Wage formation and bargaining power during the Great Depression[*]

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Abstract

We present an econometric analysis of wage behaviour in Norway during the interwar years. The analysis is based on a panel of manufacturing industry data using GMM estimation methods. Our empirical analysis shows that wage formation in the interwar period can be understood with the help of modern bargaining theory and well established wage equations. We estimate a long-run wage curve that has all the standard features of being homogeneous in prices, proportional to productivity, and with a negative unemployment elasticity. We also present some new Monte Carlo evidence on the properties of the estimators used.

1 Introduction

Wage behaviour in the interwar period is often considered anomalous: it is difficult to fit within the empirical framework used to explain postwar wage formation. This can be seen from much of the evidence to date, which shows fragile empirical wage equations for the interwar period. For the United Kingdom, Hatton (1988), Dimsdale, Nickell, and Horsewood (1989) and Broadberry (1986) estimate several wage equations, including a wage-bargain model and a Phillips-curve type of model, using quarterly time series data, but no empirically well-specified model 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 one finds only a weak feedback from unemployment to the real wage.[1]

One reason for these results might be a decline in the bargaining power. During the depression years of the interwar period, European manufacturing workers were often in danger of losing their jobs due to business cycle fluctuations. Employment protection and worker rights in Europe were much weaker than in postwar years, and the social security system was not nearly as well developed. Alternative employment opportunities in informal labour markets were largely nonexistent, although some employment could be found in agriculture and fishing, paying subsistence wages. Empirical analysis of labour markets in the interwar years may thus provide new and interesting evidence on the bargaining power of unions.[2]

Most previous studies have been poorly equipped to identify a stable and well identified relationship, however, 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, provide a limited basis for identifying stable relationships between key variables.[3]

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[1]An exception to these studies is Falch (2001), who estimates well-specified *sector-wise* wage curves for Norwegian teachers for the period 1905-1939.

[2]Likewise, data from the United States indicate a change in the cyclical behaviour of real wages between the interwar period and the postwar years—see Bernanke and Powell (1986) and Hanes (1996). This fact does not necessarily imply that there were changes in the structural parameters of labour demand and supply equations. Such changes could also stem from differences in the relative magnitudes of labour demand and supply shocks in the two time periods, but 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. However, one explanation for these conflicting empirical findings may be that wage formation in interwar labour markets was indeed different from the postwar period.

[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 one can draw inference from the cross-section variation in the data in addition to the time series aspects of the early 1930s. We use a panel data set constructed by Klovland (1999) for Norwegian manufacturing. 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. The wage equations are estimated using the GMM estimator of Arellano and Bond (1991), as well as the system GMM estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998). Both estimators control for the presence of unobserved industry-specific effects and for the possible endogeneity of the explanatory variables.

Our results indicate that unions did have bargaining power during the depression. Moreover, we find theoretically plausible and empirically sound wage equations for the interwar period. Thus, we conclude that, once a more powerful data set and the proper estimation methods are used, wage formation in the interwar period is no longer anomalous.

Section 2 reviews some features of interwar labour markets in Norway that are of specific relevance to the theories examined here. Section 3 presents the estimation results for an unrestricted dynamic model, focusing on the economic interpretation of the long-run results as well as methodological issues related to estimation methods. The model is sufficiently general to nest many specialized 'non-competitive' wage-bargaining models. We therefore test several long-run theory predictions in Section 3.1. The modelling of the dynamic wage equation, subject to long-run constraints, is reported in Section 4. A fuller discussion of the methodological issues is contained in Appendix A, where we present some new Monte Carlo evidence on the properties of the estimators used.

2 Some features of the interwar labour market in Norway

With a deflationary period beginning in 1926 and lasting until 1933, Norway was back on the gold standard at the prewar parity in May 1928.[4] Manufacturing output, which is shown together with the retail price index in Figure 1, was significantly affected by the international depression beginning in the autumn of 1929. The output level of 1929 was not surpassed until 1934, but even this five-year growth pause was a reasonably good performance relative to other countries. The fact that Norway followed pound sterling and went off the gold standard in September 1931 may be a key factor here, as suggested by the international cross-section analysis in Eichengreen and Sachs (1985). In the second half of the 1930s manufacturing output recovered quite well, very much in line with other Scandinavian countries and other Sterling block countries.[5] Increasing labour productivity and capital deepening implied that output could expand significantly without leading to any shortage of labour. Although unemployment went down somewhat in the latter half of the 1930s, it was still high among trade union members in manufacturing at the end of the decade.
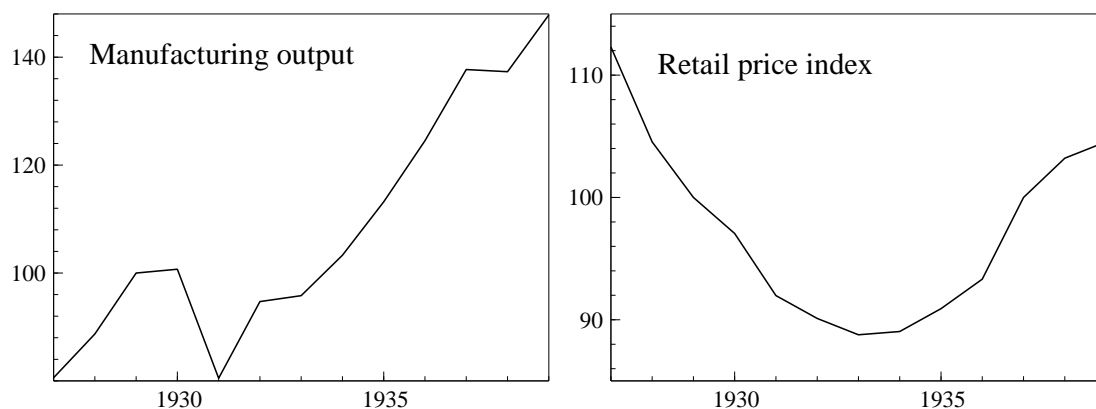

Figure 1: Indices of manufacturing output and retail prices, 1927-1939. 1929=100.

Wage bargaining was organized at a fairly disaggregated industry level. In 1930 there were 32 national trade unions, 21 of which had members working in manufacturing and construction industries. Data on some key labour market characteristics are given in Table

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[4]Klovland (1998) contains some background on the monetary policy in the interwar years.

[5]See Klovland (1997) for new data on manufacturing output in Norway and some international comparisons.

Table 1: Labour market conditions: Norwegian manufacturing, mining and construction 1927-1939.

Year	Degree of unionization (percent of workforce)	Working days lost in labour disputes (days per employed worker)
1927	21.6	5.2
1928	23.1	1.3
1929	26.3	0.6
1930	28.3	0.8
1931	29.5	28.9
1932	30.3	1.3
1933	30.1	1.1
1934	32.4	0.5
1935	37.0	0.4
1936	43.1	1.1
1937	49.1	2.6
1938	50.5	0.7
1939	52.2	0.3

Notes: The number of organized workers and working days lost due to labour market disputes in manufacturing, mining and construction is computed from data found in *Statistical Yearbook of Norway*, Statistics Norway, Oslo, various issues 1929-1941. Annual estimates of the workforce and employed workers in the same industries are taken from Grytten (1994, pp. 166-167).

