags of the endogenous variables as instruments, based on the orthogonality assumptions in (3); i.e. under the null hypothesis of no serial correlation in \(\varepsilon_{it}\). The transformed disturbance (\(\Delta \varepsilon_{it}\)) in (4) will be uncorrelated with lags of the endogenous variables dated \(t - 2\) and before. The insider variables, \((p + y - n)_{it}\), \(w_{it}\) and \(n_{it}\) are treated as endogenous variables, while the other variables are treated as exogenous.

The first difference transformation implies that coefficients on variables possessing only cross-sectional variation (i.e. variables that are constant over time) are not identified. However, this potential methodological cost has no practical implications for our model since all firm information available in this study vary in both dimensions. Nevertheless, there is a cost to differencing, which increases the variance of the error term and reduces the signal from the regressors.\(^4\)

In Table 1 we report three regressions. In model (A) I have put restrictions on (4) according to Nickell and Wadhwani (1990) and in model (B) the restrictions are along the lines of Lee and Pesaran (1992). Model (B) differs from model (A) in the lag structure, and also includes the inflation rate. Model (C) allows for richer dynamics than (A) and (B) and encompasses both of these models. Furthermore, in model (C) I allow for different elasticities of the price wedge, \((cp = pk) = \log (CP/P_k)\), and the tax wedge, \((prt_j - tax) = \log [(1 + PRT)/(1 - TAX)]\). \(CP\) is the consumer price index, \(P_k\) product price index for sector \(k\), \(PRT\) is the payroll tax rate, while \(TAX\) is the income tax rate. All regressions are estimated with the restriction that the sum of the long run elasticities of the productivity and alternative wages equal unity, i.e. \(\delta_1(\Lambda) + \delta_2(\Lambda) + \delta_3(\Lambda) = 1\).

Turning to the estimates, we find in all regressions that the short-run effect of the alternative wage is large and significant. The corresponding long-run

\(^4\)The noise is increased because the transformed error term has twice the variance of the untransformed error, \(\text{Var}(\Delta \varepsilon) = 2\text{Var}(\varepsilon)\). The signal is reduced because the variance of the transformed regressors is less than the untransformed when they are positively autocorrelated, \(\text{Var}(\Delta x) < \text{Var}(x)\) when \(\text{cov}(x_t, x_{t-1}) > 0\). Both effects increase the standard errors of the estimators.
TABLE 1
GMM2 estimates of wage bargaining models similar to Nickell and Wadhwani (1990), (A), and Lee and Pesaran (1992), (B), applied to Norwegian data. Dependent variable is $w_t$; absolute t-values in parentheses

<table>
<thead>
<tr>
<th>Regressors</th>
<th>Model (A)</th>
<th>Model (B)</th>
<th>Model (C)</th>
</tr>
</thead>
<tbody>
<tr>
<td>constant</td>
<td>-0.003 (1.51)</td>
<td>-0.003 (12.62)</td>
<td>-0.003 (9.32)</td>
</tr>
<tr>
<td>$w_{t(t-1)}$</td>
<td>0.341 (40.05)</td>
<td>0.312 (39.56)</td>
<td>0.313 (38.67)</td>
</tr>
<tr>
<td>$(p + y - n)_{ht}$</td>
<td>0.041 (4.64)</td>
<td>0.047 (7.49)</td>
<td>0.042 (6.42)</td>
</tr>
<tr>
<td>$(p + y - n)_{h(t-1)}$</td>
<td>0.004 (1.02)</td>
<td>-0.000 (0.003)</td>
<td></td>
</tr>
<tr>
<td>$\bar{w}_t$</td>
<td>0.613 (-)</td>
<td>0.641 (-)</td>
<td>0.645 (-)</td>
</tr>
<tr>
<td>$\Delta n_{ht}$</td>
<td>-0.055 (4.60)</td>
<td>-0.031 (2.98)</td>
<td>-0.030 (2.66)</td>
</tr>
<tr>
<td>$u_{ht}$</td>
<td>-0.002 (5.52)</td>
<td>-0.003 (6.00)</td>
<td></td>
</tr>
<tr>
<td>$p_{(t-1)}$</td>
<td>0.002 (0.80)</td>
<td>-0.000 (1.00)</td>
<td></td>
</tr>
<tr>
<td>$\Delta p_{ht}$</td>
<td>0.182 (9.19)</td>
<td>0.186 (8.81)</td>
<td></td>
</tr>
<tr>
<td>$c_{(t-1)}$</td>
<td>-0.038 (2.89)</td>
<td>-0.038 (2.89)</td>
<td>-0.036 (1.35)</td>
</tr>
<tr>
<td>$cp_{ht} - tax$</td>
<td>-0.002 (5.52)</td>
<td>-0.003 (6.00)</td>
<td></td>
</tr>
<tr>
<td>$c_{(t-1)} - tax$</td>
<td>0.002 (0.80)</td>
<td>-0.000 (1.00)</td>
<td></td>
</tr>
<tr>
<td>WALD-TD (15)</td>
<td>404.3</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>AR(1)</td>
<td>-36.6</td>
<td>-36.2</td>
<td>-35.4</td>
</tr>
<tr>
<td>AR(2)</td>
<td>5.1</td>
<td>4.2</td>
<td>4.2</td>
</tr>
<tr>
<td>WALD (df)</td>
<td>1,731 (5)</td>
<td>1,828 (5)</td>
<td>1,797 (9)</td>
</tr>
<tr>
<td>SARGAN (df)</td>
<td>754 (356)</td>
<td>840 (357)</td>
<td>816 (356)</td>
</tr>
<tr>
<td>$\sigma^2$</td>
<td>0.025</td>
<td>0.024</td>
<td>0.25</td>
</tr>
</tbody>
</table>

Notes:
(i) $N = 7323$ and $T_t = 10, ..., 17$.
(ii) All equations estimated as first differences using the GMM method of Arellano and Bond (1991). $w_{t(t-1)}$, $\Delta n_{t(t-1)}$, $(p + y - n)_{ht}$ and $(p + y - n)_{h(t-1)}$ are treated as endogenous, while $\bar{w}_t$, $u_t$, $cp_t$, $p_t$, $tax$, $prt$ and time dummies are treated as exogenous. Additional instruments are $w_{t(t-2)}$, $n_{t(t-2)}$, $(p + y - n)_{h(t-2)}$ where $s = 2, ..., T$.
(iii) Two-step estimates, test statistics are robust to heterogeneity.
(iv) WALD-TD is a Wald test for joint significance of the time dummies. SARGAN-test for instrument validity. The AR(2) statistic is asymptotically $N(0,1)$, while the SARGAN and WALD statistics are $\chi^2$.

