ISSN 1503-299X

WORKING PAPER SERIES

No. 18/2002

A WAGE CURVE FOR THE INTERWAR LABOUR MARKET: EVIDENCE FROM A PANEL OF NORWEGIAN MANUFACTURING INDUSTRIES

Gunnar Bårdsen
Jurgen Doornik
Jan Tore Klovland

Department of Economics

Norwegian University of Science and Technology
N-7491 Trondheim, Norway
www.svt.ntnu.no/iso/wp/wp.htm

A wage curve for the interwar labour market: Evidence from a panel of Norwegian manufacturing industries*

Gunnar Bårdsen
Norwegian University of Science and Technology, Dragvoll, Trondheim

Jurgen Doornik
Nuffield College, Oxford

Jan Tore Klovland
Norwegian School of Economics and Business Administration, Bergen

Abstract

We present an econometric analysis of wage behaviour in Norway during the interwar years. Applying GMM estimation methods to a newly constructed panel of manufacturing industry data, we find that the interwar years do not seem to be such an anomalous time period as has been suggested with respect to wage behaviour. We estimate a long-run wage curve that has all the modern features of being homogeneous in prices, proportional to productivity, and having an unemployment elasticity of -0.1. We also present some new Monte Carlo evidence on the properties of the estimators used.

Discussion paper 5/00
May 2000, revised April 2001

JEL Classification: E24,N24

*We would like to thank Ola H. Grytten, Kåre Johansen, and Fredrik Wulfsberg for helpful comments.

1 Introduction

Empirical wage curves, which relate nominal wages to consumer and producer prices, productivity and unemployment, have been successfully identified for many countries on postwar data, as documented in Blanchflower and Oswald (1994). These equations usually reflect homogeneity with respect to the price level, and exhibit reasonable stability over time.

In contrast, wage behaviour in the interwar period has often been seen as somewhat anomalous, being difficult to fit entirely into the empirical framework used to explain postwar wage formation. Empirical wage equations from the interwar period appear to be much more fragile, seldom fulfilling the whole list of desirable theoretical properties referred to above. For the United Kingdom Hatton (1988), Dimsdale et al. (1989) and Broadberry (1986) estimated several wage equations for the interwar years, including a wage-bargain model and a Phillips-curve type of model, using quarterly time series data, but no empirically well-specified model, which is fully consistent with theory, was obtained. The results from other European countries reported by Newell and Symons (1988) are somewhat more in line with standard wage equations than is the case for Britain, but even here there is only a weak feedback from unemployment to the real wage.

One explanation for these empirical findings may of course be that wage formation in interwar labour markets was indeed different from the postwar period, either because of the large fluctuations in prices and unemployment in that period, or as a result of long-term structural changes. Data from the United States indicate that there was a change in the cyclical behaviour of real wages between the interwar period and the postwar years.[1] This fact does not necessarily imply that there were changes in the structural parameters of labour demand and supply equations, however. Such changes might for example stem from differences in the relative magnitudes of labour demand and supply shocks in the two time periods.[2]

In this paper we take another approach to the seemingly instability of interwar labour market equations - data requirements and estimation methods. Our hypothesis is that a 'standard' postwar labour market model may be able to explain the interwar period as well, once a more powerful data set is available and the proper estimation methods are applied. Most previous studies have been poorly equipped to identify a stable and well identified relationship, being confined to use the relatively small samples of time series data available for the interwar years. Even quarterly data, typically over a period of at most 15 years, may provide a relatively poor basis for identifying stable relationships, given the varying data quality of the key variables during the period.[3]

The novel feature of our approach is to estimate standard wage equations using a *panel data set* recently constructed by Klovland (1999) for Norwegian manufacturing.

[1] See Bernanke and Powell (1986) and Hanes (1996) for evidence on the changing cyclicality of real wages.

[2] On the other hand, Hanes (1996) rejected the hypothesis of relative changes in demand and supply shocks in favour of an explanation in terms of a shift towards more finished goods in the consumption bundle of consumers, making the real consumption wage more procyclical over time.

[3] The fact that Bernanke (1986) obtained quite well-behaved real earnings equations using US *monthly* manufacturing data of relatively high quality from the interwar period may indicate that better data may be of some importance.

Panel data estimation is likely to provide more information than time series estimation over a relatively short sample period, since we can draw inference from the cross-section variation in the data in addition to the time series volatility of the early 1930s. The data base contains annual values of key output and labour market variables for 55 manufacturing industries over the period 1927 to 1939: nominal average hourly earnings, producer price indices, labour productivity (real value added per hour) and, at a somewhat less disaggregated level, unemployment rates.

Section 2 describes the theoretical model, which we think is sufficiently general to encompass wage behaviour in both prewar and postwar years.[4] We report the empirical modelling of the wage equation for the years 1927 - 1939 in Sections 3 to 6, focussing on the economic interpretation of the results as well as methodological issues related to estimation methods. An example of the latter is contained in Section 4, where we present some new Monte Carlo evidence on the properties of the estimators used.

