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REGIONAL WAGE CURVES EMPIRICAL EVIDENCE FROM NORWAY

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Regional Wage Curves
Empirical Evidence from Norway*

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Abstract

The papers studies wage formation using panel data for a large sample of Norwegian municipalities covering the time period 1970–1992. The main conclusions are the following. Regional manufacturing wages are negatively related to regional unemployment, but the effects of open as well as total regional unemployment are small. We find a numerically important wage dampening effect of labour market programs. However, we can not reject the null that only total unemployment matters. Finally, manufacturing profitability is important in shaping regional manufacturing wages and reduced regional payroll taxes are almost fully transmitted into reduced regional wage costs.

Keywords: Panel Data; Wage Curve; Labour Market Programs
JEL Classification: C23; J30; J60

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1 Introduction

The main issue in the present paper will be to test whether or not Norwegian regional wages are affected by regional unemployment and labour market programs (LMP). The empirical analysis is based on annual panel data for 322 municipalities for the time period 1970 to 1992.

The question of regional wage responsiveness to regional unemployment is closely related to the analyses by Blanchflower and Oswald (1994) who estimate regional wage curves for a number of different countries. Using data for individuals, they provide evidence of a downward sloping wage curve. For most countries the estimated unemployment elasticity of pay cluster around –0.10. Further evidence of a downward sloping wage curve for Denmark is given by Nicolaisen and Tranæs (1996), for Belgium by Janssens and Konings (1998), for the Netherlands by Groot *et al.* (1992), for Germany by Baltagi and Blien (1998) and Wagner (1994), for Austria by Winter-Ebmer (1996), and for the US by Card (1995), Bratsberg and Turunen (1996) and Turunen (1998).

Blanchflower and Oswald (1994) also report results for Norway supporting a downward sloping regional wage curve, but the estimates are highly mixed and depend upon the exact specification. Wulfsberg (1997) and Raaum and Wulfsberg (1998) report rather low partial long-run elasticities of firm level wages with respect to open county unemployment. The results in Dyrstad and Johansen (2000), using panel data for Norwegian municipalities 1973–88, imply a long-run *regional* unemployment elasticity of pay between –0.01 and –0.02 while the elasticity with respect to *aggregate* unemployment approximates –0.07. Interestingly, the sum of the unemployment elasticities is close to –0.1. One interpretation of this result is that Norwegian wages mainly respond to labour market conditions through the central wage settlements while the effect working through local wage setting is more limited.

The present paper is partly an update of Dyrstad and Johansen (2000) and utilise data for four new years characterised by relatively high and increasing unemployment. A particular issue will be to investigate whether or not the estimated unemployment effect depends upon how regional labour markets are defined. Since most Norwegian municipalities are very small, one may argue that the state of the labour market outside the same municipality plays an important role. On the other hand, county may represent a too wide definition of the relevant regional labour market, in particular if the within-county unemployment dispersion is large and mobility is low. In this paper we therefore report results using three different definitions of regional labour markets: Municipality, County and Local Labour Office Area (LLOA).[1]

Furthermore, we make an investigation of the shape of the wage curve. Blanchflower and Oswald (1990, 1994) provide evidence in favour of a more convex wage curve than the conventional log-linear specification. Results for Norway based on aggregate time series data (Johansen, 1995, 1997) and industry panel data (Johansen, 1996, 1999) suggest that the Norwegian aggregate wage curve is highly nonlinear. The present paper makes an investigation of the shape of the Norwegian *regional* wage curve.

One possible objection to the results in Dyrstad and Johansen (2000) is biasedness because participants on LMP are not included. Raaum and Wulfsberg (1998) report that higher participation rate on LMP significantly reduces regional wages. Similar evidence for Sweden can be found in Edin *et al.* (1994). Further, the number of participants are positively correlated with open unemployment across regions. In this paper we therefore re-estimate the regional wage equation including information on LMP.

The paper also highlights the role of regional manufacturing profitability in shaping regional manufacturing wages. This issue is related to the studies of insider forces in wage determination like Nickell and Wadhwani (1990), Nickell and Kong (1992) and Nickell *et al.* (1994) who report significant long-run effects of firm or industry level profitability on firm or industry level wages. The empirical evidence regarding insider effects on Norwegian wages is mixed. Holmlund and Zetterberg (1991) find small and statistically insignificant effects of industry prices and productivity on industry wages while the long-run insider weight reported in Johansen (1996, 99) approximates 20%. Using panel data for Norwegian manufacturing firms, Wulfsberg (1997a) reports a statistically significant insider weight of 5% while the results in Raaum and Wulfsberg (1998) imply a long-run insider weight above 30%. Finally, Dyrstad and Johansen (2000) report a long-run elasticity of wages with respect to value added per worker slightly above 10% based on regional manufacturing data.

Payroll taxes have been regionally differentiated (since 1975) in order to stimulate regional employment in depressed areas. How such a policy works depends among other things on the wage responses to a tax cut. If reduced payroll taxes are transmitted into higher regional earnings, the main effect will be an income redistribution while the effects on regional employment and unemployment will be small. An important policy issue is therefore to estimate how regional wage costs are affected by changes in regional payroll taxes, both in the short run and in the long run.

The rest of the paper is organised as follows. Section 2 gives a brief presentation of the empirical model. The main results are reported in Section 3 while concluding comments are given in Section 4.