Elasticities of the alternative wage are 0.931 in model (A), 0.932 in model (B), and 0.939 in model (C). Thus, the insider weight is 6.9% in model (A), 6.8% in model (B), and 6.1% in model (C). These estimates are smaller than those from the studies on UK data. In all columns the revenue elasticity is positive and significant. The change in productivity is also included as regressor in model (C), but is not significant.

The insider hysteresis effect ($\Delta n$) obtains a significant negative sign, i.e. the opposite of the Nickell and Wadhwani (1990) result. A possible explanation of this negative sign might be that firms lay off workers in reverse order of seniority when reducing employment, so that average wage costs per worker increase (assuming that wages are positively related to experience).

The elasticity of lagged unemployment in model (A) is positive but insignificant, while contemporaneous unemployment has a significant negative
impact in models (B) and (C). The corresponding long-run unemployment elasticities of wages for a given alternative wage, which we shall refer to as the partial unemployment elasticity of wages, are $-0.003$ and $-0.004$ respectively. Nickell and Wadhwani (1990) (using industry unemployment) report a partial unemployment elasticity of $-0.101$ while Lee and Pesaran (1992) report a partial elasticity of $-0.023$. The partial unemployment elasticity estimated here may seem small compared to evidence from aggregate time-series studies. However, if we take into account the effect of unemployment on alternative wages by imposing the alternative wage to equal the wage in the long run, we obtain a total unemployment elasticity of wages of $-0.044$ and $-0.066$ for the models in models (B) and (C) respectively. The effect of a change in unemployment in model (C) is found to be insignificant.

In the encompassing regression of model (C) we include the tax and price wedge. Only the price wedge is significant, but neither the price wedge elasticity nor the tax wedge elasticity has the expected positive sign. The growth rate of the consumer price is also significant, but as the only aggregate variable in the equation it might capture other aggregate effects that are common to all firms. Furthermore, as shown by Moulton (1990), the standard errors of coefficients representing aggregate effects in micro studies are biased downwards, which yields an upward bias in the $t$-statistics. The reason is that there will be correlation between the error terms of different observations within the same group) (industry, region, etc.). The bias is probably reduced by controlling for time constant individual effects, but it will not be eliminated.

The diagnostic tests suggest that these models do not explain the data well. First, the significant second order autocorrelation test $(AR(2))$ indicates that there are important unexplained dynamic effects. Second, the Sargan test rejects the validity of the instruments. If the assumption of uncorrelated errors in the levels equation does not hold, lagged variables will not be valid instruments. This might be a further indication of a more complex dynamic structure in the data. A third indication of unexplained dynamics in the data is the high degree of autocorrelation. One possible source of unexplained dynamics is that increases in the wage rates take place in negotiations during the year, while the data measures average wages over the year. An increase in wage rates during year $t$ will thus yield a rise in average wages from year $t - 1$ to year $t$ and from year $t$ to year $t + 1$.

Although some of the estimated coefficients seem plausible the tests for misspecification indicate that the model misses out important dynamic effects. This misspecification of the model might follow from rather ad hoc assumptions on the dynamic structure. The model includes mostly variables in levels

---

5 Since the equation is estimated in first differences, it follows that if the error term from the level eq. (3) is uncorrelated, the differenced error term in (4) is an MA(1) process and should thus fail a test for first order autocorrelation, but not of second order.

6 Holmlund and Zetterberg (1991) find a higher degree of autocorrelation in Norway than in other countries.
which might fail to pick up short run effects. A specification which allows for richer dynamics in an equilibrium correction model.

3.2. Equilibrium correction models

Equilibrium correction models (ECMs) estimate both short- and long-run effects based on initial assumptions about the long-run relationships between variables. In our model this long-run relationship is given by (2) with $\mu_1 + \mu_2 = 1$, which in equilibrium correction terms is specified as a cointegration relationship between wages, revenue per head, and alternative wages. An ECM which encompasses (4) and satisfies the long run restrictions in (2) is

$$\Delta w_{it} = f_i + \alpha_1(\Lambda)\Delta w_{i(l-1)} + \alpha_2(\Lambda)\Delta \bar{w}_{it} + \alpha_3(\Lambda)\Delta(p + y - n)_{it}$$

$$+ \beta_1 s_{i(l-1)} + \beta_2 (w - \bar{w})_{i(l-1)} + \gamma_1(\Lambda)\phi_{it} + \gamma_2(\Lambda)u_{it} + \epsilon_{it}$$

where $s = w - (p + y - n)$ is the wage share. If a large wage increase takes place in period $t$, bringing the wage share and relative wage away from the long-run equilibrium levels, this will induce a correction the next period through the insider equilibrium correction term, $s_{i(l-1)}$, and the outsider equilibrium correction term, $(w - \bar{w})_{i(l-1)}$. Thus we expect the estimates of $\beta_1$ and $\beta_2$ to be negative, but less than unity in magnitude. The closer the estimate is to $-1$, the faster the correction.

The long-run steady state of the ECM (5) is

$$w = \frac{\beta_1}{\beta_1 + \beta_2}(p + y - n) + \frac{\beta_2}{\beta_1 + \beta_2}\bar{w} - \frac{\gamma_1(1)}{\beta_1 + \beta_2}\phi - \frac{\gamma_2(1)}{\beta_1 + \beta_2}u$$

where $\gamma_1(1)$ and $\gamma_2(1)$ are the sums of the coefficients in the lag polynomials. Comparing eq. (6) with eq. (2) we see that the relationship between the ECM parameters and the steady-state parameters are $\mu_1 = \beta_1/(\beta_1 + \beta_2)$, $\mu_2 = \beta_2/(\beta_1 + \beta_2)$ and $\mu_3 = \gamma_2(1)/(\beta_1 + \beta_2)$.