1. The degree of unionization was low relative to postwar standards, being only about 20 per cent in the mid-1920s but increasing steadily to about 50 per cent towards the end of the 1930s. In the 1920s and early 1930s trade unions were quite militant and employers enforced lockouts on several occasions. Most labour disputes arose in connection with the annual or biannual industry-wide wage negotiations. As can be seen from Table 1, the number of working days lost was nevertheless quite modest in most years; a notable exception is the year of 1931 when a wage dispute in the paper industry spread to the whole manufacturing sector.

We do not have data on replacement ratios. A general scheme of unemployment insurance for manufacturing workers guaranteed by the government was not established until 1938.[6] Before that, only members of trade unions that offered unemployment schemes were entitled to public unemployment benefits. About one third of trade union members did not have access to such schemes. The amounts paid were low, amounting only to about one third of the general wage level in manufacturing.[7] Unorganized workers and members of unions that did not have unemployment schemes were forced to seek public

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[6]Information on labour market institutions is from Grytten (2000).

[7]Grytten (2000, p. 34) concludes that 'it is not likely that the unemployment benefits paid to insured trade unionists gave any significant incentive to stay unemployed'.

relief work in case of unemployment. This was of short duration and poorly paid, being about the same level as unemployment benefits from trade unions. The most important relief work schemes were found in the construction of public roads. Relief workers accounted for about one third of the labour force in public road works in this period. The average hourly wage rates paid to pieceworkers on relief work schemes relative to ordinary workers in public road works are shown in **??** together with average unemployment rates across industries.[8] Wages in public road works were in general much lower than in manufacturing; in 1930 hourly earnings for ordinary workers in public road works were about 32 per cent lower than for skilled workers in the engineering industries, and 18 per cent lower than for unskilled workers.[9] The wage rates paid to relief workers were even lower, mostly in the range of 75 - 80 per cent of ordinary wage rates for pieceworkers in public road construction. In the worst depression years the ratio fell to 65 - 70 per cent.
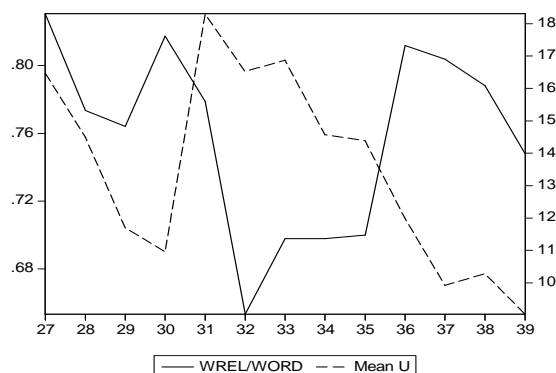

Figure 2: Relief wage relative to ordinary wage for public road works (left axis), together with the average unemployment rates across industries (right axis) over the sample.

If we can generalize from this, it seems the implied expected replacement ratios were most likely low, uncertain, relatively constant, and, as suggested by Figure 2, seemed to covary negatively with unemployment.

Figure 3 describes the distributions of different wage measures and the unemployment rates across the 55 manufacturing industries in the years 1927 to 1939 by means of box-and-whisker plots.[10] The distributions of nominal wages remains fairly constant during

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[8]The data are taken from *Statistiske Meddelelser*, the monthly statistical bulletin of Statistics Norway.

[9]Data on hourly earnings in the engineering industries are from Table 295 of *Historical Statistics 1968*, Statistics Norway, Oslo, 1969.

[10]The lower and upper limits of the box are the 25 and 75 percentiles, while the horizontal lines inside the box denotes the median. The whiskers denote the upper and lower adjacent observations. If x_{75} and x_{25} are the 75 and 25 percentile observations, then observations bigger than $x_{75} + 3/2\,(x_{75} - x_{25})$ and smaller than $x_{25} - 3/2\,(x_{75} - x_{25})$ are outside the adjacent values (and are marked as outside values).
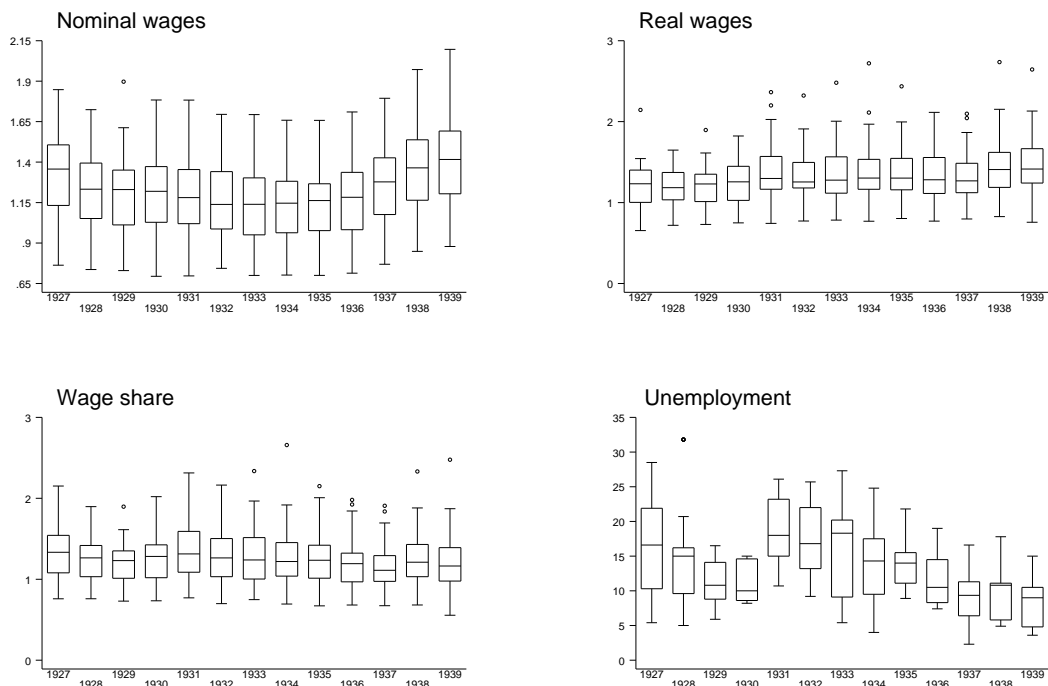
Figure 3: Distribution of nominal wages, real wages, wage shares and unemployment in industries per year.

the depression, rising back towards pre-depression levels in the late thirties. Real product wages show somewhat more dispersion across industries during depression years—but notably so in terms of observed high real wages in some industries, suggesting real wage rigidity. Labour's share of income also displays the same surprisingly stable pattern over the period. Hence, if real wages and wage shares did not exhibit any appreciable downward movements, we would expect labour demand to vary quite a lot — which it does. The lower right panel shows how the unemployment rates increase both in general and across industries as the depression hits the economy, before unemployment rates fall towards the end of the period. The same impression of a strong recession is reflected in the behaviour of retail prices, shown in Figure 1.

