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Wage and Profitability: Norwegian Manufacturing 1967-1998

Abstract:
Economic theories of imperfectly competitive labour markets predict that wages are linked to profits. In spite of this, profit variables are not explicitly specified in empirical models of wage formation that otherwise are appealing. Does this mean that theory overplays the role of profitability in wage formation? The answer is probably not: Using Norwegian wage formation as an example, we model the dereminants of profitability within a vector autoregressive model and show that existing wage equations that have been successful empirically in fact contain a close linkage between wage setting and profits.

Keywords: Cointegration, incomes policy, profitability, rate-of-return, vector autoregressive model, wage formation.

JEL classification: E2, E64, J31, J51.

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1 Introduction

The relationship between industrial profitability and wage formation has important implications for economic performance in an open economy. Excessive wage-claims relative to profitability have adverse effects on competitiveness, employment and, over a longer horizon, also on capital formation. The coordination of the Norwegian wage settlements, has probably helped produce a shared recognition of these mechanisms by both sides of the bargain, see Dølvik et al. (1997) for a comprehensive exposition of the institutional set-up. Therefore, high nominal wage claims seldom occur simultaneously with low industrial prosperity. Conversely, high ability to pay among firms is normally a good predictor of increased wage claims winning through in the bargaining process. The 'main-course' theory of wage development in Aukrust (1977), provided an early attempt to formalize these mechanisms in economic terms.

Although the importance of wage-setting institutions in individual economies should not be underestimated, the profit-to-wage linkage is probably sustained by economic motives and rules of behavior that are more general in nature. Indeed, the basic implication has a long history in economic research that assume imperfect competition in labour markets, see e.g. Carruth and Oswald (1989) for an overview of the early literature, and recent contributions by Solow (1990) and Hahn (1997). Moene (1988) shows that modern theories of wage bargaining imply that there is an element of "profit sharing" in union based wage setting.

The empirical confirmation, or otherwise, of the importance of profitability effects in wage setting is important for agencies that monitor the economy and those who make policy considerations. For example, if the economy is on the recovery after a slump and firm profitability has increased, wage growth may increase well ahead of any sign of a fall in unemployment. More fundamentally, if wages equilibrium corrects with respect to profitability, there may not be a unique supply side determined equilibrium of the economy, see Kolsrud and Nymoen (1998).

Against this background, it is surprising to observe that explicit variables of firm profitability are seldom included in empirical wage equations. Rowlatt (1987) and Carruth and Oswald (1989, Chapter 7) are exceptions from this rule, and they both find significant profit-to-earnings effects in Britain. Another observation is that when researchers include factors that affect profitability (e.g. product price and productivity indices), the overall implication for the importance of the profitability-to-wage hypothesis is often not drawn, the multi-country study in OECD (1997, Chapter 1C) is one recent example. However, following Blanchard and Summers (1986), several studies have subsumed output prices and productivity under the heading of "insider variables", see Layard et al. (1991, Chapter 4). Generally, "inside factors" are found to influence wages, but their importance varies across countries. Layard et al. (1991) point to the Scandinavian countries as examples of economies were insider forces are likely to be weak, because of the centralized wage-setting institutions of these economies.[1]

In this paper we review a decade's work on empirical modelling of Norwegian manufacturing wages with the aim of elucidating the relationship between profitability and wage setting. In section 2 we examine how the link between profitability and wages is represented in existing empirical models. In section 3 we extend the data used by Nymoen (1989a) by 10 years of quarterly observations. Building on the wage equation in Nymoen (1989a), and on the improvement of that model in Johansen

[1] See Layard et al. (1991), p. 188 and 212.

(1995a), we formulate a model that has stable parameters over the extended sample. The model implies that the wage-share depends on the rate of unemployment, but in a highly non-linear manner. With this model as our reference, we introduce the rate-of-return on capital as an explicit measure of profitability, and test its impact on model specification and on residual fit. Section 4 shows the results of joint modelling of the determinants of profitability. Section 5 summarizes the empirical findings and discusses implications and generalizations.

2 Existing studies

Previous studies of Norwegian manufacturing wage formation include Nymoen (1989a), Calmfors and Nymoen (1990), Johansen (1995a) and Rødseth and Nymoen (1999). Nymoen (1989a) uses quarterly data from 1966.1-1987.4. The three other studies are annual with samples that end in 1987, 1990 and 1994 respectively. The following equation summarizes the common features of the existing studies

$$\begin{aligned}(1)\qquad \Delta wc_t &= \gamma_0 - \gamma_1(wc - p - y)_{t-1} + \gamma_2(pc - p - at + pt)_{t-1} \\ &\quad -\gamma_3 f(U_t) + \gamma_4\Delta p_t + \gamma_5\Delta y_t + \gamma_6\Delta pc_t + \gamma_z z_t + \varepsilon_t.\end{aligned}$$

wc denotes hourly wage cost; p the producer price index; pc the consumer price index; y labour productivity. These variables are measured in logarithmic scale. U is the rate of unemployment. The degree of concavity of the $f-$function is a separate issue: Nymoen (1989a) and Rødseth and Nymoen (1999) use a logarithmic transform (γ_3 is an elasticity), while Johansen (1995a) prefers a more concave functional form:

the square inverse $f(U) = U^{-2}$.[2]

pt and at are tax-rate variables defined as $pt = \ln(1+$ *tax rate on the use of labor)* and $at = \ln(1-$ *average income tax rate).* Finally, z_t is (a vector of) other explanatory variables, e.g. dummies for incomes policies, and ε_t is an error term. With the exception of γ_3, the unknown coefficients are non-negative. With a logarithmic f-function, $\gamma_3 \geq 0$, while for the square inverse transform, $\gamma_3 \leq 0$. The subscript t is for time period. The simple dynamic structure of (1) is most representative for the annual studies.

Note that $(wc - p - y)$ is the log of the *wage share*, and that $(pc - p - at + pt)$ is the log of the *wedge* between producer real wages and consumer real wages. The two zero restrictions $\gamma_1 = \gamma_2 = 0$ imply that the model reduces to an open economy Phillips-curve equation. If on the other hand, $\gamma_1 > 0$ (1) implies a steady state equation for the wage-share

$$(2) \qquad wc - p - y = \frac{\gamma_0}{\gamma_1} + \frac{\gamma_2}{\gamma_1}(pc - p - at + pt) - \frac{\gamma_3}{\gamma_1} f(U) + \frac{(\gamma_4 + \gamma_6 - 1)}{\gamma_1} \Delta pi + \frac{(\gamma_5 - 1)}{\gamma_1} \tau$$

where we have suppressed the error term, and the steady state is defined by

i. $(\Delta(wc - p - y)_t \mid_{\Delta U = 0}) = 0$,

ii. $\Delta p_t = \Delta pc_t = \Delta pi$,

iii. $\Delta y_t = \tau$,

The wage share is constant for a given rate of unemployment, producer and consumer

[2] See also Johansen (1997).

prices grow in accordance with the international rate of inflation Δpi, productivity growth is constant at a rate τ. In addition, (2) assumes that the tax rates are constant, and that $z_t = 0$ in the steady state. Equation (2) can be simplified by imposing two testable assumptions, namely

- *No wedge*, $\gamma_2 = 0$, and
- *Dynamic homogeneity*, $\gamma_4 + \gamma_6 = 1$.