2 Data and empirical specification

The theoretical basis for the analysis is a standard firm level bargaining model, see Nickell and Andrews (1983) and Hoel and Nymoen (1988). Blanchflower and Oswald (1994, Ch. 3) offer a survey on theoretical models predicting a downward sloping wage curve. The starting point for the empirical study is an equilibrium correction model which contains feedback effects of region specific variables, while any effects of aggregate variables, common to all municipalities, will be captured by a full set of time dummies.

The study uses a panel of annual time series data from 322 Norwegian municipalities covering the time period 1970–92.[2] The regional wage variable, W_j, is manufacturing wage costs per worker, including payroll taxes. Our measure of regional profitability, P_j, is manufacturing value added at factor prices per worker. We include the regionally differentiated payroll tax rate, TP_j, and lagged levels of employment, N_{jt-1}, to test for insider (or membership) hysteresis effects. Three alternative definitions of open regional unemployment will be used: the municipality unemployment rate, U_j, the county unemployment rate, U_c, and finally the unemployment rate in the LLOA, U_l.

The benchmark regional wage equation to be estimated below is given by

$$\begin{aligned}\Delta w_{jt} &= -\alpha w_{jt-1} + \beta_1 p_{jt-1} + \beta_2 u_{it-1} + \beta_3 tp_{jt-1} + \beta_3 n_{jt-1} \\ &\quad + \gamma_1 \Delta p_{jt} + \gamma_2 \Delta u_{it} + \gamma_3 \Delta tp_{jt} + \eta_t + \eta_j + \varepsilon_{jt}, \qquad (1)\end{aligned}$$

where lowercase letters denote natural logs and $tp = \ln(1 + TP)$, subscript $i = j, c, l$ reflects alternative definitions of the regional unemployment rate, η_t are the common time specific effects, η_j are time invariant municipality specific effects while ε_{jt} are the remaining error terms, assumed to be $\mathsf{iid}(0, \sigma_\varepsilon^2)$.

To test for nonlinearity we follow Blanchflower and Oswald (1990) and expand equation (1) with the cube of log regional unemployment. We have information on LMP only at the county level, and only for the time period 1980–92. For this period we re-estimate the wage equation including total county unemployment, TU_c, and the ratio of open to total county unemployment, U_c/TU_c, as separate variables.[3]

3 Empirical results

3.1 Results for the benchmark model

Since the dynamic wage equation (1) includes municipality specific effects, the within-groups estimators are biased even if the residuals are white noise.

Furthermore, value added per worker and employment should be regarded as endogenously determined and potentially correlated with the municipality specific effects. The same may be true for regional unemployment. In this case we report results given alternative assumptions of the exogeneity status of the unemployment variable.

To obtain consistent estimators we apply a system GMM estimator suggested by Arellano and Bover (1995) and Blundell and Bond (1998). The system GMM estimator uses equations in first-differences, from which the municipality specific effects are eliminated by the transformation. In these equations endogenous variables lagged two or more periods are valid instruments provided there is no serial correlation in the time-varying component of the error terms. The differenced equations are combined with equations in levels. In these equations the instruments must be orthogonal to the municipality specific effects. As shown in Blundell and Bond (1998), first-differences of variables can be uncorrelated with the individual specific effects even if the levels are correlated with the effects.

Table 1 report results where w_{jt-1}, Δp_{jt}, p_{jt-1} and n_{jt-1} are treated as endogenous.[4] The set of instruments are defined in note iii to the Table. It should be noted that regional unemployment is assumed strictly exogenous and also uncorrelated with the municipality specific effects.

Equation I contains the municipality unemployment rate while equation II contains county unemployment and equation III the unemployment rate in the LLOA. The short-run impact effect of regional unemployment is small for all specifications and statistically insignificant in equation II and III. The estimated level effect is significantly below zero in all equations and robust with respect to specification.[5] The estimate of lagged regional wages is significantly below zero which means that the Phillips curve specification is rejected. Using the results in Table 1 we find that the estimated long-run unemployment elasticity of pay ranges from –2.0% in equation III to –2.7% in equation II.[6] These estimates are well below the estimates reported in Blanchflower and Oswald (1994) and confirm the finding in Dyrstad and Johansen (2000) that regional wage responses to regional labour market conditions are rather weak. Interestingly, Nicolaisen and Tranæs (1996), using Danish data for individuals linked to county unemployment, report an unemployment elasticity of –2%. Furthermore, Holmlund and Skedinger (1990), estimating wage drift equations on data from the Swedish wood industry, also obtain negative but small effects of regional unemployment. The long-run elasticity with respect to county unemployment is –1.6%.

In order to discriminate between the three different specifications we estimated a general model which contains the three unemployment rates (in first differences and lagged levels). The individual coefficients of Δu_{ct}, Δu_{lt}, u_{ct-1},

u_{lt-1} were all statistically insignificant while the individual coefficients of Δu_{jt} and u_{jt-1} were both significantly below zero. A test of the joint null hypothesis that the coefficients of Δu_{ct}, Δu_{lt}, u_{ct-1}, u_{lt-1} are all equal to zero yields $\chi^2(4) = 5.79$ with a p-value of 0.22. On the other hand, a test of the null that the coefficients of Δu_{jt} and u_{jt-1} are equal to zero yields $\chi^2(2) = 15.59$ which is highly significant.