The impression of wage rigidity is reinforced when we compare retail prices with the means of wage shares and unemployment rates, shown in Figure 4. While retail prices fall heavily, inflation being positive only after 1933, the mean of labour's share of product income is virtually constant. This observation is consistent with the notion that wages adjust to the "scope" for wage increases—the firm surplus per employee—see Forslund, Gottfries, and Westermark (2008).[11] The mean of the unemployment rates, on the other

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[11] This is consistent with the "main-course" theory of Aukrust (1977). See Bårdsen, Eitrheim, Jansen,
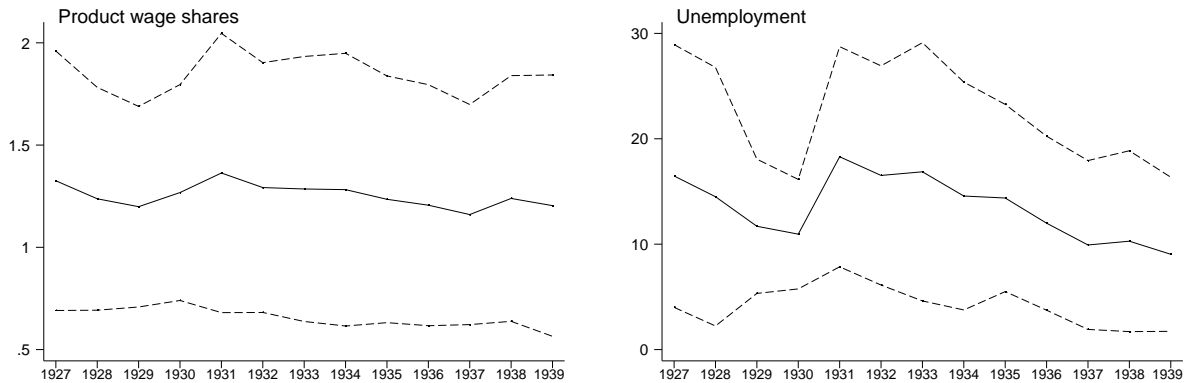
Figure 4: Means $\pm$ two standard errors of product wage shares and unemployment rates across industries over the sample.

hand, changes from 11% 1930 to over just over 18% in 1931.

3 A general dynamic specification of wage determination

There exists a number of specialized models of "non-competitive" wage setting in the literature. Rather then relating to theory-specific models, we want to start by representing the common features of these approaches within a general empirical specification, drawing upon the expositions of Bårdsen, Eitrheim, Jansen, and Nymoen (2005) and Bårdsen and Nymoen (2009). This will then serve as a test-bed for more specialized theory predictions—drawing in particular upon Nymoen and Rødseth (2003) and Forslund, Gottfries, and Westermark (2008).

A dynamic specification general enough to represent these common features[12], and to describe the data presented above, is

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\begin{aligned}
(1 - \alpha_1 L) \, w_{it} = {} & (\beta_0 + \beta_1 L) \, p_{it} + (\gamma_0 + \gamma_1 L) \, q_{it} \\
& + (\delta_0 + \delta_1 L) \, u_{it} + (\zeta_0 + \zeta_1 L) \, pc_t + \eta_i + \varepsilon_{it}.
\end{aligned}
\tag{1}
$$

The variables are (logs of) nominal hourly earnings w_{it}, producer prices p_{it}, labour productivity q_{it}, the unemployment rate u_{it}, and retail prices pc_t.[13] The subscript i denotes

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and Nymoen (2005, Chapter 3) for a recent exposition.

[12]See for example Nymoen and Rødseth (2003) for a similar specification.

[13]We disregard tax rates, which were rather low during the interwar period. We do not have data on replacement ratios.

Table 2: The different instrument sets used in the estimations.

Instrument sets:	GMM instruments	Other instruments
Set 1: GMM estimation, p_{it}, q_{it}, u_{it} exogenous	w_{it-2}, w_{it-3}	$\Delta pc_t, \Delta pc_{t-1}, \Delta p_{it}, \Delta p_{it-1},$ $\Delta q_{it}, \Delta q_{it-1}, \Delta u_{it}, \Delta u_{it-1}$
Set 2: GMM estimation, p_{it}, q_{it}, u_{it} endogenous	$w_{it-2}, w_{it-3}, p_{it-2}, p_{it-3},$ $q_{it-2}, q_{it-3}, u_{it-2}, u_{it-3}$	$\Delta pc_t, \Delta pc_{t-1}$
Set 3: System GMM estimation, p_{it}, q_{it}, u_{it} exogenous	$w_{it-2}, w_{it-3}, \Delta w_{it-1}$	$\Delta pc_t, \Delta pc_{t-1}, \Delta p_{it}, \Delta p_{it-1},$ $\Delta q_{it}, \Delta q_{it-1}, \Delta u_{it}, \Delta u_{it-1}$ For levels: $pc_t, pc_{t-1}, p_{it}, p_{it-1},$ $q_{it}, q_{it-1}, u_{it}, u_{it-1}$
Set 4: System GMM estimation, p_{it}, q_{it}, u_{it} endogenous	$w_{it-2}, w_{it-3}, p_{it-2}, p_{it-3},$ $q_{it-2}, q_{it-3}, u_{it-2}, u_{it-3}$ $\Delta w_{it-1}, \Delta p_{it-1}, \Delta q_{it-1}, \Delta u_{it-1}$	$\Delta pc_t, \Delta pc_{t-1}$ For levels: pc_t, pc_{t-1}

Notes: The variables are defined in Section 3.

the industry, while L is the lag operator: $Lx_{it} = x_{it-1}$. The variables w_{it}, p_{it}, q_{it} and u_{it} are industry-specific.[14] Retail prices pc_t captures economy-wide effects that are not transmitted through the unemployment rate. Nominal wage growth is assumed to respond positively to increases in both producer and (if relevant) retail prices, labour productivity, and negatively to increased unemployment. The dynamic specification is in line with the annual or biannual industry-wide wage negotiations—see Forslund, Gottfries, and Westermark (2008). This specification is in accordance with Forslund, Gottfries, and Westermark (2008) when replacement ratios are held constant.

The wage equation (1) is estimated using the GMM estimator of Arellano and Bond (1991), as well as the system GMM estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998). Both estimators control for the presence of unobserved industry-specific effects and for the possible endogeneity of the explanatory variables.

GMM estimation takes first-differences of the equation to eliminate industry-specific fixed effects. Endogenous variables in levels lagged two or more periods will then be valid instruments, provided there is no autocorrelation in the time-varying component of the error term. This can checked by examining tests for serial correlation in the first-differenced residuals, following Arellano and Bond (1991).