Since

$$wc - p - y = \ln(1 - \text{capital share}),$$

the *No wedge* restriction $\gamma_2 = 0$ is seen to turn (2) into a long-run profitability equation. In this case, the equilibrium correction form of the dynamic model (1) becomes

$$\begin{aligned} \Delta wc_t &= -\gamma_1\{(wc - p - y)_{t-1} - wshare_{t-1}\} \\ &\quad + \gamma_4 \Delta p_t + \gamma_5 \Delta y_t + \gamma_6 \Delta pc_t + \gamma_z z_t + \varepsilon_t, \end{aligned} \tag{3}$$

where

$$wshare_t = \frac{\gamma_0}{\gamma_1} - \frac{\gamma_3}{\gamma_1} f(U_t). \tag{4}$$

If *Dynamic homogeneity* is jointly acceptable with the no-wedge restriction, long-run profitability in (2) is unaffected by the rate of inflation, and the equilibrium correction equation (3) can be written in terms of real-wage growth, $\Delta wc_t - \Delta pc_t$

on the left hand side of the equation.

Both restrictions can be argued from theory: *Dynamic homogeneity* is plausible once we think of Δp_t and Δpc_t in (1) as *expected* price increases. The *No wedge* restriction can be shown to hold theoretically for reasonable specifications of the underlying utility functions of unions, see e.g. Rødseth (1999, Chapter 8.5).

Following Nymoen (1989a), the existing studies assume that wage, prices and productivity are integrated of degree one, $\mathsf{I}(1)$ in a common notation. Although the rate of unemployment is conceptually integrated of degree zero, $\mathsf{I}(0)$, with a bounded variance, the actual time series of the (transformed) rate of unemployment is heavily autocorrelated, which makes it behave as if it too was $\mathsf{I}(1)$. The reactions of previous researchers has been to treat the transformed rate of unemployment as $\mathsf{I}(1)$, and to test whether it cointegrates with the other variables in the wage-equation.

Empirically, Nymoen (1989a) found that a wedge effect, $\gamma_2 > 0$ was necessary for cointegration. However, Johansen (1995a) shows that this result hinges on Nymoen's use of an import price index as the operational measure for the product price index p_t. When Johansen instead uses a deflator of factor income, the no-wedge restriction is accepted statistically. Rødseth and Nymoen (1999) measures the producer price by the value added deflator and obtains an empirical model that corroborates Johansen's finding: The wedge variable is statistically insignificant, not only in the equation for Norwegian manufacturing wages, but also in the similar models for the other Nordic countries. Dynamic homogeneity is not rejected in Rødseth and Nymoen (1999), but the earlier studies either do not test the hypothesis or it is rejected statistically. In Rødseth and Nymoen (1999) the dynamic homogeneity property is data acceptable irrespective of whether the current price increase terms in the equation are instrumented or not.

Existing studies basically use single equation estimation method, although contemporaneous variables are sometimes instrumented, as in the study on annual data by Rødseth and Nymoen (1999). In the quarterly model of Nymoen (1989a), the short run part of the model is dominated by lagged terms, so the need to instrument that part of the model did not arise. However, the use of ordinary least-squares may lead to (finite sample) bias in the estimation of the parameters of the cointegrating vector. Following, Johansen (1992) the efficiency of single equation estimation rests on the validity of weak exogeneity of the conditioning (levels) variables with respect to the parameters in the cointegrating vector. That being said, the previous studies show that the estimated parameters are recursively stable over the available samples, which is informal evidence that any remaining bias is of "second order importance".

In the empirical sections of this paper we first, in Section 3, apply the same methodology as in Nymoen (1989a), in order to focus on the impact of an explicit measure of profitability on the existing models. In section 4 we use the multivariate cointegration methods and test the exogeneity assumptions.

3 Single equation estimation of profitability effects

In this section we use an extension of the data in Nymoen (1989a) up to 1998q2 to elucidate two main issues. First, we investigate whether the wage-share formulation in (3), similar to Johansen's (1995a) annual model, is valid for the quarterly data set. Second, we test whether an explicit profitability measure contains extra predictive power for wage growth.

The profitability measure that we use is the gross operating surplus, per currency unit of capital, RR :

$$(5) \qquad RR_t = \frac{P_t Q_t - WC_t HW_t}{PJ_t\ K_t} = \frac{P_t\ Y_t}{PJ_t K_t / HW_t}[1 - \frac{WC_t}{P_t Y_t}],$$

where Q_t is value added (at constant prices), P_t is the producer price index, Y_t is labour productivity, HW_t is the total number of hours worked (so $WC_t HW_t$ is the wage bill), PJ_t is the price index of gross investments and K_t is the capital stock.

We address both issues by first setting out a model that encompass both a wage-share and a rate-of-return specification. For that purpose we use a log-linearization of the rate-of-return:

$$(6) \qquad rr_t \approx a_0 - a_1[wc - (p + y)]_t + [wc - (pj + k - hw)]_t,$$

where a_1 depends on the sample mean ($a_1 = (\overline{P\,Y})/(\overline{P\,Y} - \overline{WC}) > 1$). For the period 1968q1–1998q2 $a_1 = 2.77$. This linearization enables us to embed both the wage-share and the rate-of-return specification in a wage equation that retains linearity in the parameters

$$(7) \qquad \begin{aligned} \Delta wc_t &= \gamma_0 - \gamma_1(wc - (p + y))_{t-1} + \gamma_{rr}(wc - (pj + k - hw))_{t-1} \\ &\quad + \gamma_3 f(U)_t + \gamma_4 \Delta p_t + \gamma_5 \Delta y_t + \gamma_6 \Delta pc_t + \gamma_z z_t + \varepsilon_t. \end{aligned}$$

$\gamma_{rr} = 0$ yields the wage-share model (3). On the other hand, $\gamma_{rr} = \gamma_1 / a_1 = \gamma_1 / 2.77$ is consistent with a rate-of- return specification.