The appropriate approach when specifying ECMs is to first investigate the time series and cointegrating properties of the variables involved. The state of the art in testing for unit roots and cointegration in panel data is, however, still juvenile\(^7\) and breaking new ground on this issue is beyond the scope of this paper. Nevertheless, from time-series econometrics (see e.g. Banerjee et al., 1993) we know that the ECM can be used for testing for cointegration between both $w$ and $(p + y - n)$, and $w$ and $\bar{w}$. In this test one should note that the $t$-statistic under the null of no cointegration has a non-standard distribution with a critical value of $-2.99$ for a time series with $T = 25$ at the 10% significance level (Banerjee et al., 1993, Table 7.6). However, these critical values are probably not relevant to panel data. Panel data will in general increase the degrees of freedom and hence probably reduce the critical

\(^7\)Levin and Lin (1992), Harris and Tzavalis (1996) and Im et al. (1996) are recent suggestions.
values. To my knowledge critical values for panel data do not exist in this context.

The ECM is consistent with at least two possible interpretations of the long-run relationship between the variables in the equilibrium correction terms. Either \( w, \bar{w}, \) and \((p + y - n)\) are jointly cointegrated, which implies that relative wages cointegrate with the wage share or they are not jointly cointegrated, but the relative wage and the wage share are both \( I(0) \) (Banerjee et al., 1993, p. 155). However, in both cases \( \beta_1 \) and \( \beta_2 \) are different from zero and \( \mu_1 \) can be interpreted as the insider weight.\(^8\)

In order to get an idea of the time series properties of the relative wage we have plotted its logarithm over time in Fig. 1. We see that the distribution of relative wages does not change dramatically over the years. Each series of a relative wage may nevertheless still be \( I(1) \). If this is true there should be little correlation between the first and last observation of each relative wage considering the fairly large time span. As the correlation coefficient is 0.29, it is not unreasonable to assume that the relative wage is stationary.

Returning to (5) we take first differences in order to control for the unobserved individual fixed effects, and estimate the resulting differenced equation using GMM.\(^9\) Table 2 reports various regressions including time

\(^8\) An alternative possibility to cointegration is of course that all variables are stationary (i.e. \( I(0) \)); the only implication being that (5) can be interpreted as a partial adjustment model.

\(^9\) Note that the equilibrium correction model (5) is not differenced, it is just a reparametrisation of a distributed lag model in levels.
**Table 2**

GMM2 estimates of wage bargaining models specified as equilibrium-correction models. Endogenous variable is $\Delta w_{it}$; time dummies are included.

<table>
<thead>
<tr>
<th>regressors</th>
<th>model</th>
<th>(D)</th>
<th>(E)</th>
<th>(F)</th>
<th>(G)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta w_{it(-1)}$</td>
<td>$-0.112$ (14.32)</td>
<td>$-0.112$ (14.29)</td>
<td>$-0.112$ (14.26)</td>
<td>$-0.112$ (14.28)</td>
<td></td>
</tr>
<tr>
<td>$\Delta w_{it(-2)}$</td>
<td>$-0.041$ (8.94)</td>
<td>$-0.041$ (8.92)</td>
<td>$-0.040$ (8.90)</td>
<td>$-0.041$ (8.92)</td>
<td></td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{it}$</td>
<td>$0.065$ (5.22)</td>
<td>$0.065$ (5.19)</td>
<td>$0.064$ (5.11)</td>
<td>$0.064$ (5.12)</td>
<td></td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{it(-1)}$</td>
<td>$0.015$ (4.79)</td>
<td>$0.015$ (4.72)</td>
<td>$0.015$ (4.69)</td>
<td>$0.015$ (4.73)</td>
<td></td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{it(-2)}$</td>
<td>$0.006$ (3.05)</td>
<td>$0.006$ (3.07)</td>
<td>$0.006$ (3.08)</td>
<td>$0.006$ (3.11)</td>
<td></td>
</tr>
<tr>
<td>$\Delta \tilde{w}_{it}$</td>
<td>$0.173$ (8.87)</td>
<td>$0.158$ (8.07)</td>
<td>$0.158$ (8.07)</td>
<td>$0.158$ (8.08)</td>
<td></td>
</tr>
<tr>
<td>$\Delta \tilde{w}_{it(-1)}$</td>
<td>$-0.189$ (9.46)</td>
<td>$-0.182$ (9.09)</td>
<td>$-0.183$ (9.12)</td>
<td>$-0.184$ (9.18)</td>
<td></td>
</tr>
<tr>
<td>$\Delta \tilde{w}_{it(-2)}$</td>
<td>$-0.103$ (5.70)</td>
<td>$-0.101$ (5.54)</td>
<td>$-0.102$ (5.58)</td>
<td>$-0.102$ (5.62)</td>
<td></td>
</tr>
<tr>
<td>$s_{it(-1)}$</td>
<td>$-0.020$ (2.15)</td>
<td>$-0.020$ (2.17)</td>
<td>$-0.020$ (2.12)</td>
<td>$-0.020$ (2.11)</td>
<td></td>
</tr>
<tr>
<td>$(w - \tilde{w})_{it(-1)}$</td>
<td>$-0.455$ (26.85)</td>
<td>$-0.455$ (26.89)</td>
<td>$-0.456$ (26.93)</td>
<td>$-0.456$ (26.92)</td>
<td></td>
</tr>
<tr>
<td>$\Delta pr_{it}$</td>
<td>$0.712$ (4.87)</td>
<td>$0.734$ (4.98)</td>
<td>$0.695$ (6.18)</td>
<td>$-0.064$ (0.31)</td>
<td></td>
</tr>
<tr>
<td>$pr_{it}$</td>
<td>$-0.064$ (0.31)</td>
<td>$-0.087$ (0.42)</td>
<td>$0.006$ (1.14)</td>
<td>$0.006$ (1.10)</td>
<td></td>
</tr>
<tr>
<td>$u_{it}$</td>
<td>$0.006$ (1.14)</td>
<td>$0.006$ (1.10)</td>
<td>$-0.003$ (0.82)</td>
<td>$-0.003$ (0.81)</td>
<td></td>
</tr>
<tr>
<td>$\Delta u_{it}$</td>
<td>$-0.003$ (0.82)</td>
<td>$-0.003$ (0.81)</td>
<td>$-33.9$</td>
<td>$-33.8$</td>
<td></td>
</tr>
<tr>
<td>$AR(1)$</td>
<td>$-33.9$</td>
<td>$-33.8$</td>
<td>$-33.8$</td>
<td>$-33.8$</td>
<td></td>
</tr>
<tr>
<td>$AR(2)$</td>
<td>$0.5$</td>
<td>$0.5$</td>
<td>$0.5$</td>
<td>$0.5$</td>
<td></td>
</tr>
<tr>
<td>$\hat{\sigma}^2$</td>
<td>$0.023$</td>
<td>$0.023$</td>
<td>$0.023$</td>
<td>$0.023$</td>
<td></td>
</tr>
</tbody>
</table>

Notes: see Table 1.