Table 1 about here

The main finding so far is that the wage responses to regional labour market tightness are weak, a conclusion that seems robust with respect to the alternative definition of regional labour market regions. The small effects of regional unemployment may reflect a mix of a downward sloping wage curve and a positive unemployment effect due to compensating wage differentials, cf. Harris and Todaro (1970).[7] As noted by Blanchflower and Oswald (1994) the Harris-Todaro concept of compensating differentials and a downward sloping wage curve may well coexist. Because migration is costly, migrants' choices are not likely to respond to transitory shocks to regional unemployment. It is the *permanent* or expected values of regional wages and regional unemployment that might be positively correlated in a long-run equilibrium. Since it seems likely that municipality fixed effects will capture any effects of permanent unemployment differences, we made a comparison of the within groups (WG) estimates with the plain OLS estimates. For all three unemployment rates, the absolute value of the OLS estimates were larger than the corresponding WG estimates.[8] Although this experiment should not be interpreted as a formal test of the hypothesis of compensating wage differentials, the results are not supportive to this hypothesis. However, the differences between the OLS and WG estimates may reflect an effect of left out variables that are correlated with regional unemployment. One obvious candidate is regional consumer prices.

Both the short-run impact effect of value added per worker, and the long-run level effect are well determined in all equations. The long-run insider weight approximates 23%. This estimate implies that the effect of a hypothetical change in value added per worker from the sample mean to its maximum value would increase regional wages by 64%.

The estimated insider weight reported in Table 1 is higher than the corresponding estimates in Dyrstad and Johansen (2000). On the other hand, the new estimates are largely in accordance with results based on panel data for Norwegian industries reported in Johansen (1996, 1999). A positive long-run relation between regional wages and regional manufacturing profitability is evidence against competitive forces as well as completely centralised wage setting. Our interpretation is that such permanent effects reflect wage bargaining at the firm level.

We find no evidence of an insider hysteresis effect as lagged values of regional manufacturing employment, n_{jt-1}, always enter with a positive sign, and the estimates are statistically significant from zero. As an alternative to lagged employment levels we included current values of employment changes, Δn_{jt}, properly instrumented. The estimates of Δn_{jt} were always negatively signed and statistically insignificant from zero confirming the results in Dyrstad and Johansen (2000).

The short-run impact elasticity of regional payroll taxes vary from 0.8 to 0.9 while the long-run elasticity vary from 0.85 to 0.96. All estimates are statistically significant from zero and never statistically different from unity. The estimated long-run effects are largely in accordance with the results reported in Dyrstad and Johansen (2000) and Wulfsberg (1997b). Reduced regional payroll taxes are mainly transmitted into reduced wage costs. Hence, the regionally differentiated payroll tax system seems to be an effective policy instrument, given that labour costs are important for labour demand. In fact, preliminary results based on the same data set imply a long run elasticity of municipality manufacturing employment with respect to wages approximately equal to –1.

Table 1 reports several diagnostic test statistics. First, the Arellano and Bond (1991) test of no second-order serial correlation in the differenced residuals, $AR2$, is always below critical values. On the other hand, the $AR1$ statistic clearly indicate negative first-order serial correlation. Taken together these results imply that the levels of the error terms are white noise.

Sargan[232] is the Sargan (1958, 1988) test of over-identifying restrictions. This is a joint test of the validity of all orthogonality restrictions. Sargan diff1[2] tests the validity of treating regional unemployment as exogenous and uncorrelated with the municipality specific effects while Sargan diff2[84] tests the validity of the orthogonality restriction for the levels equations.[9] Finally, two Hausman-type tests are reported. Hausman[8] tests the null that all parameters are equal in the equations in levels and first differences while Hausman[2] tests the null that the coefficients of regional unemployment are equal across these equations. The two Hausman tests enter with a p-value slightly above 5% in equation I. In all other cases all tests are well below their corresponding critical values.

Although all tests reported in Table 1 look comfortable we investigated the sensitivity of the results with respect to different estimators. First, we re-estimated the wage equations in Table 1 using the GMM-difference estimator proposed by Arellano and Bond (1991). Second, we excluded Δu_{it} and u_{it-1} from the set of instruments, taking endogeneity and potential correlation with municipality effects into account. The main result of interest was that the absolute value of the estimated unemployment effects were smaller for all alternative estimators as compared with their corresponding counterparts reported in Table 1. This evidence suggest that the results in Table 1 do not under-estimate the absolute value of the true regional unemployment coefficient.

3.2 Nonlinear unemployment effects

The next issue will be to further investigate the shape of the wage regional curve. In order to test for more complicated nonlinearity we follow Blanchflower and Oswald (1994) and expand the benchmark wage equation with the cube of log unemployment. The estimated wage equation is then given by

$$\begin{aligned} \Delta w_{jt} \;=\; & -\alpha w_{jt-1} + \beta_1 p_{jt-1} + \beta_{20} u_{it-1} + \beta_{21} u_{it-1}^3 + \beta_3 tp_{jt-1} + \beta_3 n_{jt-1} \\ & + \gamma_1 \Delta p_{jt} + \gamma_2 \Delta u_{it} + \gamma_3 \Delta tp_{jt} + \eta_t + \eta_j + \varepsilon_{jt}, \end{aligned} \tag{2}$$

where we expect β_{21} to be positive.