Table 1 shows an empirical counterpart to (7) estimated on a sample 1968q1-1998q2. The dynamic specification of the model closely resembles the equations

Table 1: A model that encompasses wage-share and profitability specifications. Standard errors in parentheses.

$$
\begin{array}{llllllll}
\Delta_4 wc_t & = & + & \underset{(0.035)}{0.059} & & - & \underset{(0.006)}{0.046} & DUM_t \\
& & + & \underset{(0.037)}{0.824} & \Delta_3 wc_{t-1} & + & \underset{(0.076)}{0.802} & \Delta_2 pc_{t-1} \\
& & - & \underset{(0.015)}{0.060} & (wc-(p+y))_{t-4} & + & \underset{(0.014)}{0.024} & (wc-(pj+k-hw))_{t-4} \\
& & + & \underset{(0.0010)}{0.0046} & U_{t-3}^{-2} & + & \underset{(0.019)}{0.054} & \Delta_3 y_t \\
& & - & 1 & \Delta nh_t & + & \underset{(0.152)}{0.367} & \Delta PT \\
& & + & \underset{(0.0011)}{0.0034} & \Delta U_{t-1}^{-2} & & &
\end{array}
$$

Sample 1968.1-1998.2

R^2	$= 0.940$
σ %	$= 1.19$

Misspesification tests, p-values in brackets:

$F_{ar,1-5}(5, 107)$	$= 0.35$ [0.88]
$F_{arch,1-4}(4, 104)$	$= 1.06$ [0.38]
$\chi^2_{nd}(2)$	$= 2.15$ [0.34]
$F_{het}(18, 93)$	$= 1.44$ [0.13]
$F_{reset}(1, 111)$	$= 0.06$ [0.80]

in Nymoen (1989a) and Nymoen (1989b), although the sample ended in 1987q4 in those two studies. The equation explains the 4-quarter growth rate in nominal hourly wage costs, $\Delta_4 wc_t$. But we note that, because of the $\Delta_3 wc_{t-1}$ term on the right hand side, the equation can be re-arranged with the quarterly growth rate, Δwc_t on the left hand side. Hence, the role of the $\Delta_3 wc_{t-1}$ term is to capture the negative autocorrelation in Δwc_t, which reflects the pattern of the centralized wage settlements (2. quarter each year). It is also consistent with the earlier quarterly studies that there is no within quarter effect of CPI-inflation, and that the first lags of Δpc have huge effects as captured by the 0.8 elasticity of the $\Delta_2 p_{t-1}$ variable. Dynamic homogeneity can be imposed on the specification in Table 1, since the restriction on $\Delta_3 wc_{t-1}$ and $\Delta_2 p_{t-1}$ is not rejected statistically (test whether $3 * 0.824 + 2 * 0.802 = 4.076$ differ significantly from 4): the Wald-statistic is $\chi^2(1) = 0.30[0.58]$. The wedge-coefficient has been restricted to zero (the test for adding a wedge term yields $F(1,111) = 0.31[0.58]$).

Productivity growth is also averaged, in the form of $\Delta_3 y_t$, perhaps to extract more permanent effects. The tendency to provide full short-run wage compensation for the reductions in normal hours (Δnh_t) over the sample, is captured by the unity coefficient of (Δnh_t). ΔPT is the change in social cost on the use of labor, mainly taxes. DUM_t is a dummy for the laws imposed on wage and price growth in 1979 and 1988–89. Finally, there is squared inverse transformation of the rate of unemployment, (U_{t-3}^{-2}) and a difference of the same variable (ΔU_{t-1}^{-2}), consistent with Johansen's (1995a) preferred equation on annual data.

The "t-value" of the coefficient of $(wc - (pj + k - hw))_{t-4}$ is 1.7. Formally, the F-statistic for the wage-share restriction $\gamma_{rr} = 0$ is $F(1,112) = 2.84.[0.09]$. The F-test for the rate-of-return restriction $\gamma_{rr} = \gamma_1/2.77$ is even less significant, $F(1,112) = 0.03[0.87]$. The estimated residual standard errors are 1.20% for the wage-share specification and 1.18% for the rate-of return specification. The residual diagnostic test statistics at the end of the table include tests for 5. order autocorrelated residuals ($F_{ar,1-5}$), for normally distributed (χ^2_{nd}), autocorrelated squared residuals ($F_{arch,1-4}$), for heteroscedasticity due to squares of the regressors (F_{het}) and finally the regression specification test (F_{reset}). The statistics are explained in Doornik and Hendry (1999). [3] They give no indication of residual misspecification of the model in Table 1, and they change little when the wage-share and rate-of-return restrictions are imposed.

Figure 1 gives visual evidence of the stability of the two competing simplifi-

[3] As indicated in the Table, the normality test is a Chi-square tests, the other tests are F-distributed under their respective null hypotheses.

cations of the equation in Table 1. The first row shows the sequence of 1-step OLS residuals with ± 2 residual standard errors ($\pm 2\sigma$ in the graph), and the recursively estimated coefficients (γ in the graphs) of the lagged wage-share together with ± 2 coefficient standard errors ($\pm 2se$ in the graph). The second row shows the same graphs for the rate-of-return specification.

In sum, the wage-share formulation that Johansen (1995a) showed was an encompassing model on annual data ending in 1990, is a data congruent model also on our quarterly and extended data set. The model shows a high degree of stability over 30 quarters outside Johansen sample period. In this (out of sample) period unemployment reached a post war all time high in 1992-1993. Foreign currency have been in turmoil, first in late 1992 and then again in January 1997. Incomes policy has been pursued with renewed energy by the social democratic government from that came into power in late 1990. A rate-of return specification improves the fit compared to the wage-share specification, but with a conventional 5% significance level, the increased predictive power is not statistically significant.

So far we have built on the existing single equation studies of manufacturing wages. The statistical testing rests on several assumptions that can be tested. For example, we have assumed that the log-linearization of the rate of return is $I(0)$, or alternatively, that the rate-of-return and the non-linear transform of the rate of unemployment are both $I(1)$ but cointegrated. Furthermore, there is the implicit assumption of exogeneity of prices, productivity and unemployment with respect to the parameters of the cointegration relationship. In the next section we address these issues in a joint model of wages and profitability.

