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Employment behaviour in slack and tight labour markets

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Employment behaviour in slack and tight labour markets*

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Abstract

Empirical and theoretical studies suggest that employment behaviour varies with the state of the labour market since hiring and firings costs depend on the availability of labour. Extending earlier empirical work on this subject, we test for state dependence in employment adjustment and in the effects of forcing variables such as indicators of aggregate demand. We also test whether anticipated labour shortage leads to multiple equilibria in (un)employment. In the inquiry, we employ a linear vector equilibrium correction model (VEqCM) and two states Markov switching VEqCMs. The models are based on quarterly data for Norwegian industry employment and aggregate unemployment in the period 1974–96. We find clear evidence of state dependent adjustment and response to changes in forcing variables. Yet equilibrium solutions for the employment and unemployment appear invariant to cyclical and structural changes in the sample. Shifts in long run means of the variables are shown to depend on exogenous factors, product demand and unit labour costs in particular.

Keywords: *adjustment costs, asymmetric response, multiple equilibria, cointegration, Markov switching.*

JEL classification: *E24, J32, E32, C32, E51.*

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1 Introduction

Employment adjustment costs may explain a number of empirical regularities such as sluggish employment response to shocks, labour hoarding and asymmetric cycles in employment and GDP, see e.g., Hamermesh and Pfann (1996), Nickell (1995) and Rotemberg and Summers (1990). The adjustment costs affect not only the dynamics but may also induce lasting effects of shocks if they vary with the business cycle. Such costs are generally characterised as functions of labour shortage measures, e.g., the unemployment rate, see inter alia Ball and Cyr (1966), Hughes (1971), Peel and Walker (1978), Burgess (1988), (1992a) and (1992b). Presumably, labour shortages raise hiring costs by increasing search costs for suitable workers and makes employment adjust at a slower pace towards the desired level. Thus, conventional employment determinants such as real wages and product demand are believed to have weaker effects in a tight labour market than in a slack labour market. Further, anticipated future labour shortages may be a source of persistence and multiple equilibria in the overall unemployment rate, as implied by Moene et al. (1997).

However, existing empirical studies do not seem to present evidence of the joint occurrence of all these aspects of cycle dependent adjustment costs: cycle dependency of: (i) the adjustment process, (ii) effects of changes in forcing variables and (iii) multiple equilibria. The existing studies typically present evidence of (i) or (ii), but not of both (i) and (ii) occurring jointly, see e.g., Smyth (1984), Acemoglu and Scott (1994), Burgess (1988), (1992a) and (1992b). Furthermore, increasing number of studies report evidence of multiple unemployment equilibria, see Peel and Speight (1995), Skalin and Teräsvirta (1999), Bianchi and Zoega (1998) and Akram (1999). However, the evidence is based on univariate models, which do not identify the mechanisms that may have led to the appearance of multiple equilibria in a given sample; Multiple equilibria are implied by a range of mechanisms besides cyclical adjustment costs, see e.g., Cooper and John (1985), Manning (1990), Murphy et al. (1989), Pagano (1990) and Saint-Paul (1995).

We investigate the joint occurrence of the three aspects of adjustment costs using multivariate models of employment and unemployment that condition on relevant forcing variables. We also take into account the possibility of asymmetric response to positive and negative changes in forcing variables when testing for cycle dependent employment response. The possibility of sign dependent response arises if hiring costs are greater than firing costs, as observed by e.g., Hamermesh and Pfann (1996), Pfann and Verspagen (1989), Chang and Stefanou (1988) and Borrego (1998).

Econometrically, we build on Krolzig (2001) who employs a Markov regime switching vector equilibrium correcting model (MS-VEqCM) to allow for state dependence in the parameters, see Krolzig (1997) and Hamilton (1989). In his two-step approach, cointegration between US employment and output is established by following the procedure developed by Johansen (1988). Thereafter, the vector autoregressive model (i.e., VAR) is reformulated as a vector equilibrium correction model (VEqCM) and its parameters are allowed to shift by a first order Markov chain. We follow the same route to a large extent, but start out with a VAR for the Norwegian aggregate unemployment rate, industry employment and working hours, conditioning on a set of macroeconomic variables. This VAR is developed into an interpretable linear simultaneous equation model, hereafter referred to as a structural VEqCM, see Bårdsen and Fisher (1999) and Boswijk (1995). In the second step, we allow the parameters of the structural VEqCM to shift in the Markov way. Finally, within the derived Markov switching employment model, we allow for asymmetric response to over- and undermanning (relative to equilibrium employment) and to positive and negative shocks from forcing variables.