2 The wage equation

Theories of wage formation with of imperfect competition in goods and labour markets[5] state that the bargained nominal wage level mainly depends on:

- firm-side variables, e.g. productivity, producer prices and the payroll-tax rate
- factors affecting workers' take home pay, e.g. retail prices and the income tax-rate
- labour-market pressure
- institutional features, e.g. the existence of centralized wage-bargaining institutions and the degree of mismatch
- earlier outcomes of the relevant variables.

The general dynamic specification, adapted from Kolsrud and Nymoen (1998), we use is

$$\begin{aligned}(1-\alpha_1 L)\,w_{\mathsf{it}} &= (\beta_0+\beta_1 L)\,p_{\mathsf{it}}+(\gamma_0+\gamma_1 L)\,q_{\mathsf{it}} \\ &\quad +(\delta_0+\delta_1 L)\,u_{\mathsf{it}}+(\zeta_0+\zeta_1 L)\,pc_{\mathsf{t}}+\eta_{\mathsf{i}}+\varepsilon_{\mathsf{it}}. \qquad (1)\end{aligned}$$

The variables are nominal hourly earnings W, producer prices P, labour productivity Q, the unemployment rate U, and retail prices PC.[6] Small letters denote natural logarithms of the corresponding variables denoted in capitals, so $x_{\mathsf{it}} \equiv \ln X_{\mathsf{it}}$. The letter L denotes the lag operator, defined by $Lx_{\mathsf{it}} = x_{\mathsf{i(t-1)}}$. Hence w_{it} denotes the logarithm of the nominal wage in the $i'th$ industry in period t. The variables p_{it}, q_{it} and

[4] In Norway there was no publicly administered unemployment insurance scheme before 1938, but some trade union members received unemployment benefits from their unions. This amount was fairly small, about one third of the average wage, and did not vary much during the period (Grytten (2000)). Hence, the level of unemployment benefits is not a strong candidate for inclusion in wage equations in the interwar years.

[5] See e.g. Carlin and Soskice (1990) and Lindbeck (1993).

[6] We disregard tax rates, which were rather low during the interwar period.

u_{it} are industry-specific variables, while economy-wide effects that are not transmitted through the unemployment rate, say, are captured by the retail price index pc_{t}.

Nominal wage growth responds positively to increases in producer and retail prices, labour productivity, and negatively to increased unemployment. A natural property of a wage equation is that in the long run the nominal wage level is homogenous of degree one with respect to the two price variables (industry-specific output prices and general retail prices), but that there is some degree of wage level stickiness in the short run. We would also expect that productivity growth increases real wages in the same proportion in the long run. The Equilibrium Correction Model (EqCM) provides an intuitively appealing way to implement such considerations empirically.[7] The EqCM reparameterization of (1) is

$$\Delta w_{\mathsf{it}} = \beta_0 \Delta p_{\mathsf{it}} + \gamma_0 \Delta q_{\mathsf{it}} + \delta_0 \Delta u_{\mathsf{it}} + \zeta_0 \Delta pc_{\mathsf{t}} - \alpha_1 \left(w - w^*\right)_{\mathsf{i(t-1)}} + \eta_{\mathsf{i}} + \varepsilon_{\mathsf{it}}, \qquad (2)$$

where $\Delta x_{\mathsf{it}} = x_{\mathsf{it}} - x_{\mathsf{i(t-1)}}$ and w^*_{it} is the steady-state wage level

$$\begin{aligned} w^*_{\mathsf{it}} &= \left(\frac{\beta_0 + \beta_1}{1-\alpha_1}\right) p_{\mathsf{it}} + \left(\frac{\gamma_0 + \gamma_1}{1-\alpha_1}\right) q_{\mathsf{it}} + \left(\frac{\delta_0 + \delta_1}{1-\alpha_1}\right) u_{\mathsf{t}} + \left(\frac{\zeta_0 + \zeta_1}{1-\alpha_1}\right) pc_{\mathsf{t}} \\ &= \beta^* p_{\mathsf{it}} + \gamma^* q_{\mathsf{it}} + \delta^* u_{\mathsf{t}} + \zeta^* pc_{\mathsf{t}}. \end{aligned} \qquad (3)$$

Price level homogeneity requires that $\beta^* + \zeta^* = 1$. We also test the long-run proportionality assumption of labour productivity, $\gamma^* = 1$. Institutional and structural features are reflected in the coefficients of (3). Changes in the impact of institutions on wage setting can therefore be tested by looking at the empirical stability of (3) over the sample period. It is quite likely that wages interact simultaneously with all the explanatory variables—with the likely exception of the retail price index. In the present setting, however, we would like to focus on the behaviour of wages. We do, of course, take the possible simultaneity into account when estimating the model by using instrumental variables.

