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by

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# A European-type wage equation from an American-style labor market: Evidence from a panel of Norwegian manufacturing industries in the 1930s

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20 March 2004

**Abstract**

Using a newly constructed panel of manufacturing industry data for interwar Norway, we estimate a long-run wage curve for the 1930s that has all the modern features of being homogeneous in prices, proportional to productivity, and having an unemployment elasticity of $-0.1$. This result is more typical of contemporary European than U.S. wage equations, even if the labour market in interwar Norway possessed distinctively more 'American' features than those associated with present-day European welfare states. We also present some new Monte Carlo evidence on the properties of the estimators used.

*JEL Classification: E24,N24*

*Keywords: wages, depression, panel data, dynamics*

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# 1 Introduction

There are two main empirical approaches to the explanation of wage behavior. First, the dynamic *Phillips curve*, giving a negative relationship between wage *growth* and the unemployment rate, has a prominent role in the empirical literature. The second approach is the dynamic *wage curve,* which gives a negative long-run relationship between the wage *level* and the unemployment rate. The empirical evidence favours the Phillips curve specification for the US, while wage curve specifications dominate the European literature. Blanchflower and Oswald (1994) in particular have made a strong case for the wage curve as a general phenomenon. While they also report wage-curve specifications on US data, their results are refuted by Blanchard and Katz (1997).

Blanchard and Katz (1997, 1999) provide an elegant attempt at reconciling the conflicting evidence by utilizing the fact that the Phillips curve is nested within the wage-curve specification. This makes it easy to discriminate empirically between the two models. To illustrate their point, consider the following stylized wage-curve specified as an error correction model (ECM):

$$\Delta w_t = c + \Delta pc_t + \alpha \Delta q_t - \delta u_t - \alpha \left[ w - pc - q \right]_{t-1} + \varepsilon_t, \tag{1}$$

where the variables are nominal hourly earnings $W$, labor productivity $Q$, the unemployment rate $U$, and retail prices $PC$. Lowercase letters denote natural logarithms of the corresponding variables denoted in capitals, so $x_t \equiv \ln X_t$ and growth rates are given as $\Delta x_t \equiv x_t - x_{t-1}$. Blanchard and Katz (1999) argue that

$$\alpha = (1 - \mu\lambda), \quad 0 \leq \{\mu, \lambda\} \leq 1,$$

where $(1 - \lambda)$ is the direct effect of productivity on the expected real wage, and $(1 - \mu)$ is the direct effect of productivity on the reservation wage. Thus, if there are no effects from productivity, so that $\mu = \lambda = 1$, the ECM-term $[w - pc - q]_{t-1}$ drops out and the Phillips-curve specification remains.

According to Blanchard and Katz (1999), underlying labor market conditions and institutional settings are the crucial determinants of wage behavior with systematic structural differences between Europe and the United States. Productivity effects on wages are assumed to be higher in Europe than in the US, which implies a small magnitude of $\mu$ and $\lambda$. This explains the presence of an ECM-term in European equations. The small magnitude of $\mu$ is related to the greater role of unions and more stringent hiring and firing regulations in the European labor markets. The smaller European $\lambda$ could be caused by a bigger informal sector, although the evidence here is less well documented. A more general statement is perhaps that $\lambda$ will be higher the weaker the rights of workers, and that $\mu$ will be higher the more diverse the total labor market is.

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 labor markets were largely nonexistent, although some employment could be found in agriculture and fishing, paying subsistence wages. In many respects, interwar European labor markets possess features that are closer to typical American labor settings than to present-day European markets. Empirical analysis of European labor markets in the interwar years may thus provide new and interesting evidence on the two conflicting hypotheses. According to the explanation given by Blanchard and Katz (1999), we should expect to find a Phillips curve rather than a wage curve when looking at European data for the interwar years. On the other hand, it could be that the wage curve model emerges as the best specification. In that case, other theories are called for to explain the differences between US and European wage setting.