The results reported in Table 2 show that the estimate of the cube of log regional unemployment is positive for all alternatives and statistically significant in equation I and II. The estimated coefficients of u_{it-1}^3 are small, in particular for municipality unemployment. Nevertheless, the results provide strong evidence of a more convex regional wage curve than implied by the usual log-linear specification.

Table 2 about here

Figure 1 graphs the long-run regional wage curves based on the results in Table 2. The wage level is normalised to unity for regional unemployment rates equal to 1%. The wage curve based on county unemployment is U-shaped and minimises at an unemployment rate equal to 4.6%. The other two are monotonically decreasing for unemployment rates within their respective sample ranges.[10]

Figure 1 about here

We also estimated the wage equations in Table 2 using data for the sub-sample 1981–1992 which is the sample period used in the next section. For this sub-sample we still found a positive and significant estimate of the cubic term using data for municipality unemployment. However, the cubic term was highly insignificant for the other regional unemployment rates. This result may reflect the fact that there are no extremely low unemployment rates at these levels of aggregation during the period 1981–1992 and that the log-linear specification works well when unemployment is not too low.

3.3 Labour market programs

Finally, we investigate the effects of labour market programs. Standard models of wage bargaining predict ambiguous wage effects of expanding LMPs, see Calmfors and Lang (1995) for a formal discussion. On the one side workers may prefer LMPs to open unemployment. Expanding programs will therefore reduce the expected welfare loss of being laid off. The partial effect is higher wages. On the other hand, LMPs may have positive effects on the re-employment prospect for participants through enhancement of skills, motivation and labour force participation. In this case LMPs will increase search effectiveness of the not-employed workers which implies increased job competition and therefore reduced wages.

The empirical specification is based on recent work by Rødseth and Nymoen (1999) and Raaum and Wulfsberg (1998) and reads

$$\begin{aligned}\Delta w_{jt} &= -\alpha w_{jt-1} + \beta_1 p_{jt-1} + \beta_2 tp_{jt-1} + \beta_3 tu_{ct-1} + \beta_4 (u - tu)_{ct-1} \\ &\quad +\gamma_1 \Delta p_{jt} + \gamma_2 \Delta tp_{jt} + \gamma_3 \Delta tu_{ct} + \gamma_4 \Delta (u - tu)_{ct} \\ &\quad +\eta_j + \eta_t + v_{jt}, \qquad (3)\end{aligned}$$

where tu is the log of total unemployment (the sum of openly unemployed and program participants divided by the labour force) while u is the log of the rate of open unemployment. This specification enable us to test two specific hypotheses. The first one is that only total unemployment matters which means that $\beta_4 = 0$. In this case an expansion of LMP will not affect wages when the participants are recruited from open unemployment while wage pressure is reduced if they are recruited from regular employment. The second hypothesis is that only open unemployment matters which means that $\beta_4 = \beta_3 < 0$. In this case an expansion of LMP will increase wages if the participants are recruited from open unemployment while there is no wage effect if they are recruited from regular employment (programs are perfect substitutes to regular employment). A positive value of β_4 means that expanding programs will reduce wages even if participants are recruited from open unemployment. Finally, in the cases where $\beta_4 < \beta_3$, programs will increase wages even if the participants are recruited from regular employment. The same comments also apply for the short-run impact effects.

We have data for LMP only at the county level and only for the period 1980–92. Equation (3) is therefore estimated including county unemployment for this period. Table 3 reports the results. We first note that all diagnostic test statistics look comfortable for all equations. In order to test for parameter stability we first re-estimate equation II in Table 1 using data for the sub-sample. The results in the first column of Table 3 reveal that the level effect of county unemployment is higher for this sub-sample period as compared with the full sample estimates. The absolute value of the long-run unemployment elasticity of pay increases from 2.7% to 3.6%. One possible explanation is that reduced time period makes the cross-section variation more important which may affect the estimates. We therefore re-estimated the equation using data for the period 1973–80. Based on data for this period we obtained a long-run unemployment elasticity equal to 2%. These results therefore indicate increasing regional wage flexibility over time, possibly because firm level bargaining have become more important during our sample period, cf. Holden (1989) and Rødseth and Holden (1990) for institutional evidence.

Column II reports results based on the unrestricted version of equation (3). The short-run impact effect of total unemployment is negatively signed but statistically insignificant while the long-run level effect is significantly below zero. The absolute value of the long-run elasticity is slightly higher than the corresponding estimate for open unemployment reported in column I. The estimated effects of $\Delta (u - tu)_{ct}$ and $(u - tu)_{ct-1}$ are both positively signed. An increase in the program ratio (reduced value of $u - tu$) will therefore reduce wage pressure both in the short run and in the long run even if the program participants are recruited from open unemployment.

Table 3 about here

To illustrate the effects of higher unemployment and expanding programs we consider the following experiments. Assume that total unemployment is initially equal to 5%, that the rate of open unemployment is 4% while 1% of the labour force is allocated to labour market programs. In case one we increase the rate of open unemployment by 1 percentage point holding the participation rate constant. Given the estimates in Table 3 this will induce a long-run wage moderation of 0.42%. In case two we increase the participation rate by 1 percentage point holding the rate of open unemployment constant. In this case the long-run wage moderation is 2.11%. In case three, total unemployment is constant while the participation rate is increased by 1 percentage point. That is, 1% of the labour force is allocated from open unemployment to LMP. Such a policy would reduce wages by 2.18% in the long run.