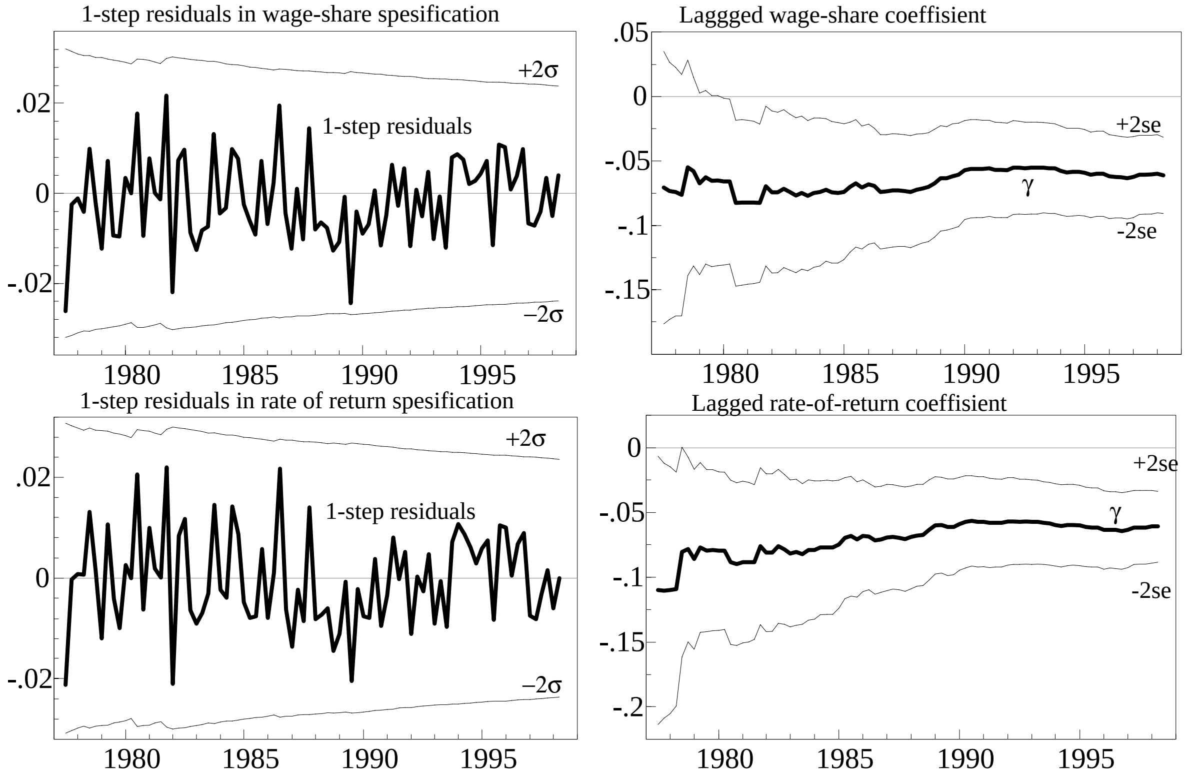

Figure 1: Recursive stability of wage-share and rate-of-return specifications of the equation in Table 1.

4 Modelling wages and profitability

We use a 4. order vector autoregressive model (VAR) in the three variables $\{wc_t, (p+y)_t, (pj+k-hw)_t\}$, i.e. wage costs, the value of fixed capital and value added, all measured per working hour. The rate of unemployment is included in the long-run part of the system as a non-modelled variable, i.e. U_{t-3}^{-2}. The other non-modelled variables are also taken form the model in Table 1: Δpc_{t-1}, Δpc_{t-2}, U_{t-3}^{-2}, ΔU_{t-1}^{-2}, Δnh_t, ΔPT_t and DUM_t.[4] In addition there are three new dummies, $i1984q1$, $i1986q3$ and $i1997q2$ that were included to mop up three large residuals in the equation for $(pj+k-hw)_t$.

Table 2 reports residual standard errors and test for misspecification for the

[4] The only variable in Table 1 that is omitted is Δy_t.

Table 2: Diagnostics for unrestricted 4. order conditional VAR for 1968.1-1998.2, p-values in brackets.

	wc	$p+y$	$pj+k-hw$	VAR
$\hat{\sigma}100$	1.14%	4.46%	2.94%	
$F_{ar,1-5}(5,90)$	1.97[0.09]	0.43[0.82]	0.77[0.58]	
$F_{arch,1-4}(4,87)$	1.56[0.19]	0.06[0.99]	0.61[0.65]	
$F_{het}(46,48)$	0.93[0.60]	1.05[0.44]	0.84[0.72]	
χ^2_{nd}	6.74[0.03]	0.06[0.97]	4.70[0.10]	
$F^v_{ar,1-5}(45,235)$				0.84[0.76]
$F^v_{het}(276,265)$				0.74[0.99]
$\chi^{2,v}_{nd}(6)$				10.9[0.09]
Trend, *constant* and 3 *centred seasonals* included.				

three equations and for the system. The single equation misspecification tests reveal no misspecification except rejection of normality with respect to the wage-costs equation at a 5% significance level. However, the corresponding vector test (in the VAR column) do not exhibit any non-normality, so the system is taken to be reasonably data congruent and forms the basis of the cointegrating analysis[5], see Doornik and Hendry (1996).

For the investigation of parameter constancy the 1-step residuals are shown in Figure 2, The 1-step residuals all lie within their two respective standard error bands. Thus, the three dimensional VAR(4) is taken to be stable.

Let β denote the matrix that contains the cointegrating vectors. If x_t is given by $x_t = [wc_t, (p+y)_t, (pj+k-hw)_t, U^{-2}_{t-3}]'$, cointegration entails that

$$\beta' x_t = \beta'[wc_t, (p+y)_t, (pj+k-hw)_t, U^{-2}_{t-3}]' \sim \mathsf{I}(0). \tag{8}$$

If the rank of β is one, there is a single cointegrating vector. In the case of multiple cointegrating vectors (the rank of β is 2 or larger), $\mathsf{I}(0)$ on the right hand side in (8)

[5] We also note that the validity of the Johansen-procedure does not strictly depend on the normality assuption (Lütkepohl (1991))