In 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-

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[14]According to Blanchflower and Oswald (1994) local unemployment rates should be the relevant ones. In sparsely populated areas like most of Norway, industrywise and local unemployment are closely related. See the appendix for details of the construction of the unemployment rates.

specific effects under stationarity. This allows the use of lagged differences as instruments for the levels equation.

For each of the two types of GMM estimators, we report two specifications. In the first, productivity q_{it}, producer prices p_{it}, and unemployment u_{it} are treated as exogenous, whereas in the second they are endogenous. This distinction in terms of classification is reflected in the choice of instruments, as shown in Table 2. Note that the retail price index is treated as exogenous throughout. The Appendix provides more detailed examples of the precise form of the instrument matrices. In each case, the validity of the instruments can be tested by means of the Sargan test of over-identifying restrictions. The lag length in the GMM instruments is kept fixed (again see the Appendix) to avoid overfitting, which would remove the effect of instrumental variables estimation.

The estimated wage equations using the different specifications are reported in Table 3. The results are obtained with Ox version 5, Doornik (2007), and the DPD package of Doornik, Arellano, and Bond (2006). In each case, two-step estimation is used, while the reported standard errors and test statistics are asymptotically robust, using the correction of Windmeijer (2005).

All specifications seem to capture the relevant dynamics, since no second order residual correlation is evident. The system estimators produce more reasonable estimates than the first difference estimators. The differences are in particular striking for the lagged dependent variable, 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 parameter of the lagged dependent variable. The system estimator is therefore preferred.

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

Table 3: Wage equations. The dependent wariable is w_{it}.

Estimation method:	GMM estimation		System GMM estimation	
Instrument set:	Set 1	Set 2	Set 3	Set 4
w_{it-1}	0.17 (0.12)	0.26 (0.10)	0.71 (0.05)	0.85 (0.05)
p_{it}	0.06 (0.05)	-0.01 (0.09)	0.03 (0.07)	-0.10 (0.10)
p_{it-1}	0.04 (0.05)	0.08 (0.09)	0.02 (0.06)	0.17 (0.11)
q_{it}	0.14 (0.05)	0.19 (0.07)	0.13 (0.05)	0.11 (0.07)
q_{it-1}	0.06 (0.04)	0.04 (0.05)	-0.05 (0.04)	0.008 (0.05)
u_{it}	0.002 (0.01)	0.007 (0.01)	0.013 (0.016)	0.013 (0.013)
u_{it-1}	-0.02 (0.01)	-0.04 (0.015)	-0.03 (0.01)	-0.04 (0.01)
pc_t	0.18 (0.13)	0.13 (0.19)	0.87 (0.14)	0.94 (0.14)
pc_{t-1}	0.40 (0.17)	0.33 (0.20)	-0.66 (0.13)	-0.83 (0.14)
Diagnostics				
Sargan test: $\chi^2(\cdot)$	38.01^{**} (20)	53.49 (77)	50.97^* (31)	52.62 (121)
$AR(1)$ test: $N(0,1)$	-1.38	-2.54^*	-3.97^{**}	-4.05^{**}
$AR(2)$ test: $N(0,1)$	-1.40	-1.37	0.49	0.17
long-run wage formation: $w_{it}^* = \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_t + \zeta^* pc_t$				
β^*	0.12 (0.07)	0.10 (0.15)	0.17 (0.24)	0.51 (0.52)
γ^*	0.24 (0.07)	0.32 (0.11)	0.28 (0.20)	0.79 (0.41)
δ^*	-0.024 (0.016)	-0.05 (0.02)	-0.07 (0.06)	-0.19 (0.10)
ζ^*	0.69 (0.08)	0.61 (0.15)	0.73 (0.27)	0.71 (0.54)

Notes: The instrument sets are defined in Table 2.

Monte Carlo experiment using a simplified data generating process more relevant for the analysis at hand. The results of the experiment are reported in the Appendix, and they clearly indicate that the system estimator is favoured over the difference estimator—the latter being severely downward biased for the coefficient of the lagged dependent variable.

A final issue relates to the exogeneity assumptions. The exogeneity of the explanatory variables q_{it}, p_{it} and u_{it} is rejected by the Sargan tests, with p-values of 0.008 and 0.0134, respectively. This again supports the system estimator with endogenous regressors. In the following we will therefore concentrate on the system estimator as our preferred specification.

3.1 Testing hypotheses on long-run wage formation

The long-run wage level w_{it}^* is given by

$$w_{it}^* = \left(\frac{\beta_0 + \beta_1}{1 - \alpha_1}\right) p_{it} + \left(\frac{\gamma_0 + \gamma_1}{1 - \alpha_1}\right) q_{it} + \left(\frac{\delta_0 + \delta_1}{1 - \alpha_1}\right) u_{it} + \left(\frac{\zeta_0 + \zeta_1}{1 - \alpha_1}\right) pc_t$$

$$= \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_{it} + \zeta^* pc_t \tag{2}$$

The results from applying the different estimators are given in the bottom part of Table 3. The estimates all have the expected signs across estimators, but the system estimator with endogenous regressors offers the most well-defined long-run wage curve:

$$\hat{w}_{it}^* = \underset{(0.52)}{0.51} p_{it} + \underset{(0.41)}{0.79} q_{it} \underset{(0.10)}{-0.19} u_{it} + \underset{(0.54)}{0.71} pc_t. \tag{3}$$
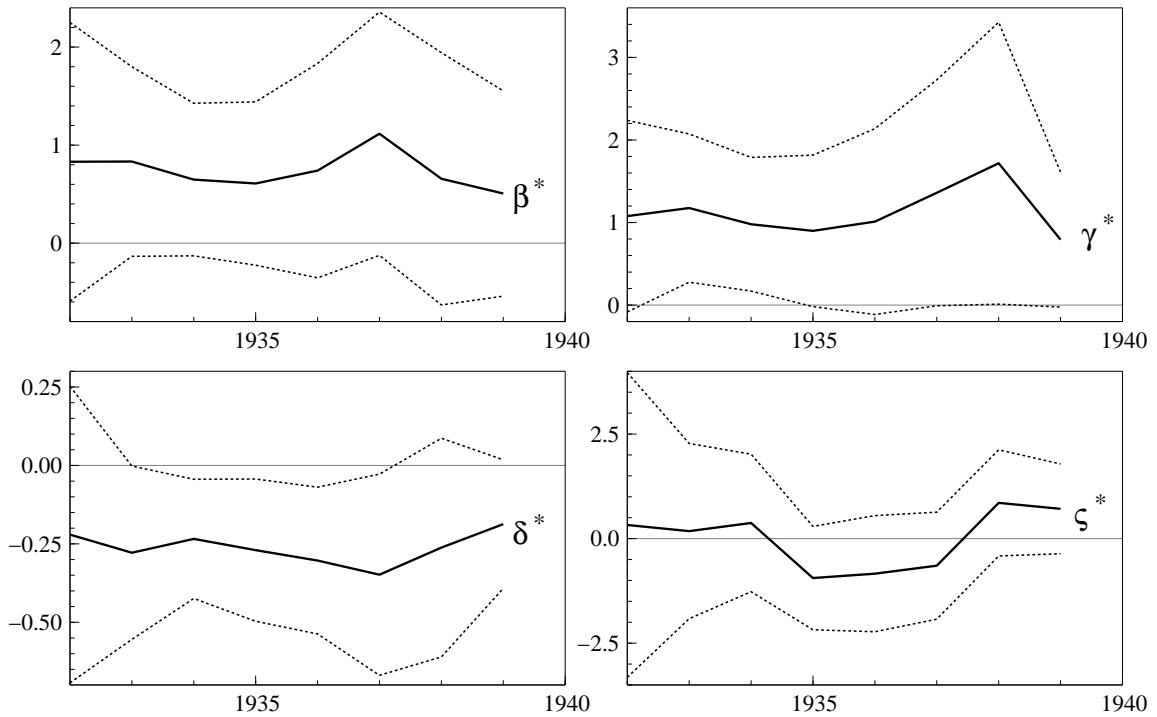