The rest of the paper is organised as follows: Section 2 sketches the way unemployment persistence and multiple equilibria may result from firms' efforts to cope with anticipated labour shortage, *friction*. Section 3 outlines the econometric framework while Section 4 presents the data set which consists of seasonally non-adjusted quarterly observations over the period 1974(1)-1996(4). Section 5 contains the structural VEqCM for industry employment, hours and aggregate unemployment. We test for friction induced multiple equilibria within the context of this model. Section 6 presents the results for the models with state dependent dynamics. The results clearly suggest that employment behaviour varies with a slack and tight labour market. Section 7 investigates whether these results are robust to an extension of the model, which allows for asymmetric response to positive and negative shocks from employment determinants. The appendix contains precise definitions of the variables, their source and tests of their time series properties.

2 Friction, persistence and multiple equilibria

A large number of studies assumes that present and anticipated labour shortage contributes to (un)employment persistence by raising employment adjustment costs, see e.g., Ball and Cyr (1966), Hughes (1971), Hazledine (1979), Smyth (1984), Peel and Walker (1978), Burgess (1988), (1992a) and (1992b). Moreover, Moene et al. (1997) suggest that anticipated labour shortage may even induce multiple (un)employment equilibria.

In order to synthesize these ideas, consider the labour demand function for a sector of the economy

$$\ln(N_t) = \Gamma_1 Z_t - f(U^e_{t+1}) + v_t, \qquad f' \geq 0, \tag{1}$$

where N_t is sectoral labour demand and v_t is a disturbance term. Z_t denotes a vector of conventional explanatory variables such as real wages and aggregate demand indicators while the function $f(U^e_{t+1})$ captures the idea that firms might be reacting directly to the anticipated future labour shortages indicated by the expected overall unemployment in period $t+1$: U^e_{t+1}. For example, high U^e_{t+1} presumably goes with low incentives to hoard labour. Following Moene et al. (1997) we refer to this direct effect of the aggregate rate of unemployment on sectoral employment as *friction*. Sectoral employment in this study is industry employment (i.e., in manufacturing and construction), which comprises 25% of all civilian employment in Norway.

In order to establish the aggregate consequences of a relationship like (1), we express the unemployment rate as

$$U = \ln(N^S) - \omega_1 \ln(N_t) - \omega_2 \ln(N^{rest}) + \varepsilon_t, \;\; \omega_1 + \omega_2 = 1, \tag{2}$$

where N^S denotes labour supply and N^{rest} is labour demand in the rest of the economy. ω_1 and ω_2 (and the residual term ε_t) are due to the log linearisation. Assume that **a)** $\ln(N_t^{rest})$ depends on a set of variables Z_t^{rest}, **b)** U_{t-1} has predictive power for U_{t+1} and that firms use this information, at least. In addition, that **c)** N^S depends linearly on past unemployment due to e.g., "discouraged worker effect", see Pencavel (1986) inter alia, and on a set of explanatory variables Z^S. Then, (1) and (2) imply:

$$U_t = \delta + \rho U_{t-1} + \omega_1 f(U_{t-1}) + \theta' Z_t + \epsilon_t, \tag{3}$$

where $\epsilon_t = \varepsilon_t - \omega_1 v_t$, $\theta' = (-\omega_1 \Gamma_1, \, -\omega_2 \Gamma_2, \, \Gamma_3)$ and $Z_t' = (Z_t, \, Z_t^{rest}, \, Z_t^S)$.