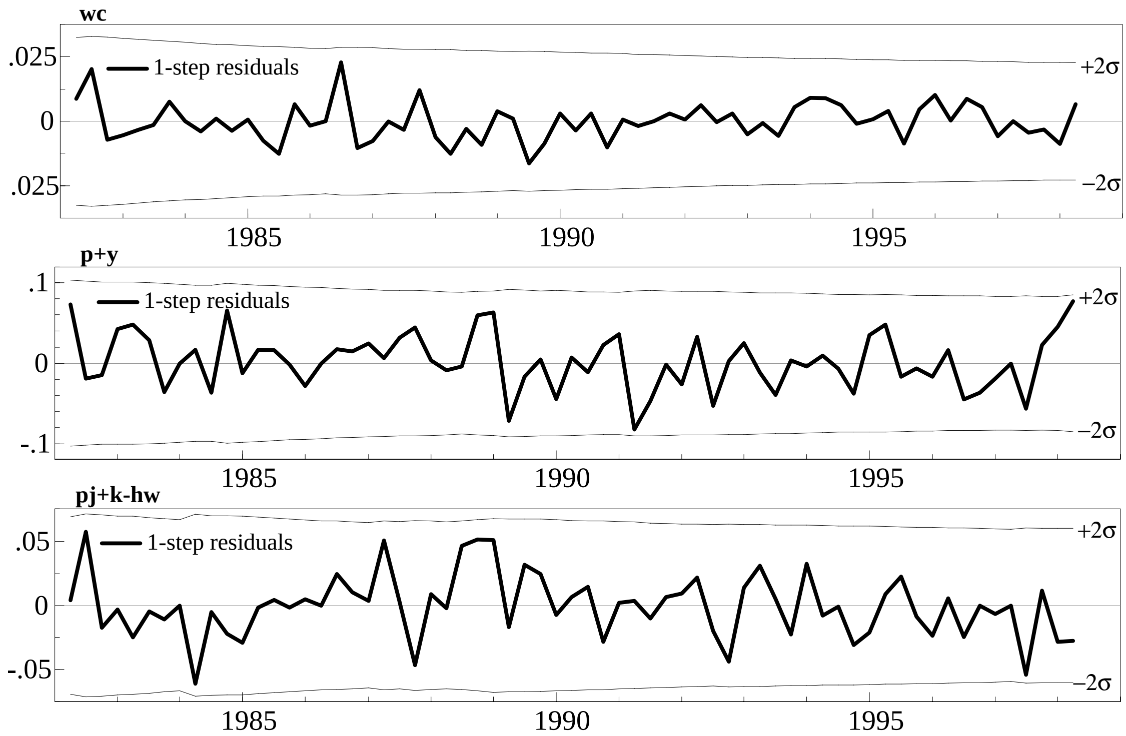

Figure 2: 1-step residuals for the VAR, with ±2 residual standard errors.

denotes a vector with the separate stationary linear combinations. Table 3 contains the cointegration analysis based on this system, see Johansen (1988), Johansen and Juselius (1990) and Johansen (1995b).

To robustify the inference against the consequences of including the rate of unemployment as a conditioning variable we also add the linear trend restricted, following the analysis in Harbo et al. (1998). The reported 5% critical values are taken from their Table 2. Nevertheless, since there are a number of non modelled differenced variables in the VAR, the formal significance of the Trace test only offer a guideline. That being said, the results seem to corroborate two cointegrating vectors (rank (r) is 2). Figure 3 shows the recursive stability of the two largest eigenvalues (λ_1 and λ_2). The eigenvalues are considered to be stable. Accepting $r = 2$, we next tested if the deterministic trend can be omitted from the cointegrating vectors.

Table 3: Cointegration tests (degrees of freedom corrected) and Harbo et. al. (1998) asymptotic critical values in parentheses. r denotes the rank of the cointegration matrix β.

Eigenvalues:	$\lambda_1 = 0.261$, $\lambda_2 = 0.228$, $\lambda_3 = 0.048$	
Trace eigenvalue test		
null	alternative	
$r = 0$	$r \geq 1$	67.2(49.6)
$r \leq 1$	$r \geq 2$	33.9(30.5)
$r \leq 2$	$r \geq 3$	5.43(15.2)
U^{-2}_{t-3} is weakly exogenous. Restricted *Trend* included.		

Table 4: The two first eigenvectors with corresponding equilibirum correction coefficients, standard error in parentheses.

$$\begin{bmatrix} \widehat{\alpha} & 1 & 2 \\ \Delta wc_t & \underset{(0.010)}{-0.055} & \underset{(0.032)}{0.012} \\ \Delta(p+y)_t & \underset{(0.036)}{0.007} & \underset{(0.120)}{0.305} \\ \Delta(pj+k-hw)_t & \underset{(0.025)}{-0.032} & \underset{(0.085)}{0.358} \end{bmatrix},$$

$$\begin{bmatrix} \widehat{\beta}' & wc & (p+y) & (pj+k-hw)_t & U^{-2}_{t-3} \\ 1 & 1 & -1.319 & 0.355 & -0.054 \\ 2 & 1 & -0.434 & -0.56 & 0.008 \end{bmatrix}$$

Remember that the trend was introduced for inference purposes, it is difficult to see the economic interpretation of having it appearing in the cointegrating vectors. In the present context therefore, a 5% significance level seems high and a test statistic of $\chi^2(2) = 6.2795[0.04]$ should not retain us from dropping the trend from the $\mathsf{I}(0)$ system.

Table 4 shows the estimation results of the two cointegrating vectors (β) and the corresponding equilibrium correction coefficients (α matrix in a common notation).

4.1 The wage-share model

The wage-share hypothesis does not identify the β'-matrix. However, if we impose the exogeneity restrictions implied by the existing single equation studies (only