One possible explanation of the strong numerical effects of LMP may be that programs have been targeted towards the long-term unemployed, and that labour market training programs have particularly favourable effect on search effectiveness for this group. However, it should be noted that the estimates of $\Delta(u-tu)_{ct}$ and $(u-tu)_{ct-1}$ are not statistically significant from zero, neither partially nor jointly, and we can not statistically reject the null hypothesis that only total unemployment matters. The imprecise estimates of $\Delta(u-tu)_{ct}$ and $(u-tu)_{ct-1}$ partly reflects an identification problem since LMPs have been targeted towards regions with high unemployment, and the correlation coefficient between tu_{ct} and $(u-tu)_{ct}$ is –0.5.

Column III reports results for a restricted version of the equation which excludes Δtu_{ct}, $\Delta(u-tu)_{ct}$ and $(u-tu)_{ct}$. These restrictions are clearly not rejected by data as the test yields $\chi^2(3) = 2.83$ with a p-value of 0.50. The estimated elasticity of total unemployment is almost entirely unaffected by imposing these restrictions. To conclude, we therefore find some evidence in favour of a wage dampening effect of LMP but these are mainly captured by total unemployment.

How do these results compare to those obtained by others? In a recent time series study of wage formation in the Nordic countries, Rødseth and Nymoen (1999) report negatively signed and statistically significant long-run effects of total unemployment for all the Nordic countries (i.e. Denmark, Finland, Norway and Sweden). The estimates of $\Delta(u-tu)$ and $(u-tu)$ are negatively signed but statistically insignificant from zero with one exception: The short-run impact effect is positive and significant for Norway which implies a short-run wage dampening effect of expanding LMP. However, their results suggest that it is the rate of open unemployment that is the fundamental long-run labour-market tightness variable in all countries as the hypothesis that $\beta_4 = \beta_3$ is not rejected. Similar results can be found in Holden and Nymoen (1998) based on the same data set.

Raaum and Wulfsberg (1998) report results for different specifications using firm level panel data linked to labour market variables at the county level. Their Model B contains tu and $u-tu$ in first differences and lagged levels as in our specification. This specification includes average wages at the county level but no aggregate variables. Both the short-run impact effect and the level effect of total unemployment are negatively signed and statistically significant. The short-run impact effect and level effect of $u-tu$ are positive and highly significant. However, all estimates of local labour market variables are strongly reduced when either aggregate unemployment or time dummies are included. In these specifications no local labour market variables enter significantly.

Forslund (1993) uses Swedish panel data for 12 unemployment insurance funds and allows for different effects of labour market training programs and relief work. The estimated wage effect of manpower training is negatively signed, statistically significant and close to the estimated effect of open unemployment. On the other hand, the effect of relief work is positive and bordercase significant.[11] Edin *et al.* (1994) offer a study of the relationship between regional labour market variables and wage pressure in Sweden using a large micro data set. Their findings suggest that manpower training programmes reduce wages while the estimated effect of relief work depends upon the exact specification. However, their results never implies a significant wage hike from relief work. Pannenberg and Schwarze (1998) use micro data for East Germany 1992–94 linked to regional labour market indicators. They provide evidence in favour of an inverse relation between wages and the total regional unemployment rate while the composition between open unemployment and program participants is of little importance.[12]

Finally, we test whether or not the effect of labour market programs depends upon the state of the regional labour market. Expanding labour market programs may induce a wage hike in regions and time periods with initially low levels of unemployment and wage moderation in regions and periods with sufficiently high unemployment. To test this hypothesis we re-estimate equation I in Table 3 including open unemployment interacted with $(u_c - tu_c)$, both in levels and first differences. The results reported in Table 4, column I, are in general supportive to the hypothesis. The estimates of $(u - tu)_{ct-1}$ and $\Delta(u - tu)_{ct}$ are both negatively signed but statistically insignificant from zero. Both interaction terms are positive and bordercase significant. In column II we report results for a restricted version which excludes $(u - tu)_{ct-1}$, Δtu_c and $\Delta(u - tu)_{ct}$. The estimated effect of the remaining variables are almost entirely unaffected by imposing these restrictions.

Table 4 about here

Turning to the other results reported in Table 3 and 4 we note that the estimated effects of value added per worker are slightly below their corresponding full-sample estimates. The estimated long-run insider weight approximates 17% using data for the last sub-sample while the full-sample estimate is 23%. Also, the short-run impact effect of regional payroll taxes is smaller for the sub-sample. The long-run elasticity is close to unity in column I and III in Table 3 but somewhat smaller in column II – and in Table 4 – which include the participation rate as a separate variable. These differences may reflect a negative correlation between LMP and regional payroll taxes.[13] Excluding programmes may therefore bias the payroll tax effect upwards. Finally, the insider hysteresis term is positive and still statistically significant from zero.

4 Concluding comments

The main issue in this paper has been to analyse empirically how Norwegian regional wages are effected by regional specific shocks to unemployment. A high degree of regional wage responsiveness to regional unemployment implies strong equilibrating mechanisms which dampen unfavourable real effects of asymmetric shocks. Vigorous wage responsiveness leads to low unemployment equilibria while regional wage stickiness implies regional unemployment persistence.