First, consider a linear $f(U_{t-1})$,

$$f(U_{t-1}) = \lambda U_{t-1},$$

which implies

$$U_t = \delta + \kappa U_{t-1} + \theta' Z_t + \epsilon_t. \tag{4}$$

Since

$$\kappa = \rho + \omega_1 \lambda \geq \rho,$$

it follows that the effect of labour market tightness on hiring decisions $\lambda > 0$ (friction) serves to increase the persistence in unemployment. In addition, friction contributes to a higher equilibrium rate of unemployment, since the conditional expectation is

$$\mathsf{E}[U_t \mid U_0, \; Z] = \frac{(1-\kappa^t)}{(1-\kappa)}[\delta + \theta' Z] + \kappa^t U_0 \, , \tag{5}$$

as long as $|\kappa| < 1$. For a large t, $\mathsf{E}[U_t \mid U_0,\, Z]$ can be approximated by

$$\mathsf{E}[U_t \mid Z] \approx \frac{\delta + \theta' Z}{(1-\kappa)}, \tag{6}$$

which implies that the conditional equilibrium unemployment rate is higher in the presence of friction, because $\kappa \geq \rho$. A mean shift in one or more of the forcing variables in Z can shift the equilibrium unemployment rate over time. The unconditional equilibrium rate of unemployment $\mathsf{E}[U_t]$ is constant, approximately $\dfrac{\delta}{(1-\kappa)}$, if $\theta' Z$ is a zero mean process.

However, if $f(U_{t-1})$ is nonlinear, a mean shift in Z is not necessary for a shift in equilibrium unemployment to occur, and low and high unemployment rates can be self sustaining. For example, the perceived difficulty in hiring labour may only impinge on firms' hiring decisions when labour market tightness exceeds a threshold. This can be represented by a logistic function:

$$f(U_{t-1}) = \frac{1}{1+e^{-\xi(U_{t-1}-c)}}, \tag{7}$$

which varies between 0 and 1 implying two extreme equilibria. c is the threshold rate of unemployment and $\xi > 0$ is a steepness parameter, which reflects the strength of firms' response to perceived labour shortage; ξ is likely to rise with the number of firms responding to perceived labour shortage. For a given c and ξ, low and high unemployment rates may reinforce themselves since $U_{t-1} << c$ and $U_{t-1} >> c$ can lead to low and high unemployment equilibria:

$$\mathsf{E}[U_t \mid Z] \approx \begin{cases} \dfrac{\theta' Z}{(1-\rho)} + \dfrac{\delta}{(1-\rho)} \equiv \mu_1 \\ \dfrac{\theta' Z}{(1-\rho)} + \dfrac{\delta + \omega_1}{(1-\rho)} \equiv \mu_2 \end{cases}, \tag{8}$$

where $\mu_1 < \mu_2$. Note that a nonlinear $f(U_{t-1})$ also implies multiple equilibria in sectoral employment. For example, (8) implies

$$\mathsf{E}[\ln(N_t) \mid Z] \approx \begin{cases} \Gamma_1 Z - \mu_1 \\ \Gamma_1 Z - \mu_2 \end{cases} . \tag{9}$$

In order to test whether non-linear friction effects can explain the existing evidence of multiple equilibria in the Norwegian labour market, it is necessary to employ multivariate models. A shortcoming of univariate studies that contain evidence of multiple equilibria is their inability to identify the underlying mechanisms at work, e.g., non-linear adjustment costs or labour hoarding, increasing returns to scale, effects on labour supply or perhaps quite simply a mean shift in one or more of the forcing variables in Z.

The next sections explain industry employment and aggregate unemployment in Norway. Specifically, we estimate generalisations of (1) together with an equation for the rate of unemployment. The average number of working hours per employed wage earner in industry is also included in the empirical model, since changes in working hours (not only persons) affect total labour input.[1]

3 The econometric framework

Consider first the following VEqCM for a vector of variables Y, conditional on a vector of non-modelled variables Z_t:

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} - \alpha(Y - Y^*)_{t-1} + \omega \Delta Z_t + \Omega \varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I). \tag{10}$$

Y^* represents the equilibrium level of Y which depends on the level of the Z variables. In our analysis, the Y vector contains the (natural) logs of employment in Norwegian industry (n), of the average working hours of industrial workers (h) and of the economy wide unemployment rate (u); The $Z-$variables include logs of wage