Figure 5: Recursive estimates of the long-run parameters $\pm$ two standard errors.

Given the turbulence of the period that we are investigating, a relevant question is whether this wage curve 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 estimate the long-run solution recursively, also see Johansen (1999). Figure 5 shows that all parameters remain stable across the 1930s, with the exception of the effect of retail prices, which is

insignificant until the latter part of the sample.

Wage formation models based on bargaining theory provide several testable long-run hypotheses, which we can formulate as linear restrictions on the long-run parameters. This approach has been criticized by Boswijk (1993), on the basis of the well-known non-invariance problems of the Wald statistic in testing non-linear restrictions and the lack of efficiency in using unrestricted estimates of the covariance matrix. Consequently, he recommends using a linear reformulation with restrictions valid under the null hypothesis imposed on the covariance matrix. To illustrate the issues involved, we rewrite the model, following the approach of Bårdsen (1989) and using the notation of Boswijk (1993), and consider the model rewritten in differences and levels in a more compact notation:

$$\Delta w_{it} = \pi_0' \Delta \mathbf{x}_{it} + \pi_1 w_{it-1} + \pi_2' \mathbf{x}_{it-1} + \eta_i + \varepsilon_{it},$$

where $\mathbf{x}_i = (p_i \ q_i \ u_i \ pc)'$.

The estimated vector of parameters of (2) are obtained as $\hat{\theta} = -\hat{\pi}_2/\hat{\pi}_1 = (\hat{\beta}^* \ \hat{\gamma}^* \ \hat{\delta}^* \ \hat{\zeta}^*)'$ and the corresponding standard errors are derived from the covariance matrix $\hat{\mathbf{V}}(\hat{\theta}) = \mathbf{J}\hat{\mathbf{V}}(\hat{\pi})\mathbf{J}'$, where $\mathbf{J} = \partial\theta/\partial\pi' = -\pi_1^{-1}(\theta \vdots I_4)$ and $\pi = (\pi_1 \ \pi_2')'$.

We can test the h long-run hypotheses

$$H_0: \ \mathbf{R}\theta = \mathbf{r}$$

using the Wald test statistic

$$W_1 = \left(\mathbf{R}\hat{\theta} - \mathbf{r}\right)' \left(\mathbf{R}\mathbf{J}\hat{\mathbf{V}}(\hat{\pi})\mathbf{J}'\mathbf{R}'\right)^{-1} \left(\mathbf{R}\hat{\theta} - \mathbf{r}\right).$$

However, based on the potential problems noted above, Boswijk recommends instead using the linear formulation

$$H_0': \ \mathbf{r}\pi_1 + \mathbf{R}\pi_2 = \mathbf{Q}\pi = \mathbf{0}$$

and the corresponding Wald test statistic:

$$W_2 = (\mathbf{Q}\hat{\pi})' \left(\mathbf{Q}\hat{\mathbf{V}}(\hat{\pi})\mathbf{Q}'\right)^{-1} (\mathbf{Q}\hat{\pi}).$$

Although both statistics are asymptotically distributed as $\chi^2(h)$, under suitable conditions, Boswijk argues that W_2 is likely to exhibit better behaviour in finite samples. In a small simulation study, Boswijk reports that the statistic W_2 has the better size, but that neither statistic dominates in terms of power. We therefore report both Wald statistics. However, noting the lack of efficiency in using unrestricted estimates of the covariance matrix, after the first test, we use a sequential approach, imposing more and more restrictions, to get more efficient covariance estimates. Since from then on only one additional restriction will be tested at each stage, the statistics will correspond to squared t-ratios.

We start with the most fundamental hypothesis, and then identify the form of the long-run equilibrium wage equation by an increasing number of testable restrictions.

H_{01} **No bargaining power**: $\beta^* = \gamma^* = \zeta^* = 0$

A robust prediction from wage-bargaining models is that the presence of workers' bargaining power is necessary to find a long-run relationship between nominal wages (given expected prices) and productivity —see, for example, Layard, Nickell, and Jackman (1991), Nymoen and Rødseth (2003) and Forslund, Gottfries, and Westermark (2008). The case of no bargaining power corresponds to the case of a vertical Phillips curve. In this case wages react if unemployment differs from the natural rate, but they are not directly affected by the "scope" for wage increases. Since our specification can be written as

$$\Delta w_{it} = \beta_0 \Delta p_{it} + \gamma_0 \Delta q_{it} + \delta_0 u_{it} + \delta_1 u_{it-1} + \zeta_0 \Delta pc_t$$
$$- (1 - \alpha_1)(w_i - \beta^* p_i - \gamma^* q_i - \zeta^* pc_t)_{t-1} + \eta_i + \varepsilon_{it}, \quad (4)$$

this hypothesis corresponds to testing

$$H_{01} : \beta^* = \gamma^* = \zeta^* = 0$$

$$W_1 \sim \chi(3) = 13.81\,[0.003]$$

$$W_2 \sim \chi(3) = 14.41\,[0.002]$$

This hypothesis is therefore clearly rejected with a p-values of 0.003 and 0.002. With the hypothesis of no bargaining power rejected, we can move on to try to identify the

characteristics of the wage curve during the depression by evaluating more specific theory predictions about the form of the equilibrium wage curve.