3 Testing specifications

The wage equations are estimated using both the GMM estimator of Arellano and Bond (1991) and the system GMM estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998). Both estimators allow control for the presence of unobserved industry-specific effects and for the possible endogeneity of the explanatory variables. Both GMM estimators use equations in first-differences to eliminate the industry-specific fixed effects. Endogenous variables in levels lagged two or more periods will be valid instruments, provided there is no autocorrelation in the time-varying component of the error terms. This is tested by examining tests for serial correlation in the first-differenced residuals, following Arellano and Bond (1991). For the system GMM estimator, the differenced equations—using level instruments—are combined with equations in levels—using differences as instruments. Blundell and Bond (1998) show that first differences of the series may be uncorrelated with the industry-specific

[7] Johansen (1996) and Wulfsberg (1997) estimate postwar panel wage equations for Norway using this approach.

Table 1: The different specifications considered

	Equations	**GMM instruments**	**Anderson & Hsiao instruments**
Diff	Differenced	$w_{\mathsf{it}-2}, w_{\mathsf{it}-3}$	$pc_{\mathsf{t}}, pc_{\mathsf{t}-1}, p_{\mathsf{it}-2}, p_{\mathsf{it}-3}, q_{\mathsf{it}-2}, q_{\mathsf{it}-3}, u_{\mathsf{it}-2}, u_{\mathsf{it}-3}$
Diff-end	Differenced	$w_{\mathsf{it}-2}, w_{\mathsf{it}-3}, p_{\mathsf{it}-2}, p_{\mathsf{it}-3}, q_{\mathsf{it}-2}, q_{\mathsf{it}-3}, u_{\mathsf{it}-2}, u_{\mathsf{it}-3}$	$pc_{\mathsf{t}}, pc_{\mathsf{t}-1}$
Sys	Differenced & levels	$w_{\mathsf{it}-2}, w_{\mathsf{it}-3}, \Delta w_{\mathsf{it}-1}$	$pc_{\mathsf{t}}, pc_{\mathsf{t}-1}, p_{\mathsf{it}-2}, p_{\mathsf{it}-3}, q_{\mathsf{it}-2}, q_{\mathsf{it}-3}, u_{\mathsf{it}-2}, u_{\mathsf{it}-3}$
Sys-end	Differenced & levels	$w_{\mathsf{it}-2}, w_{\mathsf{it}-3}, p_{\mathsf{it}-2}, p_{\mathsf{it}-3}, q_{\mathsf{it}-2}, q_{\mathsf{it}-3}, u_{\mathsf{it}-2}, u_{\mathsf{it}-3}$ $\Delta w_{\mathsf{it}-1}, \Delta p_{\mathsf{it}-1}, \Delta q_{\mathsf{it}-1}, \Delta u_{\mathsf{it}-1}$	$pc_{\mathsf{t}}, pc_{\mathsf{t}-1}$

Note: The Anderson & Hsiao instruments enter as differences or levels according to the transformation in use.

effects in the case of stationary series. We therefore use lagged differences for the variables as instruments for the levels equations. In the specifications labelled *Diff* and *Sys* the following variables are considered exogenous: productivity q_{i}, producer prices p_{i}, and unemployment u_{i}. In the specifications labelled *Diff-end* and *Sys-end* the same variables are treated as endogenous. In all specifications the retail price index is treated as exogenous. The validity of the instruments are in each case tested by means of the Sargan test of over-identifying restrictions. The exact specifications considered for the different wage equations are given in Table 1.

To avoid overfitting, and thus cancel the effects of instrumenting, we keep the number of instruments fixed as the number of time periods increases.

The results are generated using Ox version 2.20 (see Doornik, 1999)

and the DPD package (Doornik et al., 1999). The estimated wage equations using the different specifications are reported in Table 2.

We report results using the one-step estimators, with standard errors and test statistics that are asymptotically robust to general heteroscedasticity, since the standard errors of the two-step estimators are considered to produce standard errors that are downward biased.

All specifications seem to capture the relevant dynamics, since no second order residual correlation is evident. A general impression is that the system estimators produce more reasonable estimates than the first difference estimators. The differences are in particular striking for the autoregressive term, with the estimated parameter being notably higher using the system estimators. This is consistent with the analysis of Blundell and Bond (1998). They show that in autoregressive models with persistent series, the first-differenced estimator can be subject to serious finite sample biases as a result of weak instruments, and that these biases can be greatly reduced by the inclusion of the levels equations in the system estimator. This result is in particular relevant in the present setting, where the degree of nominal wage rigidity is measured by the autoregressive parameter. A first impression therefore favours the system estimators.