[1] Also, considerable evidence suggests substitution between working hours and workers, see e.g., Freeman (1998) and the references therein.

costs, indicators of product demand and capital stock. In Section 5.1, we use cointegration analysis within the context of the corresponding VAR model to estimate the relationships that define Y^*, see Johansen (1988) and (1995). Deviations between Y and Y^* in a given period is partially adjusted in the subsequent period: $0 < \alpha < 1$. $\sum_{i=1}^{k} \Gamma_i$ also conveys information about the dynamic behaviour of Y. ΔZ_t represents short run effects of the Z variables. The disturbance term is a vector $\Omega\varepsilon_t$ with zero mean and covariance matrix $\Omega'\Omega$, as ε_t is by assumption an identically, independently distributed vector with standard normal distribution.

The constant parameter VEqCM encompasses the theoretical model in Section 2 for the case of linear cyclical adjustment costs. For example, if the long run employment equation contains the rate of unemployment u, persistence in the unemployment rate can be (partly) ascribed to linear adjustment costs in employment.

A generalisation of (10) that allows for shifts in e.g., the dynamics of Y and the short run effects of forcing variables is given by

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha(s_t)(Y-Y^*)_{t-1} + \omega(s_t)\Delta Z_t + \Omega(s_t)\varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I), \tag{11}$$

with parameters expressed as function of s_t, the state of the economy at time t. This formulation also allows the unspecified exogenous shocks $\Omega(s_t)\varepsilon_t$ to be drawn from state dependent distributions, though normal.[2] We assume that s_t is an unobservable state variable that takes on discrete values in the space $\{1,\ 2,...S\}$ governed by a first-order Markov chain, see e.g., Hamilton (1989) and Krolzig (1997). Since s is unobservable, probabilistic inference about the value of s_t is based on the information available at time τ and the estimated value of the parameter vector Θ, which contains all parameter values in the system for all states. The *filtered* and *smoothed probabilities* of $s_\tau = j$ express the probability of being in state j at time τ, conditional on the information available at time $\tau = t$ and $\tau = T$, respectively.

[2] The case of constant parameters, model (10), corresponds to $s_t = 1$, $\forall\ t$.

For example, the filtered probability can be expressed as:

$$P(s_t = j \mid Y_t,\ Z_t;\ \widehat{\Theta}), \qquad j = 1,\ 2, ...S \text{ and } t = 1, 2, 3...T. \tag{12}$$

A potential shortcoming of model (11) is that it imposes symmetric effects on Y of positive and negative changes in its determinants, in a given state. It is not unlikely that employment responds slower to positive impulses than to negative ones, if e.g., cycle independent hiring costs are larger than the firing costs. The empirical relevance of this shortcoming can be assessed by considering a slightly generalised version of the model with state dependent effects. For example, one may use the following model, which allows for different response to overmanning $(Y - Y^*)^+$ and undermanning $(Y - Y^*)^-$ and to positive and negative changes in the exogenous variables, ΔZ^+ and ΔZ^-, respectively, in state s. Here, superscript "+" denotes that a variable $X^+ = X$ iff $X \geq 0$ while $X^+ = 0$ iff $X < 0$; similarly, $X^- = X$ iff $X \leq 0$ while $X^- = 0$ iff $X > 0$.

$$\begin{aligned} \Delta Y_t \;&=\; \sum_{i=1}^{p} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha^+(s_t)(Y - Y^*)^+_{t-1} - \alpha^-(s_t)(Y - Y^*)^-_{t-1} + \\ &\quad\ \omega^+(s_t)\Delta Z^+_t + \omega^-(s_t)\Delta Z^-_t + \Omega(s_t)\varepsilon_t. \end{aligned} \tag{13}$$

Given the large number of parameters to estimate, (13) requires a relatively large number of observations to provide precise coefficient estimates and conclusive results.

4 Data

The empirical analysis is based on Norwegian seasonally non-adjusted quarterly data over the period 1974(1)-1996(4). The precise definitions, source and the time series properties of the variables are reported in the appendix.