H_{02} **Price homogeneity**: $\beta^* + \zeta^* = 1$

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 still allowing for some degree of wage level stickiness in the short run. The hypothesis of long-run price homogeneity, $\beta^* + \zeta^* = 1$, is easily tested by estimating the reparameterized long-run solution:

$$w_{it}^* - p_{it} = (1 - \beta^*)(pc_t - p_{it}) + \gamma^* q_{it} + \delta^* u_{it} + (\beta^* + \zeta^* - 1) pc_t$$

$$\widehat{w_{it}^* - p_{it}} = \underset{(0.58)}{0.49}(pc_t - p_{it}) + \underset{(0.41)}{0.79} q_{it} \underset{(0.10)}{-0.19}\, u_{it} + \underset{(0.31)}{0.22} pc_t$$

The corresponding test statistics are:

$$H_{02} : \beta^* + \zeta^* = 1$$

$$W_1 \sim \chi(1) = 0.41\,[0.52]$$

$$W_2 \sim \chi(1) = 0.53[0.47]$$

This restriction is not rejected and is therefore imposed on the long-run wage curve in the following.

H_{03} **Price homogeneity in product prices**: no wedge term

According to the model of Forslund, Gottfries, and Westermark (2008), there should not be a wedge term in a wage equation representing wage-bargaining at a disaggregated level. With price homogeneity already imposed, this corresponds to testing $\beta^* = 1$ in the specification:

$$w_{it}^* - p_{it} = (1 - \beta^*)(pc_t - p_{it}) + \gamma^* q_{it} + \delta^* u_{it}$$

$$\widehat{w_{it}^* - p_{it}} = \underset{(0.48)}{0.54}(pc_t - p_{it}) + \underset{(0.39)}{0.79} q_{it} \underset{(0.14)}{-0.21}\, u_{it}$$

$$H_{03} : \beta^* = 1 \mid \beta^* + \zeta^* = 1$$

$$W_1 \sim \chi(1) = 1.26 \, [0.26]$$

$$W_2 \sim \chi(1) = 0.99 \, [0.32]$$

Accordingly, the prediction of Forslund, Gottfries, and Westermark (2008) cannot be refuted on this data-set.

H_{04} **Productivity gains are fully reflected in wages, conditional on price homogeneity and no wedge term**:

Adding the restriction of proportionality of productivity, $\gamma^* = 1$, gives a model where wages are determined by "scope", so in equilibrium the wage share is a function of the unemployment rate:

$$w_{it}^* - p_{it} - q_{it} = (\gamma^* - 1) q_{it} + \delta^* u_{it}$$

$$\widehat{w_{it}^* - p_{it}} - q_{it} = \underset{(0.37)}{-0.002} q_{it} - \underset{(0.12)}{0.25} u_{it}$$

This is a robust prediction of bargaining models—see Forslund, Gottfries, and Westermark (2008) and Nymoen and Rødseth (2003) as recent examples—but it is also consistent with the Aukrust (1977) main-course model. We note again that the additional restriction cannot be rejected:

$$H_{04} : \gamma^* = 1 \mid \beta^* = 1, \ \zeta^* = 0$$

$$W_1 \sim \chi(1) = 2.63 \times 10^{-5} \, [1.00]$$

$$W_2 \sim \chi(1) = 2.63 \times 10^{-5} \, [1.00]$$

The restricted empirical representation of (2) is therefore estimated as

$$w_{it}^* - p_{it} - q_{it} = \delta^* u_{it}$$

$$\widehat{w_{it}^* - p_{it}} - q_{it} = \underset{(0.12)}{-0.25} u_{it} \qquad (5)$$

The estimate of the unemployment response $\hat{\delta}^* = -0.25$ implies that a 10% increase in unemployment, from 10 to 11 percentage points, will reduce the wage share by 2.5%. How-

ever, the parameter is not very precisely estimated and a bit higher than usually found, which tends to be around -0.1, as documented by Blanchflower and Oswald (1994). For example, Johansen (1995b) report -0.07 on Norwegian industry data, Forslund, Gottfries, and Westermark (2008) report a lower bound of the unemployment effect of -0.12 in their Scandinavian panel specification, while Nymoen and Rødseth (2003) report an average wage elasticity with respect to unemployment of -0.13 for the Nordic countries.

Following Boswijk (1993), the 95% confidence intervals $C(W_i)$ are defined as the set of values not rejected using W_1 and W_2 as test statistics:

$$C(W_1(\delta^* \mid \beta^* = \gamma^* = 1, \ \zeta^* = 0)) = [-0.48, \ -0.02]$$

$$C(W_2(\delta^* \mid \beta^* = \gamma^* = 1, \ \zeta^* = 0)) = [-0.74, \ -0.085].$$

Both confidence intervals are too wide to provide any further guidance. We therefore choose to impose the point estimate of -0.25 when we next turn to estimating the dynamic specification in the equilibrium correction form given by (6).

Finally, the joint test statistics of imposing all the restrictions on the long-run wage equation as $w_{it}^* = p_{it} + q_{it} - 0.25 u_{it}$ are given as

$$H_{05} : \gamma^* = 1 = \beta^* = 1, \ \zeta^* = 0, \ \delta^* = -0.25$$

$$W_1 \sim \chi(4) = 2.18\,[0.70]$$

$$W_2 \sim \chi(1) = 1.89\,[0.76]\,,$$

which are in accordance with the earlier results.

4 The dynamic model

The equilibrium correction reparameterization of (1), conditional on imposed restrictions on (2), becomes:

$$\Delta w_{it} = \beta_0 \Delta p_{it} + \gamma_0 \Delta q_{it} + \delta_0 \Delta u_{it} + \zeta_0 \Delta pc_t - (1 - \alpha_1)(w - w^*)_{it-1} + \eta_i + \varepsilon_{it}, \quad (6)$$

The dynamic specifications are reported in Table 4. Column (1) contains the general

17

Table 4: Wage equations, GMM system estimates

Dep. var: Δw_{it}	(1)	(2)
Δp_{it}	-0.05 (0.09)	–
Δq_{it}	0.14 (0.05)	0.15 (0.05)
Δu_{it}	0.01 (0.01)	–
Δpc_t	0.83 (0.11)	0.75 (0.06)
$(w - w^*)_{it-1}$	-0.13 (0.02)	-0.14 (0.02)
Diagnostics		
Sargan: $\chi^2\,()$	$53.80\,(125)$	$54.65\,(127)$
$AR\,(1)$	-4.19^{**}	-4.13^{**}
$AR\,(2)$	0.18	0.35

model reparameterized in equilibrium-correction form, with the long-run solution (5) 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 0.35. The final model is reported in column (2). There is substantial nominal rigidity, as measured by the equilibrium correction coefficient with a value of -0.14, which is somewhat lower than the average European result of -0.25 reported by Blanchard and Katz (1999). Consequently, a drop in inflation is not likely to be reflected in a similar drop in wage growth, as documented by the coefficient of 0.75 on retail inflation.[15] These magnitudes are similar to the evidence from time-series studies using recent Norwegian manufacturing data by Nymoen (1989) and Johansen (1995a), as well as the panel studies of Johansen (1996) and Wulfsberg (1997).

Summing up so far, we have found a theoretically plausible and empirically sound wage equation for the interwar period in Norway, rejecting the hypothesis of no bargaining power during the depression. The next question is whether the short-run adjustment of wages during the interwar period differed from what is found in empirical 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.