However, in the Monte Carlo experiments reported by Blundell and Bond (1998) only a purely autoregressive process is considered, whereas a more realistic situation

Table 2: Wage equations, GMM estimates

Dep. var: w_{it}	Diff	**Diff-end**	Sys	**Sys-end**
w_{it-1}	0.302 (0.068)	0.266 (0.057)	0.711 (0.042)	0.844 (0.039)
p_{it}	0.051 (0.030)	−0.018 (0.058)	0.051 (0.052)	−0.111 (0.073)
p_{it-1}	0.026 (0.031)	0.084 (0.053)	0.011 (0.043)	0.188 (0.080)
q_{it}	0.147 (0.023)	0.199 (0.04)	0.137 (0.040)	0.120 (0.054)
q_{it-1}	0.011 (0.023)	0.034 (0.027)	−0.055 (0.040)	−0.008 (0.037)
u_{it}	0.003 (0.006)	0.008 (0.008)	0.014 (0.012)	0.013 (0.012)
u_{it-1}	−0.023 (0.006)	−0.043 (0.007)	−0.034 (0.011)	−0.041 (0.013)
pc_t	0.201 (0.101)	0.159 (0.118)	0.852 (0.120)	0.946 (0.088)
pc_{t-1}	0.252 (0.125)	0.305 (0.126)	−0.655 (0.099)	−0.841 (0.101)
Diagnostics				
Sargan: $\chi^2(\cdot)$	38.01 (20)	53.49 (77)	50.97 (31)	52.62 (121)
$AR(1)$	−4.78	−5.47	−4.21	−4.35
$AR(2)$	−1.70	−1.73	0.53	0.25
Steady state analysis: $w^*_{it} = \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_t + \zeta^* pc_t$				
β^*	0.110 (0.051)	0.090 (0.082)	0.214 (0.205)	0.492 (0.296)
γ^*	0.227 (0.042)	0.312 (0.060)	0.281 (0.174)	0.721 (0.279)
δ^*	−0.033 (0.014)	−0.048 (0.015)	−0.070 (0.051)	−0.181 (0.080)
ζ^*	0.650 (0.064)	0.632 (0.085)	0.682 (0.224)	0.673 (0.289)
Testing steady-state restrictions				
$\beta^* + \zeta^* = 1$	32.67 (1)	44.66 (1)	0.44 (1)	0.51 (1)
$\beta^* + \zeta^* = 1, \gamma^* = 1$	362.21 (2)	164.19 (2)	17.94 (2)	2.64 (2)
$\beta^* + \zeta^* = 1, \gamma^* = 1, \delta^* = -0.1$	517.31 (3)	323.45 (3)	33.47 (3)	3.94 (3)

would be cases like the present analysis with additional variables. To gain some further insight into the properties of the different estimators before we proceed, we therefore conducted a Monte Carlo experiment, using a simplified data generating process(DGP) more relevant for the analysis at hand.

4 A simulation experiment of the properties of the estimators

The homoscedastic DGP in Arellano and Bond (1991) is:

$$\begin{aligned} y_{it} &= \alpha y_{i,t-1} + \beta z_{i1} + \eta_i + v_{it}, \quad \eta_i \sim \mathsf{N}[0,1] \quad v_{it} \sim \mathsf{N}[0,1] \\ & \qquad i = 1,\dots,N,\ t = 1,\dots,T \\ z_{it} &= \rho z_{i,t-1} + e_{it}, \quad e_{it} \sim \mathsf{N}[0,\sigma_e^2]. \end{aligned}$$

This DGP is used in Doornik et al. (1999) to illustrate how the system GMM estimator (*Sys*) gives more precise estimates of the autoregressive parameter α than the differenced GMM estimator (*Diff*) when α is close to unity. It was also noted that *Diff* underestimates α, whereas *Sys* produces an overestimate. While Doornik et al. (1999) keep β fixed at unity, we now proceed to keep α fixed at 0.9, and vary β. We set $N = 100$, and $T = 7$ (5 after allowing for lags and differences).

The two estimators can be summarized as:

	transformation	regressors	instruments	estimation
Diff	Δ	$\Delta \mathsf{y}_{i,-1}, \Delta \mathsf{x}_i, 1$	$\mathrm{diag}(y_{i,t-3} y_{i,t-2}), \Delta \mathsf{x}_i, 1$	1-step
Sys	Δ	$\Delta \mathsf{y}_{i,-1}, \Delta \mathsf{x}_i$	$\mathrm{diag}(y_{i,t-3} y_{i,t-2}), \Delta \mathsf{x}_i$	1-step
	levels:	$\mathsf{y}_{i,-1}, \mathsf{x}_i, 1$	$\mathrm{diag}(\Delta y_{i,t-2}), \mathsf{x}_i, 1$	