- dummies as regressors, based on the ECM. The order of the lag-operator is chosen to include two lags. Significant lags of the endogenous variable imply wage persistence.

- The diagnostic tests show that the ECMs perform better in all respects than the models in Table 1. The serial correlation tests do not reject the hypothesis of no autocorrelation in (5), the WALD test statistic for joint significance is more than tripled and the SARGAN test statistic reduced by almost half. In spite of this, all the SARGAN test statistics still reject the validity of the instruments. This may indicate that some variables should not be treated as exogenous regressors. Likely candidates are the current unemployment rate and the current alternative wage. (Note, however, that the alternative wage for firm $i$ does not include the wage in firm $i$ by construction. In addition, the regions over which the alternative wage and unemployment rate are computed are fairly large.) However, using only lags of the growth in alternative wages and unemployment as instruments (not reported here) does not improve the SARGAN statistic. Arellano and Bond (1991) report simulation results showing that the SARGAN test rejects too often in the presence of heteroskedasticity which might explain our result. The satisfactory test statistics for first and second order serial correlations are also important indicators of instrument
validity, and as we see from Table 2 none of these rejects the hypothesis of serially uncorrelated regressors in the levels eq. (5).

The outside equilibrium term is strongly significant in a test for cointegration between wages and relative wages, even if we use the critical values computed for time series models (Banerjee et al., 1993). On the other hand, the inside equilibrium correction estimate is not significant in a similar test using the same critical values. But, since these critical values are based on time series, this cannot be interpreted as strong evidence of no cointegration between wages and revenue per worker. There is no reason, however, to reject a joint cointegrating relationship between these variables.

Conditional on cointegration, the equilibrium correction terms have a standard distribution. In model (D) of Table 2, eq. (5) is estimated with the restriction $\gamma_1 = \gamma_2 = 0$. 95\% confidence intervals for $\beta_1$ and $\beta_2$ are $(-0.039, -0.002)$ and $(0.488, 0.422)$ respectively. Wages are thus more responsive to changes in the relative wage. If the relative wage and wage share are $I(0)$, 45.5\% of the deviation from the equilibrium relative wage is corrected in the next period. The correction for a deviation from the equilibrium wage share is 2.0\% for a given alternative wage. The long run solution of model (D) is obtained using formula (6) which yields

\[
\text{model (D): } w = 0.043(p + y - n) + 0.957\bar{w}
\]

Thus the insider weight is 4.3\%, which is slightly smaller than the previous estimates of Table 1.

As all the models of Table 2 include time dummies to account for aggregate effects, it is not possible to identify the separate effects of the wedge because there is no cross-sectional variation in the consumer price and income tax. The time dummies will also control for effects of the arbitrations and the wage freeze in 1979. However, the payroll tax rate varies both between regions and over time. In models (E) and (F) I include both $prt$ and $\Delta prt$. Since the left hand side variable is wage cost, the small and insignificant effects of $prt$ in both models indicate that an increase in payroll taxes is borne by labour in the long run. There is nevertheless quite a significant short run effect of an increase in payroll taxes on the growth in wage costs.

In models (F) and (G) I include regional unemployment which has no significant effects either on wage growth or on the wage level. The long run solutions of models (E), (F) and (G) are

\[
\text{model (E): } w = 0.043(p + y - n) + 0.957\bar{w} - 0.135prt
\]

\[
\text{model (F): } w = 0.043(p + y - n) + 0.957\bar{w} - 0.135prt
\]

\[
\text{model (G): } w = 0.043(p + y - n) + 0.957\bar{w} - 0.135prt
\]

10 There is solid evidence for cointegration between wages and revenue per worker in aggregate data, see for example Johansen (1995).

11 $t$-ratios for the long run estimates which are shown in parentheses, are computed using the method of Bårdsen (1989).
\[ w = 0.042(p + y - n) + 0.958 \tilde{w} + 0.013 u_j - 0.182 prt \] (9)

\[ w = 0.042(p + y - n) + 0.958 \tilde{w} + 0.014 u_j \] (10)

We see that the insider weight is not affected by including the unemployment rate or the payroll tax rate.

The insider weight is considerably smaller than in Johansen (1996) who reports an insider weight close to 20% using Norwegian industry level panel data from roughly the same period (1966–87). As Johansen (1996) also estimates equilibrium correction models the different estimates cannot be explained by differences in model specifications. However, when we compare industry level estimates with firm level estimates, we should bear in mind that some variables are characterised as insider variables in the industry level equations, but as outsider variables in the firm level equations (e.g. wages in other firms within the industry). Consequently we should expect a larger insider weight from industry level data. This is consistent with the differences observed between Johansen’s results and those presented here. The present estimates are more like those of Holmlund and Zetterberg (1991), who find an estimate of 4% (not significant) for the insider weight using Norwegian industry panel data 1965–82. As Holmlund and Zetterberg (1991) use industry data one should expect their estimates to be larger than ours. However, Holmlund and Zetterberg’s estimates are imprecisely determined. Furthermore, their model exhibits significant autocorrelation and is therefore likely to be dynamically misspecified. Forslund (1994), using firm level panel data from Sweden 1984–88, reports an insignificant\(^{12}\) insider weight of 7%. Due to the short time period Forslund’s model is similar to the Nickell-Wadhwani specification discussed in Section 3.1.