When looking at evidence to date from the interwar period, the empirical wage equations appear to be somewhat fragile. For the United Kingdom, Hatton (1988), Dimsdale et al. (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 there is only a weak feedback from unemployment to the real wage. One explanation for these conflicting empirical findings may be that wage formation in interwar labor markets was indeed different from the postwar period, thus supporting Blanchard and Katz (1999). Data from the United States indicate a change in the cyclical behavior 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 labor demand and supply equations. Such changes could also stem from differences in the relative magnitudes of labor demand and supply shocks in the two time periods.[2]

Below, we report empirical evidence on interwar wage equations for one European country, Norway, using GMM estimation methods. Our purpose is twofold: to show that theoretically plausible and

[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.

empirically sound wage equations can be found for the interwar period, once a more powerful data set is available and the proper estimation methods are applied. This will then allow us to test the hypothesis of Blanchard and Katz (1999)—that the existence of a wage curve is dependent upon the presence of modern 'European' type labor market settings.

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, provide a limited basis for identifying stable relationships between key variables.[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. 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 aspects of the early 1930s. The data base contains annual values of key output and labor market variables for 55 manufacturing industries over the period 1927 to 1939: nominal average hourly earnings, producer price indices, labor productivity (real value added per hour) and, at a somewhat less disaggregated level, unemployment rates.

Section 2 briefly presents the general model, which is sufficiently general to encompass wage behavior in this period. Section 3 reviews some features of interwar labor markets in Norway that are of specific relevance to the theories examined here. We report the empirical modelling of the wage equation for the years 1927 – 1939 in Section 4, focusing on the economic interpretation of the results as well as methodological issues related to estimation methods. 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 The wage equation

A general dynamic specification, nesting equation (1), is

$$\begin{aligned}(1-\alpha_1 L)\, w_{it} &= (\beta_0+\beta_1 L)\, p_{it} + (\gamma_0+\gamma_1 L)\, q_{it} \\ &\quad + (\delta_0+\delta_1 L)\, u_{it} + (\zeta_0+\zeta_1 L)\, pc_t + \eta_i + \varepsilon_{it}. \qquad (2)\end{aligned}$$

[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.

The variables are (logs of) nominal hourly earnings $w$, producer prices $p$, labor productivity $q$, the unemployment rate $u$, and retail prices $pc$.[4] The subscript $i$ denotes the industry, while $L$ is the lag operator: $Lx_{it} = x_{i,t-1}$. The variables $w_{it}$, $p_{it}$, $q_{it}$ and $u_{it}$ are industry-specific, while $pc_t$ captures economy-wide effects that are not transmitted through the unemployment rate.

Nominal wage growth responds positively to increases in producer and retail prices, labor 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. A key hypothesis, subjected to empirical testing below, is that productivity growth increases real wages in the same proportion in the long run. The equivalent to Blanchard and Katz' model in (1) is the ECM reparameterization of (2):

$$\Delta w_{it} = \beta_0 \Delta p_{it} + \gamma_0 \Delta q_{it} + \delta_0 \Delta u_{it} + \zeta_0 \Delta pc_t - \alpha_1 \left(w - w^*\right)_{i,t-1} + \eta_i + \varepsilon_{it}, \tag{3}$$

where $w^*_{it}$ is the steady-state wage level

$$\begin{aligned} w^*_{it} &= \left(\frac{\beta_0 + \beta_1}{1-\alpha_1}\right) p_{it} + \left(\frac{\gamma_0 + \gamma_1}{1-\alpha_1}\right) q_{it} + \left(\frac{\delta_0 + \delta_1}{1-\alpha_1}\right) u_t + \left(\frac{\zeta_0 + \zeta_1}{1-\alpha_1}\right) pc_t \\ &= \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_t + \zeta^* pc_t. \end{aligned} \tag{4}$$

Price level homogeneity requires that $\beta^* + \zeta^* = 1$, while the long-run proportionality of labor productivity implies $\gamma^* = 1$. Institutional and structural features are reflected in the coefficients of (4). Changes in the impact of institutions on wage setting can therefore be tested by looking at the empirical stability of (4) over the sample period. It is quite likely that wages interact simultaneously with all the explanatory variables. We do, of course, take the possible simultaneity into account when estimating the model by using instrumental variables.

# 3 Some features of the interwar labor market in Norway

After a deflationary period in the mid 1920s, Norway was back on the gold standard at the prewar parity in May 1928.[5] Manufacturing output, which is shown 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

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

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

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.[6] Increasing labor productivity and capital deepening implied that output could expand significantly without leading to any shortage of labor. 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.