When $T = 5$, for example, the instruments Z in *Diff* estimation are:

$$Z_i = \begin{pmatrix} y_{i0} & 0 & 0 & 0 & 0 & \Delta x_{i,2} & 1 \\ 0 & y_{i0} & y_{i1} & 0 & 0 & \Delta x_{i,3} & 1 \\ 0 & 0 & 0 & y_{i1} & y_{i2} & \Delta x_{i,4} & 1 \end{pmatrix}.$$

This assumes that initially the available observations are $t = 0,\dots,4$. One observation is lost owing to the lagged dependent variable, and one more by differencing. For *Sys* estimation the instruments for the differenced equations (Z^*) and level equations (Z^+) are:

$$Z_i^* = \begin{pmatrix} y_{i0} & 0 & 0 & 0 & 0 & \Delta x_{i,2} \\ 0 & y_{i0} & y_{i1} & 0 & 0 & \Delta x_{i,3} \\ 0 & 0 & 0 & y_{i1} & y_{i2} & \Delta x_{i,4} \end{pmatrix}, \quad Z_i^+ = \begin{pmatrix} \Delta y_{i1} & 0 & 0 & x_{i,2} & 1 \\ 0 & \Delta y_{i2} & 0 & x_{i,3} & 1 \\ 0 & 0 & \Delta y_{i3} & x_{i,4} & 1 \end{pmatrix}$$

Some results for $M = 1000$ Monte Carlo replications are presented in Figure 2. MCSD is the standard deviation of the estimated $\hat{\alpha}$. The results can be compared with Table 1 of Arellano and Bond (1991) (but we use instruments $t-2, t-3$ instead of all possible lags from $t-2$ onwards), and Table 2 of Blundell and Bond (1998) (but with larger T, and an additional regressor).

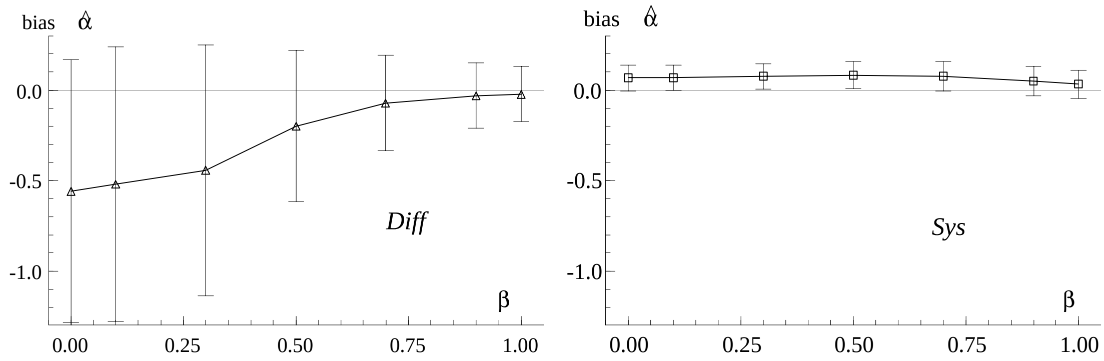

Figure 2: Mean bias of $\hat{\alpha}$, $M = 1000, \alpha = 0.9, \rho = 0.8, \sigma^2_e = 0.9$; bars are twice the MCSD; $\beta = 0, 0.1, 0.3, 0.5, 0.7, 0.9, 1$.

The results are dramatic. Despite the fact that the generated x is kept constant in replications, the bias of the *Diff* estimator is enormous for small values of β; for example when $\beta = 0.3$, the mean estimated $\hat{\alpha}$ is close to 0.5. *Sys* again overestimates α, but is much better behaved. These results shed some light on Table 2: the large discrepancy between the *Diff* and *Sys* results reported there corresponds to a low value of β in Figure 2.

The bias in $\hat{\beta}$ is never so dramatic, ranging from about 0.01 to -0.04 for *Diff*, and from 0.01 to -0.08 for *Sys*.

5 The steady state

On the basis of the experiment above, the Sys and $Sys - end$ specifications are clearly to be favoured. A further issue is the exogeneity assumptions. The exogeneity of the explanatory variables in $Diff$ and Sys in Table 2 are rejected by the Sargan tests, with p-values of 0.008 and 0.0134, respectively, so this leaves $Sys - end$ as the most reliable candidate.

A basic requirement for a well specified dynamic model ought to be a sensible steady-state solution. We therefore next focus on the long-run solution of the estimated equations, using the approach of Bårdsen (1989). The hypothesis of long-run price homogeneity is rejected in both differenced equations, while the systems specifications cannot reject the hypothesis. But only the $Sys - end$ specification accepts the joint hypothesis of price homogeneity and proportionality of productivity. Again we therefore end up with $Sys - end$ as the most reasonable specification. We will therefore use the results from this estimator in the rest of the paper.

Given the turbulent period we are investigating, a relevant question is whether the wage curve we claim to have found is indeed a genuine relationship, or just effects that happened to dominate at the end of our sample in 1939. To answer this question we estimated the steady-state solution recursively, as reported in Figure 3[8]. All parameters remain stable across the 30s, with the exception of the effect of retail prices, which is insignificant until the latter part of the sample. Whether this effect is due to lack of

[8] See also Johansen (1999).