The small insider equilibrium correction estimates (−0.02) indicate a very slow adjustment of wages to internal variables, compared to the result that almost half of the deviation from the equilibrium relative wage is adjusted within one period. The significant coefficient shows that internal variables do affect wages, but the effect is small. The much faster adjustment to the relative wage may imply that unions are more concerned with relative wages and external factors in the labour market. In other words, local wage bargaining in Norway may be as much about the wage distribution as profit sharing. The nationwide agreements have always been favourable to low-wage sectors in order to counteract the effects of local wage bargaining on wage dispersion.

The main findings in Table 2 are that the insider weight is about 4% and that it is independent of the wedge and unemployment. Wages respond very slowly to a change in internal variables. There is strong evidence of a cointegrating

\(^{12}\)The insider weight is insignificant according to the author although no standard error is reported.
relationship between wages and alternative wages. Furthermore, there is no
evidence that the unemployment affects wages.\textsuperscript{13}

3.2.1. \textit{Parameter stability and heterogeneity.} Being in the fortunate situation
of having panel data over 17 years, it is possible to test the stability of the
parameter estimates over time. As four years of observations per establishment
are lost due to lags and differencing, I have chosen to split the data into two
sub-samples of approximately the same size, one for 1976–81 and one for 1982–
88. During the 1980s the ratio of wage drift relative to tariffs has increased,
leading to questions about whether the bargaining system has become less
centralised. According to this hypothesis the insider weight for the latter
period should be larger than in the first. Johansen (1996) finds that the insider
weight increased during the 1980s and that it is larger for expanding industries
than for contracting ones. Edin and Holmlund (1993) find that the insider
weight has increased in Sweden during the 1980s.

We test the specification of model (G) in Table 2 which includes the effects of
unemployment and a short-run effect of the tax wedge. I proceed by con-
structing dummy variables for the latter sub-period and interact these with
the regressors. The estimates (omitting the less interesting short run elasticities
to save space) are reported in Table 3.

A \textit{WALD}-test statistic of 49.5 indicates rejection of the null hypothesis of no
change in the parameters. A test for no change in the long-run parameters is
also rejected with a statistic of 8.9. As seen from the table, the insider weight
has at face value decreased from the 1970s to the 1980s giving little support to
the hypothesis that the wage bargaining system became less centralised in the
1980s. However, only the reduction in the outsider equilibrium estimate (from
$-0.511$ to $-0.423$) is significant which indicates that the insider weight may
have increased. Increasing short-run effects of productivity growth provides
further evidence that this may be the case.

So far I have assumed parameter homogeneity across the firms. The reported
estimates should thus be interpreted as average elasticities if we believe that the
firms react differently. However, Pesaran \textit{et al.} (1996) show that pooled
estimators in dynamic panel data models are biased and inconsistent even
when both the number of individuals and the time dimension are large. An
important ingredient in this bias is the degree of heterogeneity in the
parameters of interest. Pesaran \textit{et al.} (1996) suggest estimating separate
models for each establishment and then computing the mean estimates. This
procedure, however, requires a large time dimension in the data. As the present

\textsuperscript{13}The null hypothesis of long run homogeneity ($\mu_1 + \mu_2 = 1$) is rejected when we include the
lagged alternative wage as regressor in model (D) of Table 2 and in models (H) and (K) in Table 5
which follows in the next section. However, the insider weight is almost unaffected. To save space I
do not report the tests. I find it hard to neglect the strong theoretical arguments for long-run
homogeneity. The results are of course available from the author upon request.
Testing for parameter stability
GMM2 estimates of wage bargaining models specified as equilibrium-correction models, endogenous variable is $\Delta w_t$.

<table>
<thead>
<tr>
<th>Regressors</th>
<th>1976–81</th>
<th>1982–88</th>
</tr>
</thead>
<tbody>
<tr>
<td>$s_{t(1)}$</td>
<td>-0.034</td>
<td>-0.016</td>
</tr>
<tr>
<td>$(w - \bar{w})_{t(1)}$</td>
<td>-0.511</td>
<td>-0.423</td>
</tr>
<tr>
<td>$u_t$</td>
<td>0.020</td>
<td>-0.024</td>
</tr>
<tr>
<td>Insider weight</td>
<td>0.063</td>
<td>0.037</td>
</tr>
<tr>
<td>WALD no change (all parameters (13))</td>
<td>49.5</td>
<td></td>
</tr>
<tr>
<td>WALD (13) long-run parameters</td>
<td>8.9</td>
<td></td>
</tr>
<tr>
<td>SARGAN (214)</td>
<td>5,847.1</td>
<td></td>
</tr>
<tr>
<td>AR(1)</td>
<td>-31.8</td>
<td></td>
</tr>
<tr>
<td>AR(2)</td>
<td>1.0</td>
<td></td>
</tr>
</tbody>
</table>

Notes:
(i) See Table 1.
(ii) The t-statistics for the 1982–88 estimates refer to a test of no difference from the first period.

Equilibrium correction estimates and insider weights by industry.

<table>
<thead>
<tr>
<th>Industry (SIC code)</th>
<th>Number of firms</th>
<th>$s_{t(1)}$</th>
<th>$(w - \bar{w})_{t(1)}$</th>
<th>Insider weight</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mining (21, 22, 23, 29)</td>
<td>221</td>
<td>-0.109</td>
<td>-0.604</td>
<td>0.152</td>
</tr>
<tr>
<td>Food, beverages, and tobacco (31)</td>
<td>1,706</td>
<td>-0.076</td>
<td>-0.408</td>
<td>0.156</td>
</tr>
<tr>
<td>Textiles (32)</td>
<td>455</td>
<td>-0.048</td>
<td>-0.581</td>
<td>0.076</td>
</tr>
<tr>
<td>Wood (33)</td>
<td>1,271</td>
<td>-0.019</td>
<td>-0.473</td>
<td>0.039</td>
</tr>
<tr>
<td>Paper and pulp (34)</td>
<td>1,119</td>
<td>-0.147</td>
<td>-0.178</td>
<td>0.452</td>
</tr>
<tr>
<td>Chemicals (35)</td>
<td>452</td>
<td>-0.113</td>
<td>-0.490</td>
<td>0.188</td>
</tr>
<tr>
<td>Mineral products (36)</td>
<td>385</td>
<td>-0.118</td>
<td>-0.319</td>
<td>0.270</td>
</tr>
<tr>
<td>Metals (37)</td>
<td>89</td>
<td>-0.138</td>
<td>-0.555</td>
<td>0.199</td>
</tr>
<tr>
<td>Machinery and equipment (38)</td>
<td>1,469</td>
<td>-0.007</td>
<td>-0.519</td>
<td>0.013</td>
</tr>
<tr>
<td>Other manufacturing (39)</td>
<td>126</td>
<td>-0.083</td>
<td>-0.450</td>
<td>0.156</td>
</tr>
<tr>
<td>Average</td>
<td></td>
<td>-0.086</td>
<td>-0.458</td>
<td>0.158</td>
</tr>
</tbody>
</table>