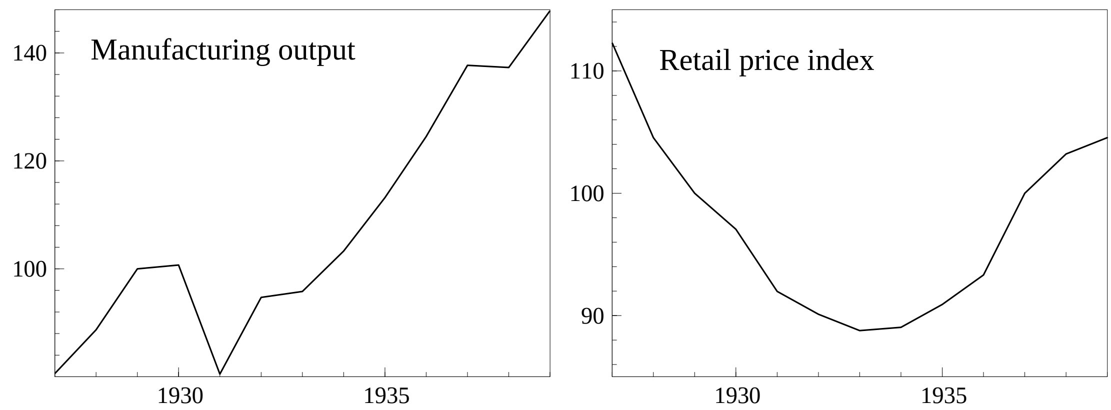


Figure 1: Indices of manufacturing output and retail prices, 1927-1939. 1929=100.

A general scheme of unemployment insurance for manufacturing workers guaranteed by the government was not established until 1938.[7] Before that, only members of trade unions that offered unemployment schemes were entitled to unemployment benefits. About one third of trade union members did not have access to such schemes. The amounts paid were low and fairly constant in real terms, amounting only to about one third of the general wage level in manufacturing. 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'. Furthermore, the level of unionization was relatively modest: roughly one in four workers were trade union members. Unorganized workers and members of unions that did not have unemployment schemes were forced to seek public 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. Thus reservation wages in the manufacturing industries

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

[7] Information on labour market instutions is from Grytten (2000).

were roughly constant at a low level, and apparently not very sensitive to productivity increases, which is the crucial link in the theory considered here.

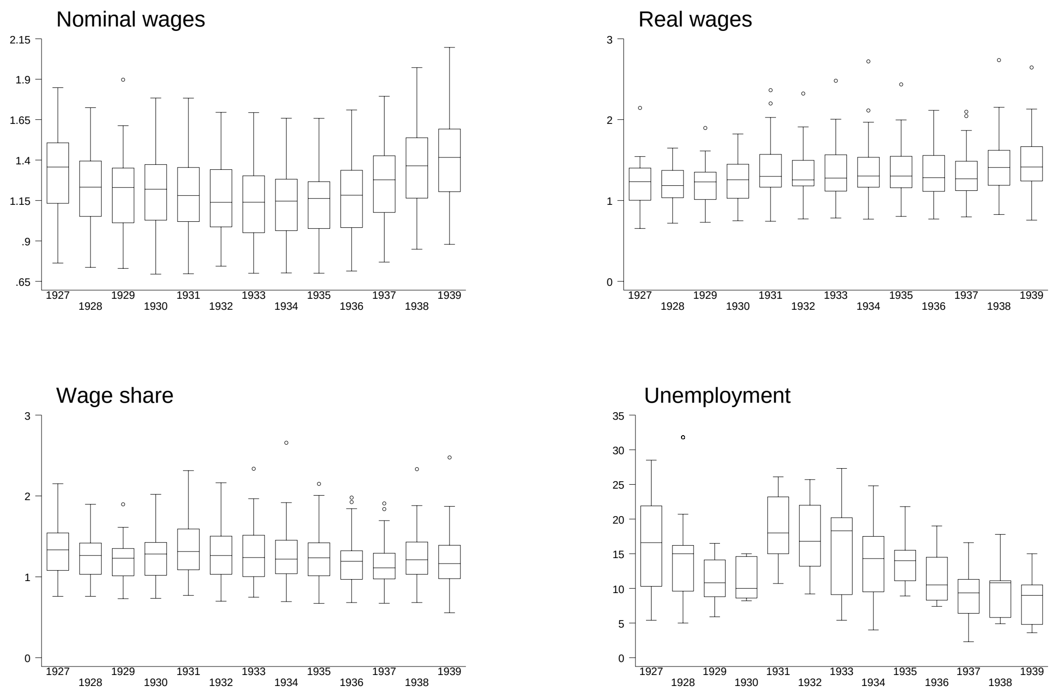


Figure 2: Distribution of nominal wages, real wages, wage shares and unemployment in industries per year.