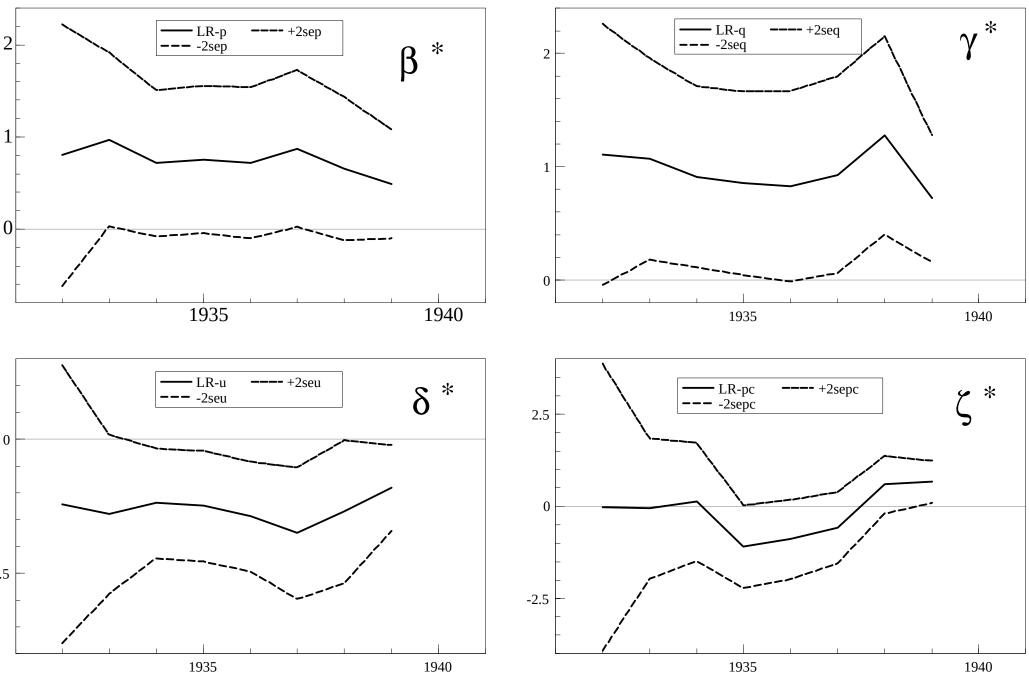

Figure 3: Recursive estimates of the steady-state parameters

cross-section variation is an issue that remains to be investigated. We do note, however, that the effect of retail prices is the parameter most invariant across specifications.

The effect of unemployment on wages is an important issue when analyzing the interwar labour market. This is in particular the case since the publication of Blanchflower and Oswald (1994), who claim to have found an empirical law stating that the unemployment elasticity of wages is -0.1, so a doubling of unemployment reduces wages by 10%. We cannot reject that hypothesis on the basis of our data. The test of the joint hypothesis of a long-run wage curve being homogeneous in prices, proportional to productivity, and having an unemployment elasticity of -0.1, produces a statistic with a p-value of 0.27. On the basis of the evidence so far, we therefore test whether the steady-state solution

$$w_{\mathrm{it}} = 0.35p_{\mathrm{it}} + q_{\mathrm{it}} - 0.1u_{\mathrm{t}} + 0.65pc_{\mathrm{t}} \tag{4}$$
$$\chi^2(4) = 4.14551[0.3867],$$

can be rejected. As the associated p-value in brackets suggests, this empirical representation of (3) cannot be rejected. It is therefore imposed when we next turn to estimating the dynamic specification in the equilibrium correction form given by (2).

6 The dynamic model

Having established the existence of a perfectly conventional long-run wage curve for Norway during the depression, we now want to investigate whether the short-run adjustment of wages during the interwar period differed from what is found in empirical

Table 3: Wage equations, GMM system estimates

Dep. var: Δw_{it}	(1)	(2)
Δp_{it}	-0.107 (0.068)	–
Δq_{it}	0.127 (0.069)	0.194 (0.054)
Δu_{it}	0.007 (0.008)	–
Δpc_t	0.817 (0.008)	0.613 (0.069)
$(w - w^*)_{i(t-1)}$	-0.175 (0.030)	-0.263 (0.041)
Diagnostics		
Sargan: χ^2 ()	53.39 (110)	54.09 (50)
$AR(1)$	-4.15	-3.87
$AR(2)$	-0.04	-0.17

studies of the postwar period. We could find no such evidence. Our preferred equation is a quite standard dynamic wage equation, with properties matching those found in comparable studies of the Norwegian economy during the postwar era. The relevant evidence is reported in Table 3. Column (1) contains the general model reparameterized in equilibrium correction form, with the long-run solution (4) imposed. The short-run effects of producer prices and unemployment are insignificant and can be dropped—the joint test statistic has a p-value of 0.31. This is of course in accordance with the corresponding results in Table 1. The final model is reported in column (2).[9] There is substantial nominal rigidity, as measured by the EqCM coefficient with a value of -0.26. Consequently, a drop in inflation is not likely to be reflected in a similar drop in wage growth, as documented by the coefficient of 0.6 on inflation. These magnitudes are similar to the evidence from time-series studies using recent Norwegian manufacturing data by Nymoen (1989) and Johansen (1995), as well as the panel studies of Johansen (1996) and Wulfsberg (1997).