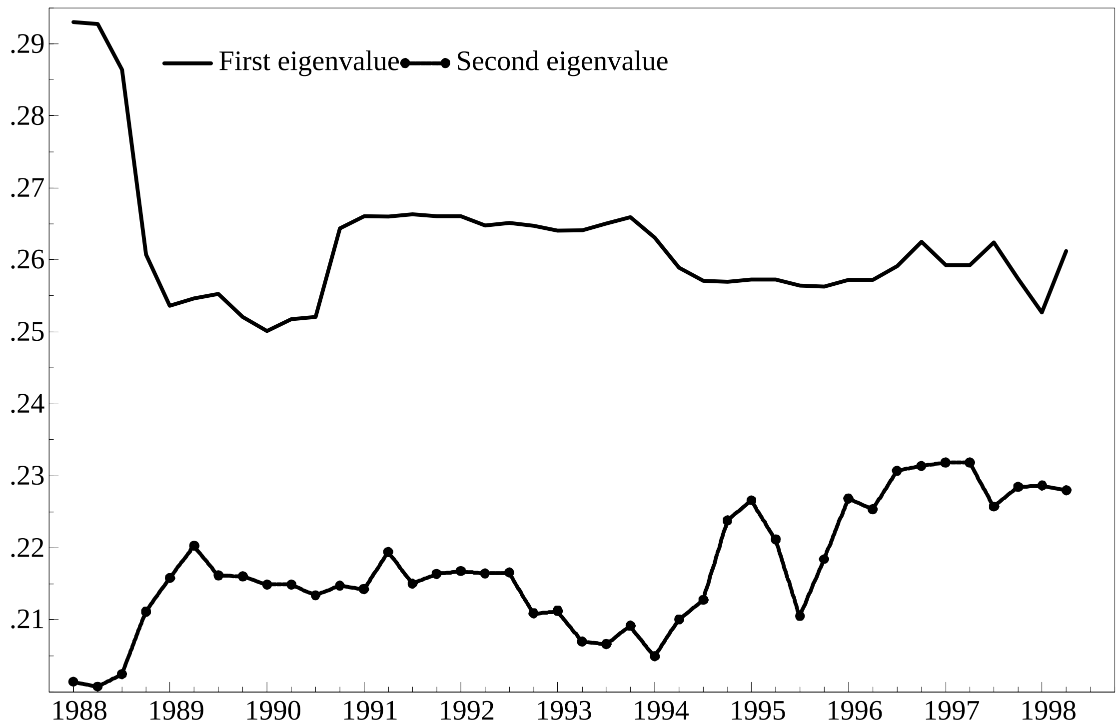

Figure 3: Recursive eigenvalues of the system in Table 3.

wages are equilibrium correcting), the system is identified. Table 5 shows estimation results for a set of restrictions that define a separating system: the first (wage-share) equilibrium correction mechanism only enters in the Δwc_t equation, the second vector only enters in the equation for $\Delta(pj + k - hw)_t$. Value added per hours worked $((p + y)_t)$ is exogenous with respect to both vectors. The joint test statistic is $\chi^2(6) = 11.144[0.084]$. Figure 4 shows that the test is insignificant for every sample size from 1983q1 to 1998q2. The single unrestricted coefficient in the β' matrix, that of U_{t-3}^{-2}, displays a noticeable drift at the end of the 1990s. Although the change in coefficient value is not statistically significant, the numerical instability indicate that Johansen's squared inverse can be disputed on this data set.

The restrictions on the $\alpha-$ matrix implies weak exogeneity of prices and productivity, and hence single equation estimation, is efficient , see Johansen (1992)

Table 5: Restricted cointegration analysis: Wage share specification $\chi^2(6) = 11.14[0.08]$, standard error in parentheses.

$$\begin{bmatrix} \widehat{\alpha} & 1 & 2 \\ \Delta wc_t & \underset{(0.013)}{-0.073} & 0 \\ \Delta(p+y)_t & 0 & 0 \\ \Delta(pj+k-hw)_t & 0 & \underset{(0.05)}{0.22} \end{bmatrix},$$

$$\begin{bmatrix} \widehat{\beta}' & wc_{t-1} & (p+y)_{t-1} & (pj+k-hw)_{t-1} & U_{t-3}^{-2} \\ 1 & 1 & 1 & 0 & \underset{(0.013)}{-0.053} \\ 2 & -1.38 & 0.38 & 1 & 0 \end{bmatrix}$$

and Banerjee et al. (1993). A model of wage growth consistent with the wage-share identification of the long-run part can thus be obtained from Table 1, but with $\gamma_{rr} = 0$ and the first vector in Table 5 imposed prior to estimation.

4.2 The rate-of-return model

In order to retain the rate-of return specification from the multivariate setting, we first restrict the β' matrix in the following manner:

wc	$(p+y)$	$(pj+k-hw)_t$	U_{t-3}^{-2}
1	-1	0	no restriction
1	0	-1	no restriction

The first (restricted) vector is the wage-share, with a freely estimated coefficient of the transformed rate of unemployment, while the second vector does not have an immediate interpretation. These β-restrictions are not significant as $\chi^2(2) = 1.08[0.58]$. The second step is to restrict the two equilibrium correction coefficients of Δwc_t in line with the linearization of log rate-of-return in section 2, i.e. $\alpha_{11} = -2.77\alpha_{12}$. The α and β restrictions on the system are jointly insignificant, with $\chi^2(3) = 1.17[0.76]$,

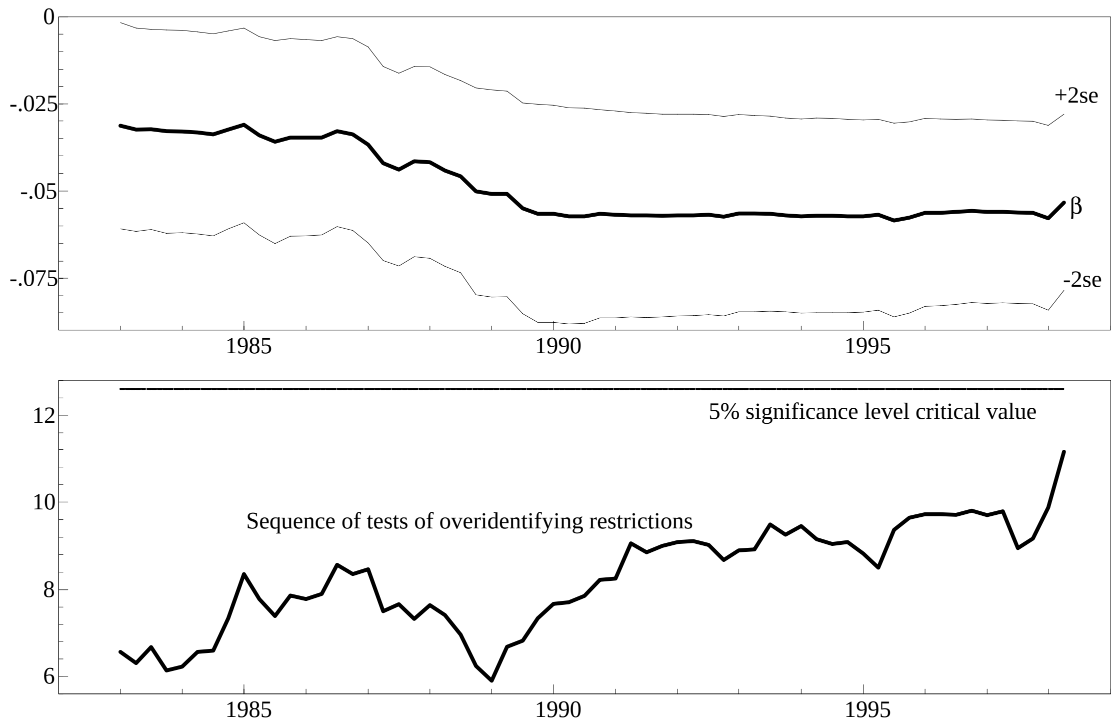

Figure 4: Wage-share model: Recursive stability of the estimated coefficient of U_{t-3}^{-2} in Table 5, and the sequence of $\chi^2(3)$ test statistics for the validity of the restricitons.