Figure 2 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.[8] The distributions of nominal wages remains fairly constant during the depression. The median displays considerable downward rigidity, 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. Real wage rigidity is even more pronounced than nominal rigidity. Labor'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,

[8]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).

we would expect labor 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 behavior of retail prices, shown in Figure 1.


Figure 3: Means ± two standard errors of product wage shares and unemployment rates across industries over the sample.

The impression of wage rigidity is further reinforced when we compare retail prices with the means of wage shares and unemployment rates, shown in Figure 3. While retail prices fall heavily, inflation being positive only after 1933, the mean of labor's share of product income is virtually constant. The mean of the unemployment rates, on the other hand, nearly doubles from 1930 to 1931.

# 4 Empirical results

The wage equation (2) 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

Table 1: The different specifications considered

| | GMM instruments | Other instruments |
|---|---|---|
| Differenced | $w_{it-2}, w_{it-3}$ | $\Delta pc_t, \Delta pc_{t-1}, \Delta p_{it}, \Delta p_{it-1}$, $\Delta q_{it}, \Delta q_{it-1}, \Delta u_{it}, \Delta u_{it-1}$ |
| Differenced | $w_{it-2}, w_{it-3}, p_{it-2}, p_{it-3}$, $q_{it-2}, q_{it-3}, u_{it-2}, u_{it-3}$ | $\Delta pc_t, \Delta pc_{t-1}$ |
| Differenced & levels | $w_{it-2}, w_{it-3}, \Delta w_{it-1}$ | $\Delta pc_t, \Delta pc_{t-1}, \Delta p_{it}, \Delta p_{it-1}$, $\Delta q_{it}, \Delta q_{it-1}, \Delta u_{it}, \Delta u_{it-1}$ For levels: $pc_t, pc_{t-1}, p_{it}, p_{it-1}$, $q_{it}, q_{it-1}, u_{it}, u_{it-1}$ |
| Differenced & levels | $w_{it-2}, w_{it-3}, p_{it-2}, p_{it-3}$, $q_{it-2}, q_{it-3}, u_{it-2}, u_{it-3}$ $\Delta w_{it-1}, \Delta p_{it-1}, \Delta q_{it-1}, \Delta u_{it-1}$ | $\Delta pc_t, \Delta pc_{t-1}$ For levels: $pc_t, pc_{t-1}$ |

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

In addition to the two types of GMM estimators, we consider 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 1. 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.

## 4.1 Empirical wage equations

The estimated wage equations using the different specifications are reported in Table 2. The results are obtained with Ox version 3.2 (see Doornik, 1999) and the DPD package (Doornik et al., 1999). In each case, two-step estimaton is used, while the reported standard errors and test statistics are asymptotically robust to general heteroskedasticity, see Windmeijer (2000).

Table 2: Wage equations, GMM estimates

| Dep. var: $w_{it}$ | Differences | | Differences & levels | |
|---|---|---|---|---|
| | exogenous | endogenous | exogenous | endogenous |
| $w_{it-1}$ | 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) |
| $u_{it}$ | 0.001 (0.01) | 0.007 (0.01) | 0.013 (0.016) | 0.013 (0.013) |
| $u_{it-1}$ | −0.02 (0.01) | −0.04 (0.01) | −0.03 (0.01) | −0.04 (0.01) |
| $pc_t$ | 0.18 (0.13) | 0.13 (0.19) | 0.86 (0.14) | 0.94 (0.13) |
| $pc_{t-1}$ | 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 |
| Steady-state analysis: $w_{it}^* = \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_t + \zeta^* pc_t$ | | | | |
| $\beta^*$ | 0.12 (0.07) | 0.10 (0.14) | 0.17 (0.24) | 0.51 (0.52) |
| $\gamma^*$ | 0.24 (0.07) | 0.31 (0.11) | 0.28 (0.20) | 0.79 (0.41) |
| $\delta^*$ | −0.024 (0.016) | −0.05 (0.02) | −0.07 (0.06) | −0.19 (0.10) |
| $\zeta^*$ | 0.69 (0.08) | 0.61 (0.15) | 0.73 (0.26) | 0.71 (0.53) |
| Testing steady-state restrictions | | | | |
| $\beta^* + \zeta^* = 1$ | $12.61^{**}$ | $16.90^{**}$ | 0.28 | 0.41 |
| $\beta^* + \zeta^* = 1$, $\gamma^* = 1$ | $128.09^{**}$ | $46.39^{**}$ | $13.97^{**}$ | 0.95 |
| $\beta^* + \zeta^* = 1$, $\gamma^* = 1, \delta^* = -0.1$ | $246.30^{**}$ | $102.33^{**}$ | $27.16^{**}$ | 1.80 |