The results suggest that regional wage responses to open regional unemployment are small. The full sample estimates based on a system GMM estimator imply a long-run unemployment elasticity of pay approximately equal to -2%. The estimated unemployment effect is robust to the definition of regional labour market areas. Although the regional wage responses to unemployment seem to increase over time, we conclude that the equilibrating mechanism working through the regional wage formation process is weak.

The low degree of regional wage responses to open unemployment is an argument in favour of an active regional policy. Interestingly, reduced regional payroll taxes will mainly reduce regional wage costs and consequently stimulate regional employment. Furthermore, we find a numerically important wage dampening effect of labour market programs. However, the estimated effects of labour market programs are not statistically significant and we can not reject the null hypothesis that only total unemployment matters.

Recent studies of the Swedish wage formation process find different wage effects of manpower training programs and relief work. An interesting issue for further research would be to allow for different effects of the different types of labour market programs used in Norway.

Finally, we find that regional manufacturing profitability is important in shaping regional manufacturing wages while there is no evidence of an insider (or membership) hysteresis effect.

A Data and sources

W_j = wage costs per worker including payroll taxes, municipality j, manufacturing. Source: Statistics Norway (SN).

P_j = value added per worker at factor prices, municipality j, manufacturing. Source: SN.

N_j = manufacturing employment, municipality j. Source: SN.

TP_j = payroll tax rate municipality j. Source: Norwegian Tax Inspectorate.

U_j = rate of open unemployment, municipality j. Defined as number of registered unemployed divided by labour force. Source: SN.

U_c = rate of open unemployment county c. Same definition as for U_j. Source: SN.

U_l = rate of open unemployment LLOA l. Same definition as for U_j. Source: SN.

R_c = participation rate county c. Number of program participants relative to labour force. Source: Directorate of Labour.

TU_c = total unemployment rate county c. Defined as $U_c + R_c$.

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NOTES

[1] There are (in 1999) 435 municipalities, 19 counties and 106 LLOAs.

[2] The number of municipalities has changed during our sample period. Due to admistritative changes some municipalities has been merged in order to have cross-sections covering the same geographical area during the whole period. More important, municipalities with less than 20 manufacturing workers have been excluded from the data set.

[3] See appendix A for definitions and sources. Dyrstad and Johansen (2000) provide descriptive statistics for most variables for the period 1970–88 while details on labour market programs can be found in Raaum and Wulfsberg (1998).

[4] All results are computed using DPD98 for GAUSS, see Arellano and Bond (1998). We report the robust one-step estimates since simulations suggest that the asymptotically more efficient two-step estimator gives downward biased estimates of the standard errors in finite samples.

[5] It should be noted that the estimated standard errors of the unemployment effects in equation II and III may be biased downwards due to common group effects, cf. Moulton (1990).

[6] The standard errors of the long-run coefficients are calculated using the procedure suggested by Bårdsen (1989).

[7] Brunstad and Dyrstad (1997) find evidence of compensating differentials on Norwegian data while the results in Johansen *et al.* (1999) are not supportive to the hypothesis. Partridge and Rickman (1997) provide recent evidence for the US.

[8] Results referred in the text but not reported are available from the author upon request.

[9] The Sargan diff1 test is computed by re-estimating the model in question excluding Δu_{it} and u_{it-1} from the set of instruments. To compute the Sargan diff2 test the models are re-estimated in first differences.

[10] The sample maximum for municipality unemployment is 10.3%, for county 4.9%, and 6.7% for LLOA.

[11] It should be noted that the labour market variables used by Forslund are all aggregate variables.

[12] The wage curves estimated by Pannenberg and Schwarze include the log of the "job searcher rate" – which broadly corresponds to the log of total unemployment – and the "accomodation rate".

[13] The correlation coefficient between tp_{jt} and $(u - tu)_{ct}$ is 0.47.

Table 1: Regional Wage Equations – GMM-System Estimates
Dependent variable is Δw_{jt} Sample period: 1973–92 (322 municipalities)

	I	II	III
w^*_{jt-1}	-0.502 (0.034)	-0.513 (0.034)	-0.511 (0.034)
p^*_{jt-1}	0.113 (0.023)	0.119 (0.025)	0.116 (0.025)
tp_{jt-1}	0.482 (0.138)	0.439 (0.141)	0.433 (0.142)
u_{jt-1}	-0.011 (0.002)	-	-
u_{ct-1}	-	-0.014 (0.005)	-
u_{lt-1}	-	-	-0.010 (0.004)
Δp^*_{jt}	0.147 (0.033)	0.155 (0.036)	0.150 (0.036)
Δtp_{jt}	0.900 (0.270)	0.779 (0.261)	0.843 (0.267)
Δu_{jt}	-0.009 (0.003)	-	-
Δu_{ct}	-	0.000 (0.007)	-
Δu_{lt}	-	-	-0.005 (0.005)
n^*_{jt-1}	0.040 (0.007)	0.044 (0.008)	0.045 (0.007)
Long run estimates			
p_j	0.225 (0.041)	0.232 (0.044)	0.227 (0.044)
tp_j	0.960 (0.268)	0.854 (0.268)	0.847 (0.271)
u_j	-0.022 (0.005)	-	-
u_c	-	-0.027 (0.010)	-
u_l	-	-	-0.020 (0.007)
Diagnostics			
$\widehat{\sigma}$	0.081	0.083	0.084
Joint significance[8]	269.8, p=0.00	276.82, p=0.00	293.79, p=0.00
Sargan [232]	243.8, p=0.28	247.0, p=0.24	247.5, p=0.23
Sargan diff1[2]	1.21, p=0.55	3.44, p=0.18	3.20, p=0.20
Sargan diff2[84]	67.4, p=0.90	75.2, p=0.74	70.4, p=0.85
Hausman[8]	15.37, p=0.052	14.59, p=0.068	13.38, p=0.100
Hausman[2]	5.77, p=0.056	3.83, p=0.147	2.77, p=0.250
AR1	-10.6, p=0.00	-10.50, p=0.00	-10.61, p=0.00
AR2	-0.36, p=0.62	-0.41, p=0.68	-0.38, p=0.71