Table 6: Restricted cointegration analysis: Rate of return specification $\chi^2(3) = 1.17[0.76]$, standard error in parentheses.

$$\left[\begin{array}{ccc} \widehat{\alpha} & 1 & 2 \\ \Delta wc_t & -0.067 & \underset{(0.004)}{0.024} \\ \Delta(p+y)_t & \underset{(0.069)}{0.145} & \underset{(0.067)}{0.166} \\ \Delta(pj+k-hw)_t & \underset{(0.048)}{0.107} & \underset{(0.046)}{0.204} \end{array}\right],$$

$$\left[\begin{array}{ccccc} \widehat{\beta}' & wc_{t-1} & (p+y)_{t-1} & (pj+k-hw)_{t-1} & U_{t-3}^{-2} \\ 1 & 1 & -1 & 0 & \underset{(0.012)}{-0.042} \\ 2 & 1 & 0 & -1 & \underset{(0.014)}{0.043} \end{array}\right]$$

with restricted estimation results shown in Table 6. The standard errors below the α-estimates indicate that both $\Delta(p+y)_t$ and $\Delta(pj+k-hw)_t$ equilibrium correct to both vectors. The two vectors both include the square inverse of unemployment, however the net effect on Δwc_t is 0.0038 compared to 0.0046 in the single equation analysis in Table 1. The recursive stability of the coefficients are shown in Figure 5, together with the sequence of χ^2-tests for the 3 restrictions embodied in Table 6. The estimated coefficient is empirically stable and the restrictions are not rejected for any sample size between 1983q1 and 1998q2.

Note that the weak exogeneity assumption underlying the single equation estimation is not supported when we use the rate-of-return restrictions to identify the α and β matrices. This result is different from the wage-share case, where we saw that exogeneity was validated. The reason for this is that weak exogeneity is a model property and is relative to the parameters of interest and not a data property, see Ericsson (1992) and Ericsson et al. (1998).

The result that no variables are weakly exogenous invites further modelling of the complete 3-variable system. A *model* of the three variables in the system is shown in Table 7. It is constructed from a 3. order DVAR model for the three endogenous

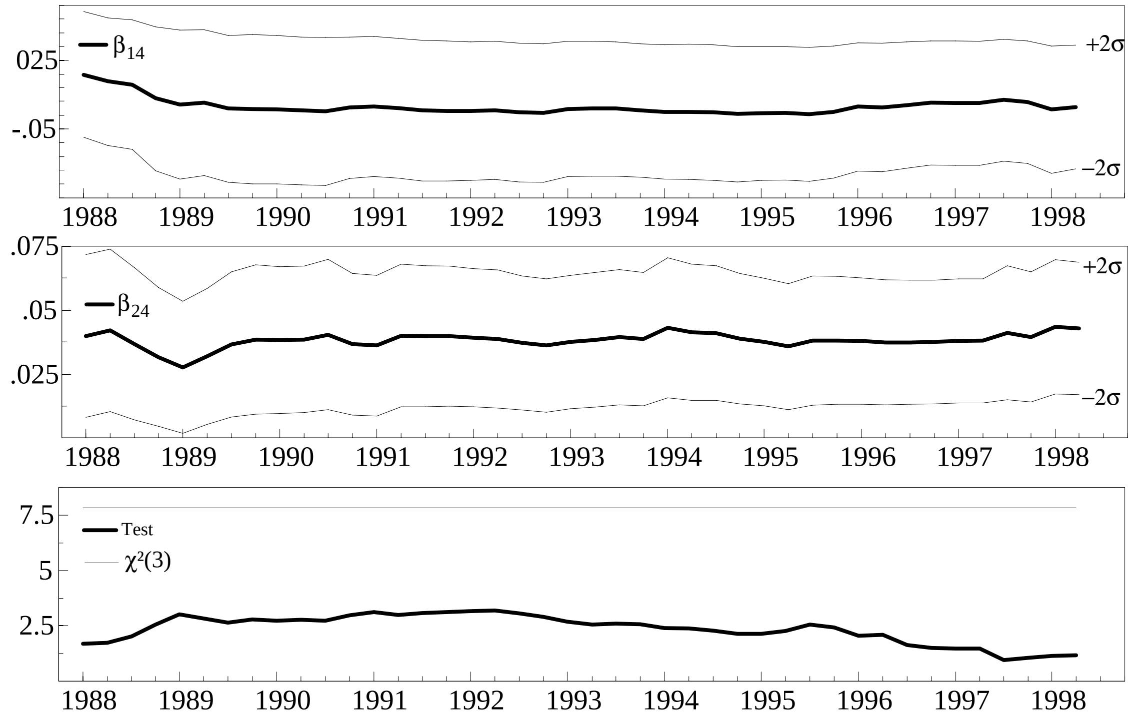

Figure 5: Recursive stability of the estimated coefficient of U_{t-3}^{-2} in Table 6 and the sequence of $\chi^2(3)$ test statistics for the validity of the restricitons.

variables Δwc_t, $\Delta(p+y)_t$ and $\Delta(pj+k-kw)_t$. The conditioning variables are the same as above, see Table 2, except that we now include the estimated equilibrium correction mechanism in the following form

$$EqCM\mathsf{RoR1}_t = wc_{t-3} - (p+y)_{t-3} - 0.042U_{t-2}^{-2},$$

$$EqCM\mathsf{RoR2}_t = wc_{t-3} - (pj+k-hw)_{t-3} + 0.043U_{t-2}^{-2}.$$

The model In Table 7 has been restricted to capture both the linearized rate of return and dynamic homogeneity in addition to other restrictions on the dynamic structure. The joint test of all the over identifying restrictions and of the restrictions yields $\chi^2(42) = 45.3[0.34]$ and $\chi^2(12) = 7.41[0.83]$ respectively. In the wage equation in Table 7 rr_{t-4} and U^{-2}_{t-3} are subsumed in $EqCM\mathsf{RoR1}_{t-1}$ and $EqCM\mathsf{RoR2}_{t-1}$[6]. In addition to the variables in the single equation in Table 1 lagged modelled variables are included in the wage equation in Table 7, i.e. $\Delta(pj+k-hw)_{t-2}$ and $\Delta_3(p+y)_{t-1}$. The additional variables reduce the estimated residual standard error from 1.18% in the single equation analysis to 1.13% in the multivariate. The coefficient estimates are otherwise almost unaffected by the change of estimation method from OLS to CFIML. Hence there is little loss of information when modelling wages in an one-equation setting compared to the multivariate analysis even when imposing the rate-of-return specification.