It might be argued that it is reasonable that such results dominate in the latter half of the sample, as Norway recovered from the great depression, but that it does not necesserily reflect actual behaviour during the depressed years in the early 1930s. To investigate this possibility we therefore estimated our preferred equation in column (2) recursively. The estimated coefficients, together with their approximate confidence bands, are shown in Figure 4, starting with 1932. The coefficients display considerable stability over time, although there is some downward drift in the coefficient on the retail price inflation until 1935. Otherwise there is little evidence of changing behaviour during the sample period.

7 Conclusions

Our empirical analysis has shown that there is no particular puzzle regarding interwar labour market behaviour in the case of Norwegian manufacturing industries. The preferred steady-state wage equation features the standard properties of homogeneity

[9] The change in coefficients partly reflects changes in the list of instruments.

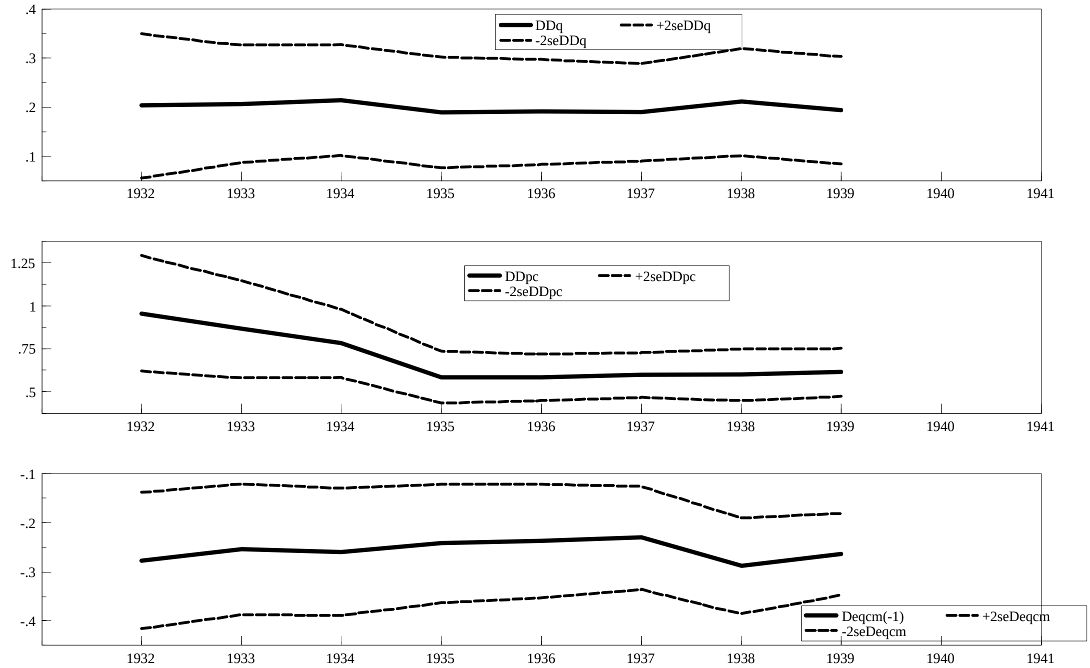

Figure 4: Recursive estimates of the model parameters.

with respect to prices and productivity, and there is an unemployment elasticity of -0.1. We also find much inertia in the dynamics of nominal wages. These results contrast with much of the empirical findings from other countries; such studies often report difficulties with replicating the standard postwar wage models on interwar data. We believe this result mainly stemsfrom the fact that we are able to use a panel data set of 55 manufacturing industries in our econometric analysis, rather than having to rely on a relatively short time series sample.

References

Arellano, M. and S. R. Bond (1991). Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations. *Review of Economic Studies*, *58*, 277–297.

Arellano, M. and O. Bover (1995). Another Look at the Instrumental-Variable Estimation of Error-Components Models. *Journal of Econometrics*, *68*, 29–52.

Bernanke, B. S. (1986). Employment, Hours, and Earnings in the Depression: An Analysis of Eight Manufacturing Industries. *American Economic Review*, *76*, 82–109.