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 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. 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 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 favored 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, which is now our preferred specification.

## 4.2 The steady state

As argued in the introduction, the crucial hypothesis to be tested is the significance of the parameters of the long-run solution (4)

$$w_{it} = \beta^* p_{it} + \gamma^* q_{it} + \delta^* u_t + \zeta^* pc_t.$$

The hypothesis of long-run price homogeneity, $\beta^* + \zeta^* = 1$, is rejected in both differenced specifications, using the approach of Bårdsen (1989). On the other hand, the system models do not reject the hypothesis. Adding the restriction of proportionality of productivity, $\gamma^* = 1$, we see that this is only accepted by the system specification with endogenous regressors. Clearly, the bias problems of the other specifications have a large impact on their steady-state estimates. In the remainder we will only consider our preferred model based on the system specification with endogenous regressors.

The evidence presented in Table 2 does not lend any support to the hypothesis that the existence of a wage curve is dependent upon 'modern European' features of the labor market. Instead it seems to be the data variation that traces out the wage curve. The variability of the unemployment rates, both over time and across industries, is clearly shifting the bargained wage.

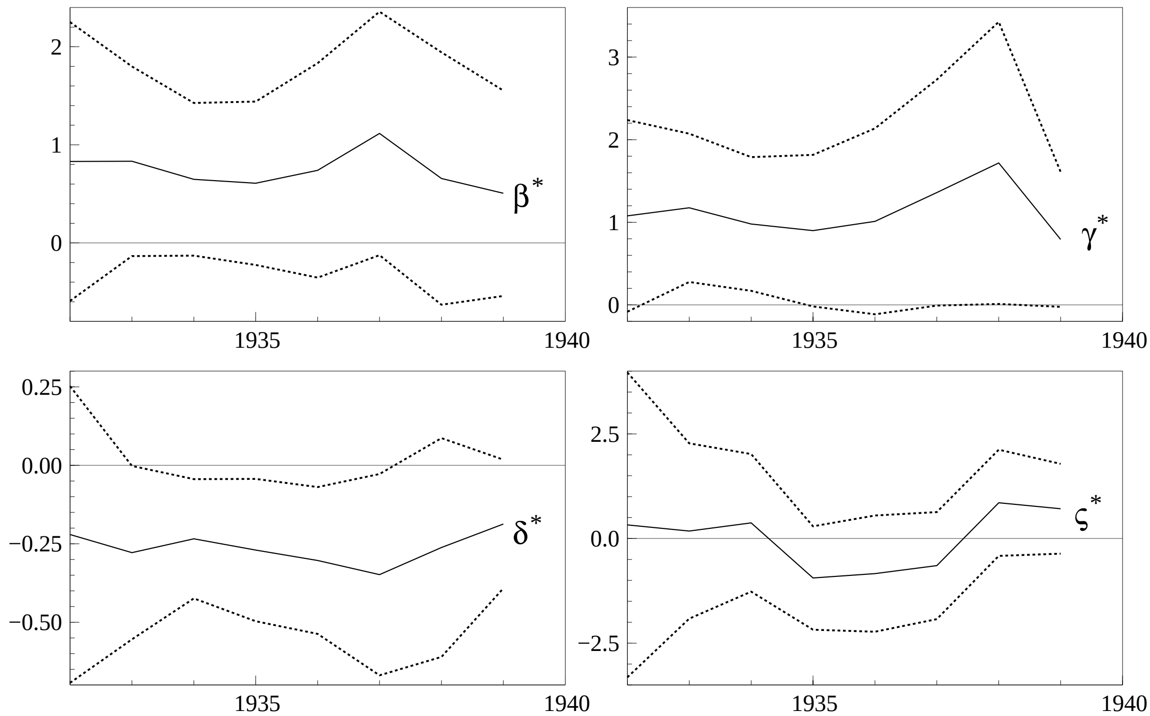


Figure 4: Recursive estimates of the steady-state parameters ± two standard errors.