5 Conclusion

We have modelled a data set that extends the data used by Nymoen (1989a) by 42 quarters, from 1987q4 to 1998q2. Despite the ten years of new data the main structure of the short run part of the model is re-established on the new data. The cointegrating part of the equation is simplified to a wage-share formulation (no wedge), in accordance with the findings of Johansen (1995a), on annual data ending in 1990. Overall therefore, the investigation shows that an equation that links wage growth to CPI-inflation and to the lagged level of the wage-share is empirically stable in the 1990s. Dynamic homogeneity is data acceptable. The role of the wage-share in the equilibrium correction part of the model contradicts the widespread view that

[6] Recall from section 3 that since

$$-2.77EqCM\mathsf{RoR1}_t + EqCM\mathsf{RoR2}_t + a_0 \approx rr_{t-3} + 0.16U^{-2}_{t-2},$$

the error correction mechanism in the wage equation in Table 7 is:

$$-0.073EqCM\mathsf{RoR1}_{t-1} + 0.026EqCM\mathsf{RoR2}_{t-1} \approx 0.026rr_{t-4} + 0.0035U^{-2}_{t-3} - 0.026a_0.$$

Table 7: CFIML results for a model of the system, standard error in parentheses.

	Δwc_t	$\Delta(p+y)_t$	$\Delta(pj+k-hw)_t$
$\widehat{\sigma}100$	1.13%	4.31%	2.90%
$EqCM\mathsf{RoR1}_{t-1}$	-0.073 (...)	$+0.081$ (0.057)	$+0.076$ (0.037)
$EqCM\mathsf{RoR2}_{t-1}$	$+0.026$ (0.004)	$+0.160$ (0.049)	$+0.193$ (0.037)
$\Delta_3 wc_{t-1}$	-0.160 (0.036)		
$\Delta(pj+k-hw)_{t-1}$			-0.603 (0.071)
$\Delta(pj+k-hw)_{t-2}$	-0.048 (...)	-0.245 (0.054)	-0.452 (0.080)
$\Delta(pj+k-hw)_{t-3}$			-0.0276 (0.072)
$\Delta_3(p+y)_{t-1}$	$+0.034$ (0.014)	-0.272 (0.070)	
Δnh	-1 (...)		-1 (...)
$\Delta_2 pc_{t-1}$	$+0.713$ (0.051)		$+0.751$ (0.176)
ΔPT_t	$+0.562$ (0.140)		
ΔU^{-2}_{t-1}	$+0.0035$ (0.0010)	$+0.013$ (0.004)	
DUM	-0.044 (0.006)		
Not included in the table: a) Intercept in all three equations.			
b) Seasonals in columns 3 and 4.			
c) i1986q2 and i1984q1 in column 4, i1997q2 in columns 3 and 4.			
DVAR misspesification tests:			
$\mathrm{F}^v_{ar,1-5}(45,283)$	0.95[0.57]		
$\chi^{2,v}_{nd}(6)$	4.57[0.60]		
$\mathrm{F}^v_{het}(240,400)$	0.98[0.55]		

centralized wage bargaining imply weak effects of 'inside factors'.

Throughout we have adopted Johansen's "squared inverse" transform of the rate of unemployment. We suspect however that the exact form of the curvature of the "wage curve" is not a structural feature in the same sense as the CPI-inflation and lagged wage-share effects. For example, there is evidence from models of annual data that the logarithmic transform is also consistent with stability, see Rødseth and Nymoen (1999). That being said, it is interesting that a wage model with the highly convex unemployment effect survives in second half of the 1990s, since the implied "wage curve" is flat in that period. Hence the considerable moderation of nominal wage claims in these years, as noted by e.g. OECD (1997) is explained by lower profitability, by the centralized settlements in the 1988 and 1989 and by the relative stability of the nominal exchange rate. There is little separate wage reduction effects form the level of unemployment that was reached in the mid 1990s.

A main objective of the paper has been to investigate whether the lagged wage-share is a good representation of the link between profitability and wages. For that purpose we formulated a slightly more general model that used the rate of return on capital as an explicit measure of profitability. That led to the discovery that lagged wages relative to the capital-labour ratio ($(wc - (pj + k - hw))_{t-1}$) could be a "missing variable" in the existing wage-share models. However, although measures of fit favors the rate-of-return specification, the statistical discrimination between the contending representations of profitability effects was not very strong. Another aspect, brought out by the cointegration analysis, is that the "missing variable" is stationary, which helps explain the empirical success of wage-share models.

Finally, the system analysis showed that the rate-of-return specification implied that wages are fundamentally interlinked with value added productivity and the capital-labour ratio. Conversely, the wage-share identification of the cointegrating vectors, implied that a wage equation could be separated out as a valid conditional sub-system. This greatly simplifies forecasting and policy analysis.

A Data

w = log of average hourly wage in manufacturing, nominal kroner. Source: NHO [Confederation of Norwegian Business and Industry].

PT = obligated social costs, share in hourly wage. Source: Norges Bank [Central Bank of Norway].

$wc = w + \ln(1 + PT)$, log of wage costs per hour in manufacturing.

UR = number of persons officially registered without ordinary work. Source: Arbeidsdirektoratet [Dircetorate of Labour].

NW = number of persons employed. Source: Statistics Norway (Labor Force Sample Surveys).

$U = \frac{UR*100}{NW}$, log of registered unemployed in per cent of total employment.

pc = log of consumer price index. Source: Statistics Norway.

p = log GDP deflator for manufacturing. Source: Quarterly National Account, Statistics Norway.

y = log of average labor productivity in manufacturing. Source: Quarterly National Account, Statistics Norway.

k = log of fixed capital in manufacturing and construction. Source: Quarterly National Account, Statistics Norway.

hw = log of man-hour in manufacturing and construction. Source: Quarterly National Account, Statistics Norway.

pj = log of gross investments deflator in manufacturing and construction. Source: Quarterly National Account, Statistics Norway.

nh = log of normal weekly working hours. Source: Norges Bank [Central Bank of Norway].

DUM = 0.5 for the period 1980.2–1981.2, −1 in 1980.3, 1 in 1981.4, 0.5 in 1988.3, −0.5 in 1989.2, and −1 in 1990.3, otherwise 0.

$i1984q1$ = 1 in 1984.1, otherwise 0.

$i1986q2$ = 1 in 1986.2, otherwise 0.

$i1997q2$ = 1 in 1997.2, otherwise 0.

The data from the Quarterly National Account have a break in 1994.1 due to a revision of the national account. The new series are level adjusted.

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