cross-section dimension is very large a first approach is to estimate the same model separately for each industry. In Table 4 we present the equilibrium correction estimates and the insider weight for ten industries at two digit level SIC code. These are obtained by running separate regressions identical to model (G) for each industry. All equilibrium correction estimates are negative and less than one in absolute

14 Detailed information is available from the author.
value. The insider weight varies from 1.3% in the machinery and equipment industry to 45.2% in the paper and pulp industry. The average insider weight (obtained from the average equilibrium estimates) is 15.8% which is significantly larger than in Table 2.

This procedure assumes, however, that the estimators are independent across industries. This is unlikely to be true given that wage bargaining is strongly coordinated between industries. Therefore, I cannot conclude that the pooled estimators are biased, but this is truly a matter for future research.

3.3. Compensating wage differentials and real wage resistance

The WALD tests for the joint significance of the time dummies in the models of Table 2 are very significant. However, the presence of time dummies makes it impossible to identify aggregate effects like real wage resistance and aggregate unemployment. In this section we drop the time dummies to explore these effects.

Bargaining theory predicts a falling locus in the wage-unemployment space. On the other hand, orthodox theory on migration and regional labour markets predicts a positive relation between wages and unemployment. In Harris and Todaro (1970) labour is mobile between regions and in the long run equilibrium wages must be higher in regions with high unemployment; this is the compensating (wage) differentials hypothesis. Our evidence so far is not conclusive with respect to this hypothesis. When we allow the unemployment elasticity to vary between regions, we find that the unemployment elasticity of wages varies between regions from -0.73 to 1.29 and that six out of 19 regional elasticities are negative.\textsuperscript{15}

The pattern of evidence changes when we let aggregate and regional unemployment interact as regressors. When we again impose identical coefficients on the regional unemployment elasticity, we would expect the aggregate unemployment elasticity of wages to be negative and the regional unemployment elasticity to be positive if the compensating wage differentials effect dominates. In Table 5 we report similar equilibrium correction models as Table 2, but with aggregate variables instead of time dummies as regressors.

When we include regional unemployment as well as aggregate unemployment, we find in model (H) a negative regional unemployment elasticity, while the aggregate unemployment elasticity is positive but insignificant. In all the models of Table 5 where the regional unemployment is included this elasticity is significantly negative. The conclusion is thus that the evidence does not support the hypothesis of compensating wage differentials. The aggregate unemployment elasticity is significant (with a negative sign) only when the regional unemployment variable is omitted.

Omitting the time dummies means we can also include wedge effects in our regression. In the models of Table 5 we put various restrictions not only on the

\textsuperscript{15} Estimates from model (F) of Table 2 with region specific unemployment elasticities of wages.
TABLE 5
GMM estimates of wage bargaining models specified as equilibrium-correction models. Endogenous variable is $\Delta w_{it}$; time dummies are not included.

<table>
<thead>
<tr>
<th>regressors</th>
<th>model</th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta w_{it(t-1)}$</td>
<td></td>
<td>-0.074 (9.08)</td>
<td>-0.074 (9.09)</td>
<td>-0.076 (9.31)</td>
<td>-0.072 (8.92)</td>
</tr>
<tr>
<td>$\Delta w_{it(t-2)}$</td>
<td></td>
<td>-0.023 (4.97)</td>
<td>-0.023 (5.00)</td>
<td>-0.024 (5.18)</td>
<td>-0.022 (4.93)</td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{it}$</td>
<td></td>
<td>0.251 (18.44)</td>
<td>0.240 (17.67)</td>
<td>0.243 (17.78)</td>
<td>0.202 (16.00)</td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{i(t-1)}$</td>
<td></td>
<td>-0.004 (0.91)</td>
<td>-0.005 (1.24)</td>
<td>-0.006 (1.34)</td>
<td>-0.009 (2.29)</td>
</tr>
<tr>
<td>$\Delta (p + y - n)_{i(t-2)}$</td>
<td></td>
<td>-0.006 (2.43)</td>
<td>-0.006 (2.51)</td>
<td>-0.006 (2.58)</td>
<td>-0.006 (2.85)</td>
</tr>
<tr>
<td>$\Delta \bar{w}_{it}$</td>
<td></td>
<td>0.342 (21.03)</td>
<td>0.358 (22.55)</td>
<td>0.371 (23.33)</td>
<td>0.417 (27.55)</td>
</tr>
<tr>
<td>$\Delta \bar{w}_{i(t-1)}$</td>
<td></td>
<td>0.003 (0.17)</td>
<td>0.024 (1.62)</td>
<td>0.037 (2.57)</td>
<td>0.078 (5.83)</td>
</tr>
<tr>
<td>$\Delta \bar{w}_{i(t-2)}$</td>
<td></td>
<td>-0.016 (1.17)</td>
<td>-0.002 (0.18)</td>
<td>-0.004 (0.31)</td>
<td>0.047 (3.93)</td>
</tr>
<tr>
<td>$s_{i(t-1)}$</td>
<td></td>
<td>-0.209 (15.69)</td>
<td>-0.205 (15.33)</td>
<td>-0.208 (15.48)</td>
<td>-0.186 (14.81)</td>
</tr>
<tr>
<td>$(w - \bar{w})_{i(t-1)}$</td>
<td></td>
<td>-0.351 (22.09)</td>
<td>-0.355 (22.32)</td>
<td>-0.347 (21.85)</td>
<td>-0.371 (24.10)</td>
</tr>
<tr>
<td>$\Delta (cp - p)_{kt}$</td>
<td></td>
<td>0.064 (4.06)</td>
<td>0.057 (3.64)</td>
<td>0.064 (4.07)</td>
<td>0.036 (2.56)</td>
</tr>
<tr>
<td>$\Delta (prt - tax)_{kt}$</td>
<td></td>
<td>0.214 (6.14)</td>
<td>0.186 (5.49)</td>
<td>0.210 (6.28)</td>
<td>0.048 (1.95)</td>
</tr>
<tr>
<td>$(cp - p)_{kt}$</td>
<td></td>
<td>0.016 (1.05)</td>
<td>0.010 (0.67)</td>
<td>-0.002 (0.11)</td>
<td>0.036 (2.56)</td>
</tr>
<tr>
<td>$(prt - tax)_{kt}$</td>
<td></td>
<td>-0.358 (7.77)</td>
<td>-0.306 (6.92)</td>
<td>-0.315 (7.18)</td>
<td>0.036 (2.56)</td>
</tr>
<tr>
<td>$u_{it}$</td>
<td></td>
<td>-0.015 (4.21)</td>
<td>-0.011 (5.94)</td>
<td>-0.015 (4.35)</td>
<td>-0.019 (6.42)</td>
</tr>
<tr>
<td>$\Delta u_{it}$</td>
<td></td>
<td>0.020 (6.27)</td>
<td>0.010 (5.91)</td>
<td>-0.015 (4.35)</td>
<td>0.019 (6.42)</td>
</tr>
<tr>
<td>$u_{t}$</td>
<td></td>
<td>0.007 (1.12)</td>
<td></td>
<td>-0.015 (4.35)</td>
<td>0.019 (6.42)</td>
</tr>
<tr>
<td>$\Delta u_{t}$</td>
<td></td>
<td>-0.018 (3.61)</td>
<td></td>
<td>0.010 (3.85)</td>
<td>0.014 (5.43)</td>
</tr>
</tbody>
</table>