It might again be argued that perhaps such results dominate in the latter half of the sample, as Norway recovered from the great depression, instead of reflecting actual

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[15]Since retail prices and unemployment rates are measured at a higher level of aggregation than wages, a downward bias in the standard errors is possible—see Moulton (1986). However, given the very high precision of the estimates of the final model, the outcome of testing the main hypothesis should not be affected.

behaviour during the depressed years in the early 1930s. To investigate this possibility we complete the analysis with recursive estimation of our preferred equation in column (2). The estimated coefficients, together with their approximate confidence bands, are shown in Figure 6, starting from 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.
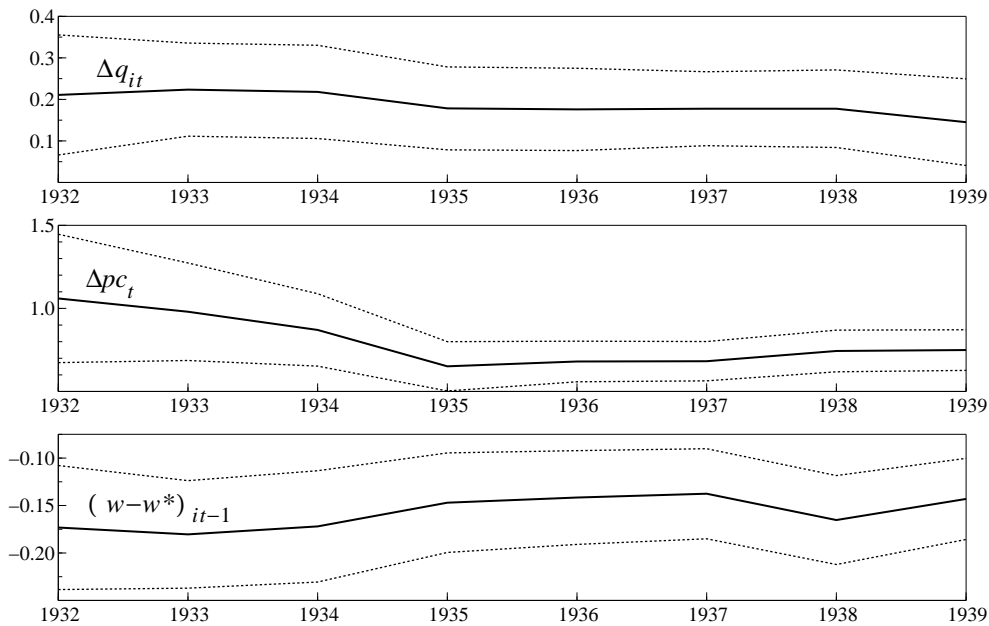

Figure 6: Recursive estimates of the model parameters ± two standard errors.

Finally, we briefly reconsider two questions by means of our chosen specification. The first is whether a Phillips curve can be encompassed by our model; the second is whether the long-run coefficient of unemployment can be restricted to -0.1, as advocated by Blanchflower and Oswald (1994). Both questions are answered by adding u_{it-1} as a regressor:

$$\Delta \hat{w}_{it} = \underset{(0.0533)}{0.1427} \ \Delta q_{it} + \underset{(0.0605)}{0.7498} \ \Delta pc_t - \underset{(0.0347)}{0.1356} \ (w - w^*)_{it-1} - \underset{(0.0143)}{0.003741} \ u_{it-1}$$

A Phillips curve would occur if the equilibrium-correction term $(w - w^*)_{it-1}$ could be omitted—a hypothesis which is clearly rejected.[16]

The confidence intervals reported in Section 3.1 for the long-run effect of unemploy-

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[16]If we have an expectations-augmented (wage) Phillips curve with backward-looking expectations, we would expect the lagged nominal wage increase to enter as a regressor. However, this term did not enter significantly.

ment were very wide and included $\delta^* = -0.1$. This hypothesis can now be retested with hopefully greater power by estimating the final specification under the null hypothesis $w_{it}^{*0} = p_{it} + q_{it} - 0.1u_{it}$:

$$\Delta\hat{w}_{it} = \ \underset{(0.0533)}{0.1427} \ \Delta q_{it} + \ \underset{(0.0605)}{0.7498} \ \Delta pc_t - \ \underset{(0.0347)}{0.1356} \ \left(w - w^{*0}\right)_{it-1} - \ \underset{(0.0105)}{0.02544} \ u_{it-1}$$

$$H_{06} : \delta^* = -0.1 \mid \gamma^* = 1 = \beta^* = 1, \ \zeta^* = 0$$

$$W_2 = \left(\frac{-0.02544}{0.0105}\right)^2 \sim \chi\left(1\right) = 5.\,87\left[0.01\right],$$

which is now rejected.

Our results therefore indicate that unemployment has a bigger effect on wages during the depression than the response usually found in post-World War II studies, as cited earlier. One simple explanation for this might be that replacement ratios are not included in our model and biases the results. However, as already indicated in Section 2, there are some interesting complications regarding the likely relationship between replacement ratios and unemployment rates during the depression. First, the probability of getting relief work P_R could be decreasing with rising unemployment U, so $P_R = P_R\left(\underset{-}{U}\right)$. Second, the wage rate for relief work W_R seemed to vary negatively with rising unemployment, $W_R = W_R\left(\underset{-}{U}\right)$, as already suggested by Figure 2.[17] We then have that the log of the (expected) replacement ratio RR is a function of the log of the unemployment rate:

$$\log(RR\left(U\right)) = \log\left(\frac{P_R\left(\underset{-}{U}\right) \times W_R\left(\underset{-}{U}\right)}{W\left(\underset{-}{U}\right)}\right) \approx \varphi_0 - \varphi_1 u,$$

say. This means that the effect of unemployment in our model is a "gross" effect, $\delta^* = (\phi + \varphi_1)$, capturing both the effect of unemployment on bargaining ϕ, holding the replacement ratio constant, and the influence through a lower replacement ratio.

Therefore, when asked to compare our coefficient of -.25 with results of other studies, like the Blanchflower-Oswald result of -.1, the difference could be attributed to a negative

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[17]A likely explanation could be that more people could get some income that way.

effect from the replacement ratio during the depression.

5 Conclusions

Our empirical analysis shows that wage formation in the interwar period can be understood with the help of modern bargaining theory and well established wage equations. In the case of Norwegian manufacturing industries during the interwar years, the preferred steady-state wage equation features the standard properties of homogeneity with respect to product prices and productivity—wages respond to "scope" and there is an unemployment elasticity of -0.25. 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 stems from 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.