Notes:

i One-step robust standard errors reported in parentheses.

ii Time dummies are included in all equations.

iii Starred variables are instrumented. Additional instruments are w_{jt-2}, w_{jt-3}, p_{jt-2}, p_{jt-3}, u_{it-2}, u_{it-3}, n_{jt-2}, n_{jt-3} in the differenced equation and $\Delta w_{jt-1}, \Delta p_{jt-1}$, $\Delta u_{it-1}, \Delta n_{jt-1}$ in the levels equations.

iv $\widehat{\sigma}$ is estimated equation standard error (levels).

v Joint significance is a Wald χ^2 test testing the joint significance of all variables included (not time dummies).

vi Sargan is the Sargan (1958) (joint) test of instrumental validity. Sargan diff1 tests the validity of treating regional unemployment exogenous. Sargan diff2 tests the validity of the orthogonality restriction for the levels equations.

vii Hausman[8] is a Hausman type test, testing the null that all parameters are equal in the levels and first differenced equations. Hausman[2] tests the null that the unemployment effects are equal.

viii ARk is the Arellano and Bond (1991) test for the presence of k-th order serial correlation in the first-differenced residuals. Asymptotic normal under the null.

Table 2: Testing Nonlinear Unemployment Effects – GMM-System Estimates

Dependent variable is Δw_{jt} Sample period: 1973–92 (322 municipalities)

	I	II	III
w^*_{jt-1}	-0.505 (0.034)	-0.514 (0.034)	-0.512 (0.034)
p^*_{jt-1}	0.116 (0.024)	0.120 (0.025)	0.116 (0.025)
tp_{jt-1}	0.487 (0.137)	0.416 (0.143)	0.426 (0.144)
u_{jt-1}	-0.014 (0.003)	-	-
$u^3_{jt-1}/100$	0.019 (0.007)	-	-
u_{ct-1}	-	-0.019 (0.006)	-
$u^3_{ct-1}/100$	-	0.271 (0.109)	-
u_{lt-1}	-	-	-0.012 (0.005)
$u^3_{lt-1}/100$	-	-	0.061 (0.068)
Δp^*_{jt}	0.151 (0.034)	0.156 (0.036)	0.151 (0.036)
Δtp_{jt}	0.883 (0.270)	0.792 (0.262)	0.846 (0.267)
Δu_{jt}	-0.007 (0.003)	-	-
Δu_{ct}	-	0.001 (0.007)	-
Δu_{lt}	-	-	-0.005 (0.005)
n^*_{jt-1}	0.038 (0.007)	0.045 (0.008)	0.045 (0.007)
Long run estimates			
p_j	0.230 (0.041)	0.233 (0.044)	0.227 (0.044)
tp_j	0.964 (0.266)	0.810 (0.269)	0.833 (0.274)
u_j	-0.029 (0.006)	-	-
$u^3_j/100$	0.038 (0.015)	-	-
u_c	-	-0.037 (0.012)	-
$u^3_c/100$	-	0.528 (0.218)	-
u_c	-	-	-0.023 (0.009)
$u^3_c/100$	-	-	0.120 (0.133)
$\widehat{\sigma}$	0.081	0.084	0.084
Joint significance[9]	290.8, p=0.00	288.8, p=0.00	278.2, p=0.00
Sargan[232]	243.3, p=0.29	248.7, p=0.22	247.0, p=0.24
AR1	-10.6, p=0.00	-10.50, p=0.00	-10.61, p=0.00
AR2	-0.35, p=0.73	-0.40, p=0.69	-0.37, p=0.71

Notes: See notes to Table 1

Table 3: Effects of Labour Market Programmes – GMM-System Estimates
Dependent variable is Δw_{jt} Sample period: 1981–92 (322 municipalities)