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 steady-state solution recursively, also see Johansen (1999). Figure 4 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. Whether this effect is due to lack of 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 another important issue when analyzing the interwar labor market. Blanchflower and Oswald (1994) in particular 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 this hypothesis on the basis of our model. The test of the joint hypothesis

Table 3: Wage equations, GMM system estimates

| Dep. var: $\Delta w_{it}$ | (1) | (2) |
|---|---|---|
| $\Delta p_{it}$ | $-0.10$ (0.08) | – |
| $\Delta q_{it}$ | 0.12 (0.08) | 0.19 (0.05) |
| $\Delta u_{it}$ | 0.007 (0.008) | – |
| $\Delta pc_t$ | 0.83 (0.008) | 0.62 (0.07) |
| $(w - w^*)_{i(t-1)}$ | $-0.18$ (0.03) | $-0.27$ (0.04) |
| Diagnostics | | |
| Sargan: $\chi^2$ () | 53.39 (110) | 54.09 (50) |
| $AR$ (1) | $-4.01^{**}$ | $-3.74^{**}$ |
| $AR$ (2) | $-0.07$ | $-0.18$ |

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.

Finally we impose the steady-state solution

$$w_{it} = 0.5p_{it} + q_{it} - 0.1u_t + 0.5pc_t \tag{5}$$

$$\chi^2(4) = 1.81[0.77].$$

The associated p-value in brackets suggests that this empirical representation of (4) cannot be rejected. It is therefore imposed when we next turn to estimating the dynamic specification in the error correction form given by (3).[9]

### 4.3 The dynamic model

We have found a theoretically plausible and empirically sound wage equations for the interwar period in Norway, rejecting the hypothesis of Blanchard and Katz (1999) that the existence of a wage curve is dependent upon the presence of modern 'European' type labor market settings. 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

[9]Note that solving for the NAIRU is not possible without further identifying restrictions—see Bårdsen and Nymoen (2003) for the details.

of the Norwegian economy during the postwar era. The relevant evidence is reported in Table 3. Column (1) contains the general model reparameterized in error 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 0.31. This is of course in accordance with the corresponding results in Table 1. The final model is reported in column (2).[10] There is substantial nominal rigidity, as measured by the ECM 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 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 behavior 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 5, 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 behavior during the sample period.

# 5 Conclusions

Our empirical analysis does not lend any support to the hypothesis of Blanchard and Katz (1999)—that the presence of a wage curve is due to relatively strong worker rights and alternative labor markets. 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 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 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.