WALD (df)    7,974 (18)  7,899 (16)  7,894 (16)  7,779 (14)  7,779 (14)
SARGAN (df)  743 (227)  744 (227)  771 (227)  810 (227)  810 (227)
AR(1)        -34.1     -34.1     -34.2     -33.9     -33.9
AR(2)        -0.7      -0.6      -0.6      -0.1      -0.1
$\hat{\sigma}^2$ 0.021  0.021  0.021  0.021  0.021

Notes: see Table 1.

unemployment variables, but also the wedge. The corresponding long-run solutions to models (H)–(K) are:

model (H): $w = 0.373(p + y - n) + 0.627\bar{w} - 0.027u_j$
\[ (12.31) \quad (4.63) \quad (4.19) \]
\[ + 0.013u + 0.028(cp - p) - 0.639(prt - tax) \quad (11) \]

model (I): $w = 0.366(p + y - n) + 0.634\bar{w} - 0.020u_j$
\[ (12.08) \quad (14.72) \quad (5.92) \]
\[ + 0.018(cp - p) - 0.548(prt - tax) \quad (12) \]

model (J): $w = 0.375(p + y - n) + 0.625\bar{w}$
\[ (12.14) \quad (14.49) \]
\[ - 0.026u - 0.003(cp - p) - 0.568(prt - tax) \quad (13) \]
model (K):  \[ w = 0.333(p + y - n) + 0.667 \bar{w} - 0.035 u \]  
(11.77)  
(15.32)  
(6.35)  

The partial unemployment elasticity varies between \(-0.020\) and \(-0.035\). Taking into account the effect of unemployment on alternative wages by setting \(w = \text{constant} + \bar{w}\) in the long run, we obtain the total unemployment elasticity which is comparable to studies on aggregate data. The total unemployment elasticity is \(-0.105\) in model (K) and between \(-0.054\) and \(-0.073\) in the other models. This is in line with several studies from a range of countries reported in Blanchflower and Oswald (1995) where the unemployment elasticity is around \(-0.10\). Regional unemployment is strongly correlated with the time dummies. Hence, including time dummies as in Table 2 might leave too little variation in the regional unemployment variable to identify the unemployment elasticity.

The significant effect of the change in the wedge indicates the existence of short-run real wage resistance. But there is no evidence of real wage resistance in the long run. The unexpected significant negative effect of the tax wedge may be the result of the proxy we are using for the income tax rate being too crude.

More importantly, the omission of the time dummies has increased the insider equilibrium correction term and the insider weight significantly. In models (H)–(K) the insider weight varies within the range of 33.3%–37.4% compared to around 4.2% in the models with time dummies. One possible explanation is that the time dummies control for any proportional change in the composition of skills in the work force which is likely to affect the wage per employee. Increasing importance of skilled labour over time will yield a positive bias in the insider weight due to productivity gains from enhanced skills. There may also be other aggregate shocks that are strongly correlated with productivity, hence the estimate of the insider weight in Table 5 should be interpreted as a gross effect taking into account the effect of the omitted time dummies.

Probably a more important explanation is that aggregate shocks to the productivity or the wage share, is followed by an equilibrium correction that is common to all firms and is captured by the time dummies. The common equilibrium correction might take place e.g. through central wage bargaining. Hence, omitting the time dummies necessarily (and correctly) increases the estimated speed of adjustment to the wage share. However, it is in the models of Table 2 where we control for aggregate equilibrium correction by including time dummies, that the insider weight, to a greater extent, will reflect firm specific variation only.

4. Conclusions

Using annual observations on more than 7,300 firms over a period of 17 years, we have investigated the relative importance of insider vs outsider variables in wage formation in the Norwegian manufacturing sector. Equilibrium correction models perform better at explaining wage setting in Norway than do
traditional empirical specifications that explain wage setting in the UK using panel data. It is thus important to model the dynamic mechanisms in panel data.

To identify the insider weight it is important to include time dummies to control for aggregate equilibrium correction. I find a positive and significant insider weight of about 5%, which is considerably lower than the one found by Johansen (1996), but slightly larger than the insignificant insider weight of Holmlund and Zetterberg (1991) both using Norwegian industry level data. The estimate of the insider weight is smaller than comparable estimates from less centralised economies such as the UK. Thus the results seem consistent with the view that the Norwegian wage setting system is more centralised than the one in the UK. There is only weak evidence of an increase in the insider weight arising from the decentralisation of bargaining institutions in the 1980s. Separate regressions for different industries indicate that the pooled estimate of the insider weight may be biased, but this is a matter for future research.