Bernanke, B. S. and J. L. Powell (1986). The Cyclical Behavior of Industrial Labor Markets: A Comparison of the Prewar and Postwar Eras. In Gordon, R. J. (ed.), *The American Business Cycle*, 583–621. University of Chicago Press, Chicago and London.

Blanchflower, D. G. and A. J. Oswald (1994). *The Wage Curve*. The MIT Press, Cambridge, Massachusetts.

Blundell, R. and S. R. Bond (1998). Initial Conditions and Moment Restrictions in Dynamic Panel Data Models. *Journal of Econometrics*, *87*, 115–143.

Broadberry, S. N. (1986). Aggregate Supply in Interwar Britain. *The Economic Journal*, *96*, 467–481.

Bårdsen, G. (1989). Estimation of Long Run Coefficients in Error Correction Models. *Oxford Bulletin of Economics and Statistics*, *51*, 345–350.

Carlin, W. and D. Soskice (1990). *Macroeconomics and the Wage Bargain*. Oxford University Press, Oxford.

Dimsdale, N. H., S. J. Nickell and N. Horsewood (1989). Real Wages and Unemployment in Britain During the 1930s. *The Economic Journal*, *99*, 271–292.

Doornik, J. A. (1999). *Object-Oriented Matrix Programming Using Ox*. London: Timberlake Consultants Press and Oxford: www.nuff.ox.ac.uk/Users/Doornik, 3rd edn.

Doornik, J. A., M. Arellano and S. R. Bond (1999). Panel Data Estimation Using DPD for Ox. Mimeo.

Grytten, O. H. (1994). *En empirisk analyse av det norske arbeidsmarked 1918-1939: Arbeidsledigheten i Norge i internasjonalt perspektiv (An Empirical Analysis of the Norwegian Labour Market, 1918-1939: Norwegian Interwar Unemployment in International Perspective)*. Ph.D. thesis, Norwegian School of Economics and Business Administration, Bergen.

Grytten, O. H. (2000). A Comparison of the Living Standards of the Employed and Unemployed in Inter-war Norway. Unpublished manuscript, Norwegian School of Economics and Business Administration, Bergen.

Hanes, C. (1996). Changes in the Cyclical Behavior of Real Wage Rates, 1870-1990. *Journal of Economic History*, *56*, 837–861.

Hatton, T. J. (1988). A Quarterly Model of the Labour Market in Interwar Britain. *Oxford Bulletin of Economics and Statistics*, *50*, 1–25.

Johansen, K. (1995). Norwegian Wage Curves. *Oxford Bulletin of Economics and Statistics*, *57*, 229–247.

Johansen, K. (1996). Insider Forces, Asymmetries, and Outsider Ineffectiveness: Empirical Evidence for Norwegian Industries 1966–87. *Oxford Economic Papers*, *48*, 89–104.

Johansen, K. (1999). Insider Forces in Wage Determination: New Evidence for Norwegian Industries. *Applied Economics*, *31*, 137–147.

Klovland, J. T. (1999). Accounting for Productivity Growth in Norwegian Manufacturing Industries 1927 - 1959. Discussion Paper 21/99, Norwegian School of Economics and Business Administration, Bergen.

Kolsrud, D. and R. Nymoen (1998). Unemployment and the Open Economy Wage-Price Spiral. *Journal of Economic Studies*, *25*, 450–467.

Lindbeck, A. (1993). *Unemployment and Macroeconomics*. The MIT Press, Cambridge.

Newell, A. and J. Symons (1988). The Macroeconomics of the Interwar Years: International Comparisons. In Eichengreen, B. and T. Hatton (eds.), *Interwar Unemployment in International Perspective*, 61–96. Kluwer, Dorcrecht.

Nymoen, R. (1989). Modelling Wages in the Small Open Economy: An Error-Correction Model of Norwegian Manufacturing Wages. *Oxford Bulletin of Economics and statistics*, *51*, 239–258.

Wulfsberg, F. (1997). An Application of Wage Bargaining Models to Norwegian Panel Data. *Oxford Economic Papers*, *49*, 419–440.

A The Data

The wage, price and productivity series are annual data 1927 - 1939 for 55 manufacturing industry groups, see Klovland (1999) for further details as to coverage and sources. The unemployment data are taken from Grytten (1994). These are only available at a more aggregated level; data for 11 industry groups were distributed on the 55 subgroups. The retail price index is taken from *Historical Statistics 1948* (Statistics Norway, Oslo, 1949).

The data definitions are:

W = nominal hourly earnings Average hourly earnings of (male and female) production workers, calculated as total wage sum divided by hours worked by production workers.

P = producer prices Paasche price index of industry gross output, shifting base year every third year.

Q = labour productivity Real industry value added divided by total hours worked. Total hours also include an estimate of hours worked by non-production workers.

U = unemployment rate based on unemployed registered at public labour exchanges, classified by industry groups.

PC = retail price index