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A A simulation experiment of the properties of the estimators

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

$$
\begin{aligned}
y_{it} &= \alpha y_{it-1} + \beta x_{it} + \eta_i + v_{it}, \quad \eta_i \sim \mathsf{IN}[0, \sigma_\eta^2] \quad v_{it} \sim \mathsf{IN}[0, \sigma_v^2] \\
&\quad i = 1, \ldots, N, \ t = 1, \ldots, T \\
x_{it} &= \rho x_{it-1} + e_{it}, \quad e_{it} \sim \mathsf{IN}[0, \sigma_e^2].
\end{aligned}
$$

This DGP is used in Doornik, Arellano, and Bond (2006) 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 and $\sigma_\eta^2 = \sigma_v^2 = 1$. It was also noted that $Diff$ underestimates α, whereas Sys produces an overestimate. While Doornik, Arellano, and Bond (2006) keep β fixed at unity, we now proceed to keep α fixed at 0.9, and vary β. We use $N = 100$, and $T = 7$ (5 after allowing for lags and differences), as well as $N = 55, T = 13$ (corresponding to the empirical application.)

Kiviet (2007) criticizes this experiment for the limited exploration of the parameter space. This applies in particular to keeping $\sigma_\eta^2 = \sigma_v^2 = 1$ throughout. He also points out that the specification of the initial conditions is important because the time-series dimension is short and fixed. In our experiments we start from zero, but discard the first twenty observations. This will closely resemble using stationary initial conditions.

The two estimators can be summarized as:

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

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

$$
Z_i = \begin{pmatrix}
y_{i1} & 0 & 0 & 0 & 0 & \Delta x_{i,3} & 1 \\
0 & y_{i1} & y_{i2} & 0 & 0 & \Delta x_{i,4} & 1 \\
0 & 0 & 0 & y_{i2} & y_{i3} & \Delta x_{i,5} & 1
\end{pmatrix}.
$$

This assumes that initially the available observations are $t = 1, \ldots, 5$. 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:

$$
Z_i^* = \begin{pmatrix} y_{i1} & 0 & 0 & 0 & 0 & \Delta x_{i,3} \\ 0 & y_{i1} & y_{i2} & 0 & 0 & \Delta x_{i,4} \\ 0 & 0 & 0 & y_{i2} & y_{i3} & \Delta x_{i,5} \end{pmatrix}, \quad Z_i^+ = \begin{pmatrix} \Delta y_{i2} & 0 & 0 & x_{i,3} & 1 \\ 0 & \Delta y_{i3} & 0 & x_{i,4} & 1 \\ 0 & 0 & \Delta y_{i4} & x_{i,5} & 1 \end{pmatrix}
$$

Some results for $N = 55, T = 13$ using $M = 1000$ Monte Carlo replications are presented in Figure 7. 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) (and an additional regressor). In both cases, the sample size is different as well.
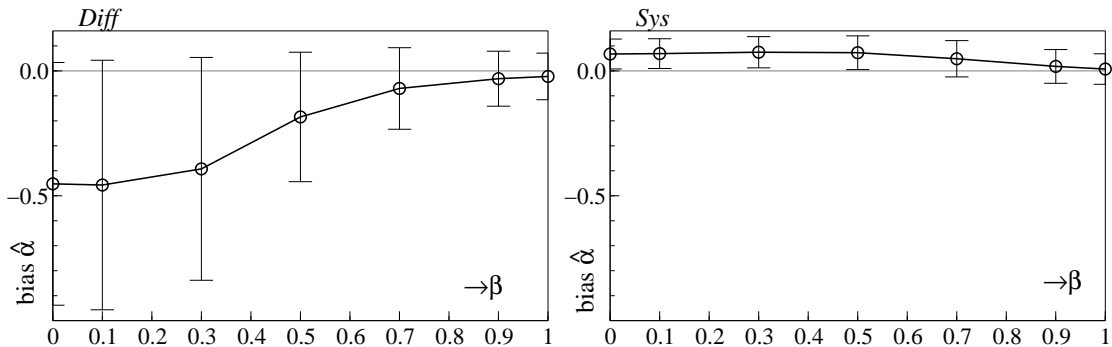

Figure 7: Mean bias of $\hat{\alpha}$, $N = 55, T = 13$, $M = 1000, \alpha = 0.9, \rho = 0.8, \sigma_e^2 = 0.9$, $\sigma_\eta^2 = \sigma_v^2 = 1$; bars are twice the MCSD; $\beta = 0, 0.1, 0.3, 0.5, 0.7, 0.9, 1$.

These 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, instead of the true value 0.9. Sys overestimates α, but is much better behaved. These results shed some light on Table 3: the large discrepancy between the $Diff$ and Sys results reported there corresponds to a low value of β in Figure 7. 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.

Finally, we address some of the points raised by Kiviet (2007). First, both α and β are varied. Secondly, we choose three variance ratios. Kiviet (2007) recommends doing

this relative to the autoregressive parameter as:

$$\psi^2 = \frac{(1 + \alpha)}{(1 - \alpha)} \frac{\sigma_\eta^2}{\sigma_v^2}.$$

Three values are selected: $\psi^2 = 1$, corresponding to $\sigma_\eta^2 = \sigma_v^2 = 1$ when $\alpha = 0$; $\psi^2 = 19$ for a variance ratio of unity when $\alpha = 0.9$, and $\psi^2 = 190$. Figure 8 shows the results, with *Diff* in the top row, and *Sys* in the bottom row. Each row uses the same scale for the bias on the Z-axis. The *Diff* has substantial downward biased for the autoregressive parameter for high α and low β. The *Sys* estimates always over-estimate, but are are much better, except for the very high value of ψ. In that case, the bias is between 0.1 amd 0.2 for low to medium α (and largely independent of β).
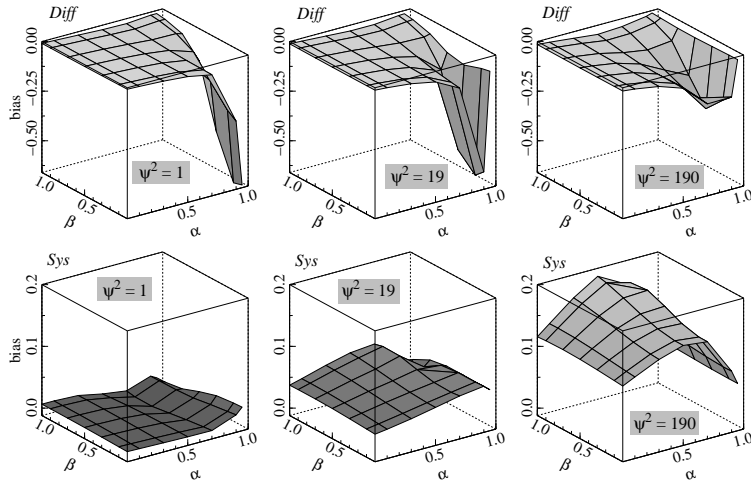

Figure 8: Mean bias of $\hat\alpha$, $N = 100, T = 5$, $M = 1000, \rho = 0.8, \sigma_e^2 = 0.9$.

B   The Data

The data-set is downloadable from `http://www.svt.ntnu.no/iso/Gunnar.Bardsen/default.htm`. 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.