I do not find evidence that regions with relatively high unemployment have relatively high wage levels, lending little support to the hypothesis of compensating wage differentials. The unemployment elasticity is found to be negative and equal to \(-0.10\) in certain specifications. This wage flexibility does not stand out as very high compared to other countries.

The wedge between real labour costs and real consumption wages does not have long-run effects on wage costs, but I find some evidence of short-run effects, which may be interpreted as short-run real wage resistance.

ACKNOWLEDGEMENTS

This paper is a part of the project 'Unemployment, institutions and policy' at Centre for Research in Economics and Business Administration (SNF). It was partly written while I was staying at the Centre for Economic Performance, London School of Economics, and I would like to thank the Centre for their tremendous hospitality. I have benefited from discussions with and comments from Charlie Bean, Danny Blanchflower, Richard Dickens, Steinar Holden, Steve Machin, Alan Manning, Lucy Matthew, Steve Nickell, Ragnar Nymoen, Andrew Oswald, Eva Pichler, Asbjørn Radseth, and four anonymous referees. No-one but myself is responsible for the errors or the conclusions. I am also in debt to Norwegian Social Sciences Data Service and Jan M. Dyrstad for providing some of the data and Tungregnevalget, Norwegian Research Council, for financial support.

REFERENCES


APPENDIX I

A right-to-manage model

This presentation is similar to Nickell and Wadhwani (1990); more explicit and thorough presentations of wage bargaining models are found in Layard et al. (1991), Nickell and Kong (1992), Lee and Pesaran (1992), and Manning (1987).

Assuming that the union cares about the utility of a representative union member and that he or she cares about the wage, the union's objective is to maximise

\[ V_i = V_i\left(\frac{W_i}{P_i\Phi_i}, N_i, A_i\right) V_i > 0 \]  \hfill (15)

where \( V_i \) is union utility, \( W_i \) is wage cost including payroll taxes, \( N_i \) is employment, \( P_i \) is the producer price, and \( A_i \) is outside option. \( \Phi_i \) is the wedge between real labour costs and real after-tax consumption wages, so that \( W_i/(P_i\Phi_i) \) is the real after tax consumption wage. Union utility is increasing in all arguments.

An increase in \( A_i \) raises the utility of a worker who is laid off, thus it also raises the expected utility of the union, \( V_i \). \( A_i \) depends on the size of the unemployment benefit, wages in alternative employment (alternative wage), and the rate of unemployment, the latter influencing the probability of obtaining alternative employment. Hence \( A_i \) can in turn be written as a function

\[ A_i = A_i(\bar{W}_i, U, B) \]  \hfill (16)

where \( U \) is the unemployment rate, \( \bar{W}_i \) is the alternative wage cost including payroll taxes, and \( B \) are unemployment benefits. A competitive firm sets employment given the outcome of the wage negotiations to maximize profits

\[ \Pi_i = R_i(Q_iE_iN_i) - W_iN_i \]  \hfill (17)

\( R_i \) is the revenue function allowing for efficiency wage effects through the efficiency function \( E_i = E(W_i/\bar{W}_i, U) \), while \( Q_i \) is a productivity parameter.

The first order condition for employment is

\[ \frac{\partial R_i}{\partial N_i} = \frac{W_i}{E(W_i/\bar{W}_i, U)Q_i} \]  \hfill (18)

Thus the maximum profit can be written as a function of wage costs and exogenous variables that affect efficiency

\[ \Pi_i = \Pi_i\left(\frac{W_i}{Q_iF_i}, U, \frac{W_i}{\bar{W}_i}\right) \]  \hfill (19)

The wage is assumed to be the solution to the Nash bargaining problem, subject to the labour demand condition (18)

\[ W_i = \text{argmax} \ [\ (V_i - V_i^0)\beta (\Pi_i - \Pi_i^0)] \]  \hfill (20)

\( \beta \) is the relative bargaining power of the union and \( (V_i^0, \Pi_i^0) \) is the disagreement point which according to strategic bargaining theory, is the parties' pay-offs during negotiations until a new agreement is reached (Binmore et al. 1986). The constraints imply that the bargaining outcome has to be as least as big as the parties' outside options; \( A_i \) and \( \bar{V}_i \) respectively. Assuming an interior
solution the first order condition of (21) is
\[ \beta \frac{\partial V_i/\partial w_i}{V_i - V_i^1} + \frac{\partial \Pi_i/\partial w_i}{\Pi_i - \Pi_i^1} = 0 \]  \hspace{1cm} (21)
which implicitly defines the wage function
\[ W_i = W_i(\beta, \bar{W}_i, U, B, Q_i, \Phi_i) \]  \hspace{1cm} (22)

APPENDIX 2

Data

Subscripts \(i, j, k, \) and \(t\) denote firm, region, sector and period respectively. There are \(N = 7,323\) firms with different observation periods given by

\[
\begin{array}{cccccccc}
T_i & 10 & 11 & 12 & 13 & 14 & 15 & 16 & 17 \\
N_i & 392 & 401 & 367 & 339 & 337 & 378 & 768 & 4,341 \\
\end{array}
\]

The variables are:
- \(n_i\): Average yearly employment computed as the average of employment measured in February, April, June, September and November. Source: The Annual Manufacturing Statistics, Statistics Norway.
- \(w_{it}\): Wage bill including payroll taxes divided by average employment. Source: The Annual Manufacturing Statistics, Statistics Norway.
- \(\bar{w}_{it}\): Alternative wage cost. Average wage cost divided by average employment outside the firm in the same labour district. Own computations based on \(w_{it}\).
- \(prt_i\): The regional payroll tax. Source: J. M. Dyrstad, University of Trondheim.
- \(cp_i\): Consumer price index. Source: Statistics Norway.
- \(u_{it}\): Registered unemployment. Own computation based on data from Norwegian Social Science Data Service and Statistics Norway.
- \(p_{ki}\): Producer price index for sector \(k\) (five digit level, SIC). Source: Statistics Norway.
- \(tax_i\): Income tax rate. Social security contributions plus other direct taxes as a percentage of GDP. Source: Statistics Norway.