	I	II	III
w^*_{jt-1}	-0.485 (0.041)	-0.488 (0.041)	-0.484 (0.041)
p^*_{jt-1}	0.080 (0.030)	0.082 (0.030)	0.081 (0.030)
tp_{jt-1}	0.480 (0.190)	0.333 (0.216)	0.456 (0.180)
u_{ct-1}	-0.017 (0.007)	-	-
tu_{ct-1}	-	-0.020 (0.008)	-0.020 (0.007)
$(u-tu)_{ct-1}$	-	0.037 (0.034)	-
Δp^*_{jt}	0.093 (0.041)	0.095 (0.042)	0.094 (0.041)
Δtp_{jt}	0.536 (0.318)	0.428 (0.336)	0.522 (0.321)
Δu_{ct}	0.008 (0.012)	-	-
Δtu_{ct}	-	-0.003 (0.015)	-
$\Delta (u-tu)_{ct}$	-	0.057 (0.039)	-
n^*_{jt-1}	0.043 (0.011)	0.046 (0.011)	0.041 (0.011)
LR estimates			
p_j	0.167 (0.057)	0.168 (0.056)	0.168 (0.057)
tp_j	0.989 (0.386)	0.682 (0.434)	0.944 (0.366)
u_c	-0.036 (0.015)	-	-
tu_c	-	-0.040 (0.017)	-0.041 (0.015)
$u_c - tu_c$	-	0.076 (0.070)	-
Diagnostics			
$\widehat{\sigma}$	0.083	0.084	0.083
Joint sign	157[8], p=0.00	173[10], p=0.00	144.9[7], p=0.00
Restr	-	-	2.38[3], p=0.50
Sargan	136[124], p=0.28	130[124], p=0.32	136[127], p=0.27
AR1	-9.47, p=0.00	-9.47, p=0.00	-9.47, p=0.00
AR2	-0.25, p=0.80	-0.26, p=0.80	-0.28, p=0.78

Notes: Restr is a Wald χ^2 test, testing the validity of the three restriction imposed to Equation III. See also notes to Table 1.

Table 4: Effects of Labour Market Programmes – GMM-System Estimates
Interaction Effects
Dependent variable is Δw_{jt} Sample period: 1981–92 (322 municipalities)

	I	II
w^*_{jt-1}	-0.487 (0.041)	-0.486 (0.041)
p^*_{jt-1}	0.081 (0.029)	0.081 (0.029)
tp_{jt-1}	0.304 (0.214)	0.291 (0.196)
tu_{ct-1}	-0.016 (0.008)	-0.016 (0.007)
$(u - tu)_{ct-1}$	-0.009 (0.039)	-
$U_{ct-1} \times (u - tu)_{ct-1}$	0.013 (0.008)	0.012 (0.007)
Δp^*_{jt}	0.093 (0.041)	0.093 (0.040)
Δtp_{jt}	0.398 (0.342)	0.388 (0.335)
Δtu_{ct}	0.004 (0.011)	-
$\Delta (u - tu)_{ct}$	-0.009 (0.042)	-
$\Delta [U_{ct} \times (u - tu)_{ct}]$	0.020 (0.009)	0.020 (0.009)
n^*_{jt-1}	0.046 (0.011)	0.046 (0.011)
LR estimates		
p_j	0.166 (0.056)	0.167 (0.055)
tp_j	0.626 (0.431)	0.598 (0.398)
tu_c	-0.032 (0.017)	-0.033 (0.015)
$u_c - tu_c$	-0.018 (0.080)	-
$U_c \times (u_c - tu_c)$	0.027 (0.017)	0.025 (0.014)
Diagnostics		
$\widehat{\sigma}$	0.084	0.084
Joint sign	175[12], p=0.00	160[9], p=0.00
Restr	-	0.119[3] , p=0.99
Sargan	125[124], p=0.46	126[127], p=0.52
AR1	-9.50, p=0.00	-9.49, p=0.00
AR2	-0.21, p=0.84	-0.21, p=0.83

Notes: Restr is a Wald χ^2 test, testing the validity of the three restriction imposed to Equation III. See also notes to Table 1.

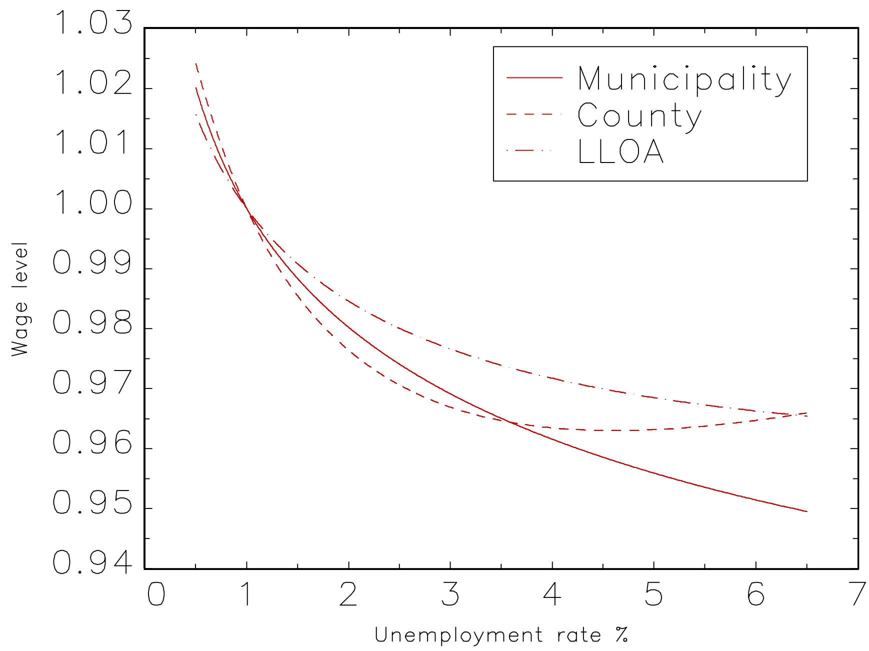

Figure 1: Nonlinear Regional Wage Curves