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

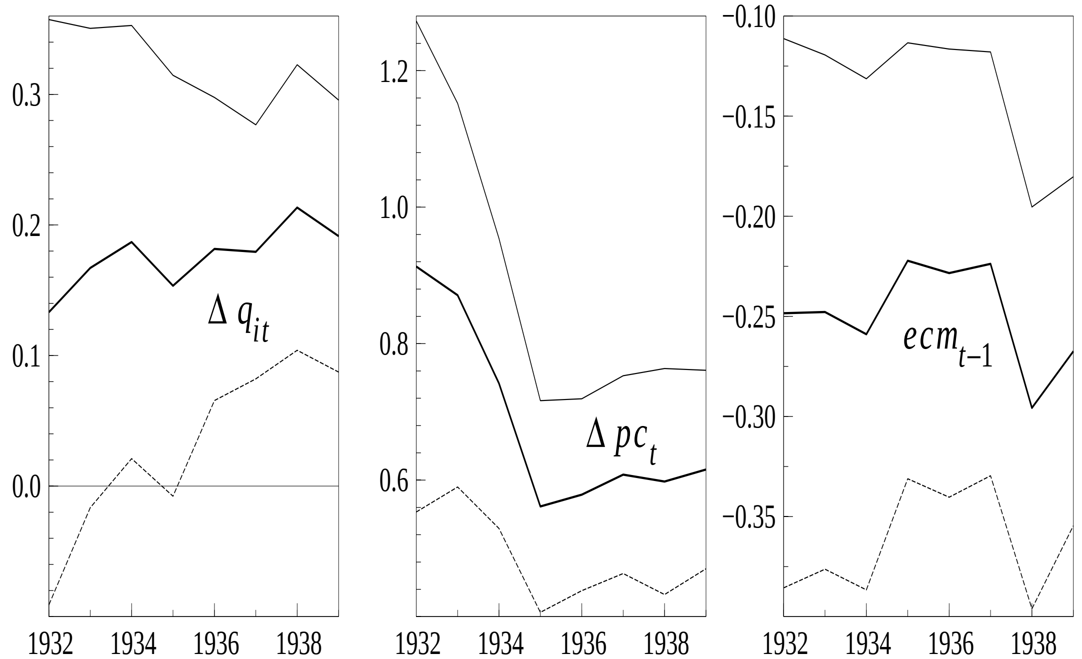


Figure 5: Recursive estimates of the model parameters ± two standard errors.

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# A A simulation experiment of the properties of the estimators

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

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

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 $\alpha$ than the differenced GMM estimator (*Diff*) when $\alpha$ is close to unity. It was also noted that *Diff* underestimates $\alpha$, whereas *Sys* produces an overestimate. While Doornik et al. (1999) keep $\beta$ fixed at unity, we now proceed to keep $\alpha$ fixed at 0.9, and vary $\beta$. 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* | $\Delta$ | $\Delta\mathbf{y}_{i,-1}, \Delta\mathbf{x}_i, \mathbf{1}$ | $\mathrm{diag}(y_{i,t-3} y_{i,t-2}), \Delta\mathbf{x}_i, \mathbf{1}$ | 1-step |
| *Sys* | $\Delta$ | $\Delta\mathbf{y}_{i,-1}, \Delta\mathbf{x}_i$ | $\mathrm{diag}(y_{i,t-3} y_{i,t-2}), \Delta\mathbf{x}_i$ | 1-step |
| | levels: | $\mathbf{y}_{i,-1}, \mathbf{x}_i, \mathbf{1}$ | $\mathrm{diag}(\Delta y_{i,t-2}), \mathbf{x}_i, \mathbf{1}$ | |

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 $t = 0,\ldots,4$. 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 6. 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).

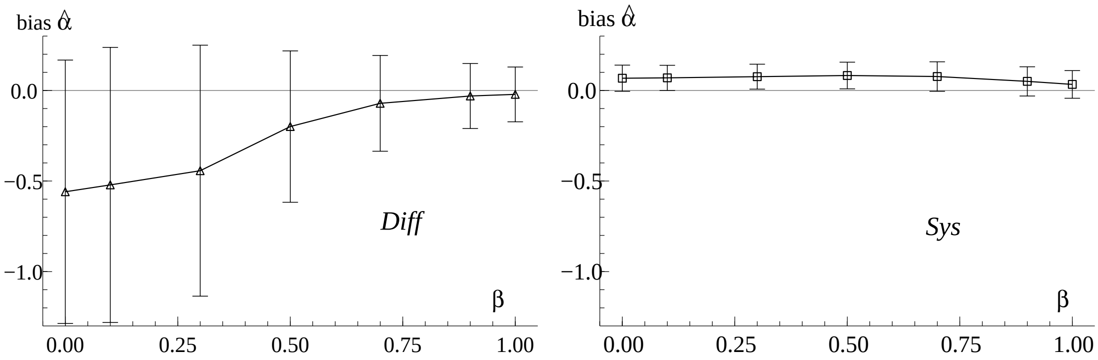


Figure 6: Mean bias of $\hat{\alpha}$, $M = 1000, \alpha = 0.9, \rho = 0.8, \sigma_e^2 = 0.9$; bars are twice the MCSD; $\beta = 0, 0.1, 0.3, 0.5, 0.7, 0.9, 1$.

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 $\beta$; for example when $\beta = 0.3$, the mean estimated $\hat{\alpha}$ is close to 0.5. *Sys* again overestimates $\alpha$, 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 $\beta$ in Figure 6.

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*.

# B 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 = labor 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 labor exchanges, classified by industry groups.

**PC = retail price index**

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