The elements of the Y vector, in levels, are displayed in Figure 1. The number of persons employed in the manufacturing and construction sector displays a

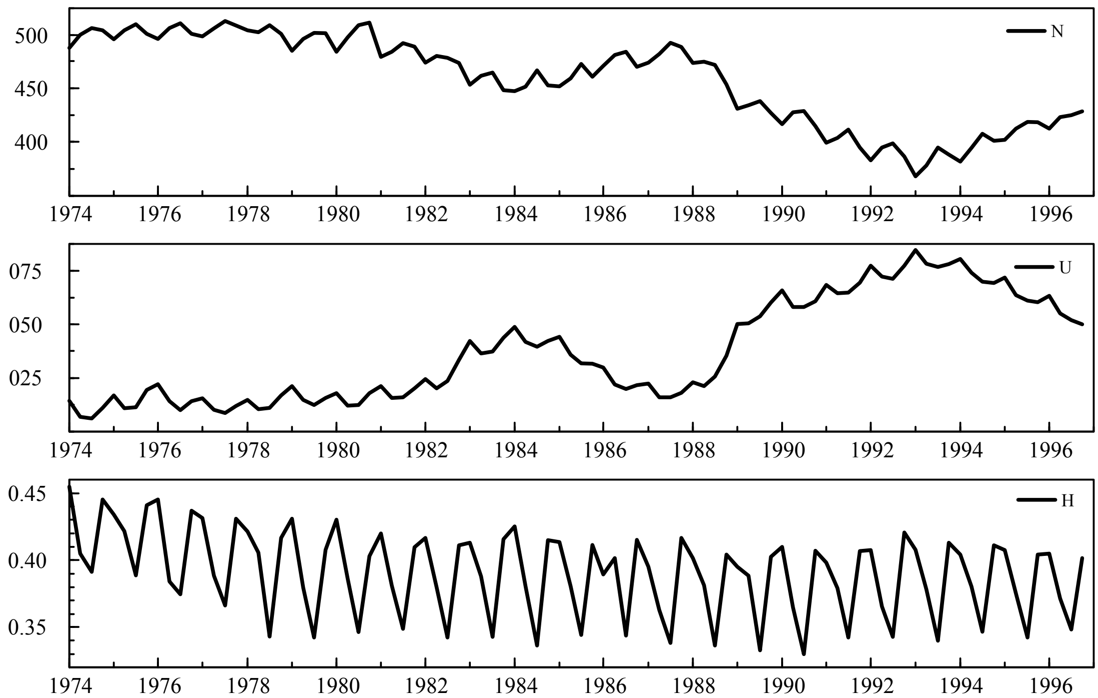

Figure 1: *Time series of the Y variables (in levels) over the sample period: 1974(1)-1996(4). Persons employed in manufacturing and construction in thousands (N), the aggregate unemployment rate (U) and average working hours in manufacturing and construction (H).*

downward trend over the sample period, especially since the late 1980s. In 1993 the employment level is about 25% lower than in 1987. However the number of employed rises from 1993 to the end of the sample.

The aggregate unemployment rate displays large fluctuations from the early 1980s, relative to its subdued behaviour in the 1970s. In 1984 the unemployment rate is more than twice the rate in 1981. In the period 1986-1989 it returns to the low levels of the 1970s. However, there is a large increase in the unemployment rate in 1989/90, and it peaks in 1993 at a rate more than four times higher than the rate in 1981. Despite the downward tendency in the unemployment in the remaining sample period, it evolves at relatively high levels. A number of studies argue that the Norwegian unemployment experienced a structural break in 1989/90 that led to a shift in its long run mean, see e.g., Bianchi and Zoega (1998) and Skalin and Teräsvirta (1999). Similarly, the downward shift in the industry employment in the late 1980s can be interpreted as a shift in the long run mean of the employment.

Average working hours exhibits a downward trend over the whole sample period and seems to be unresponsive to the cyclical variations in the sample. Seasonality though, is pronounced in this time series.

Augmented Dickey Fuller (ADF) tests presented in the appendix suggests that logs of N, U and H (denoted in small letters) may be considered as integrated of order 1.

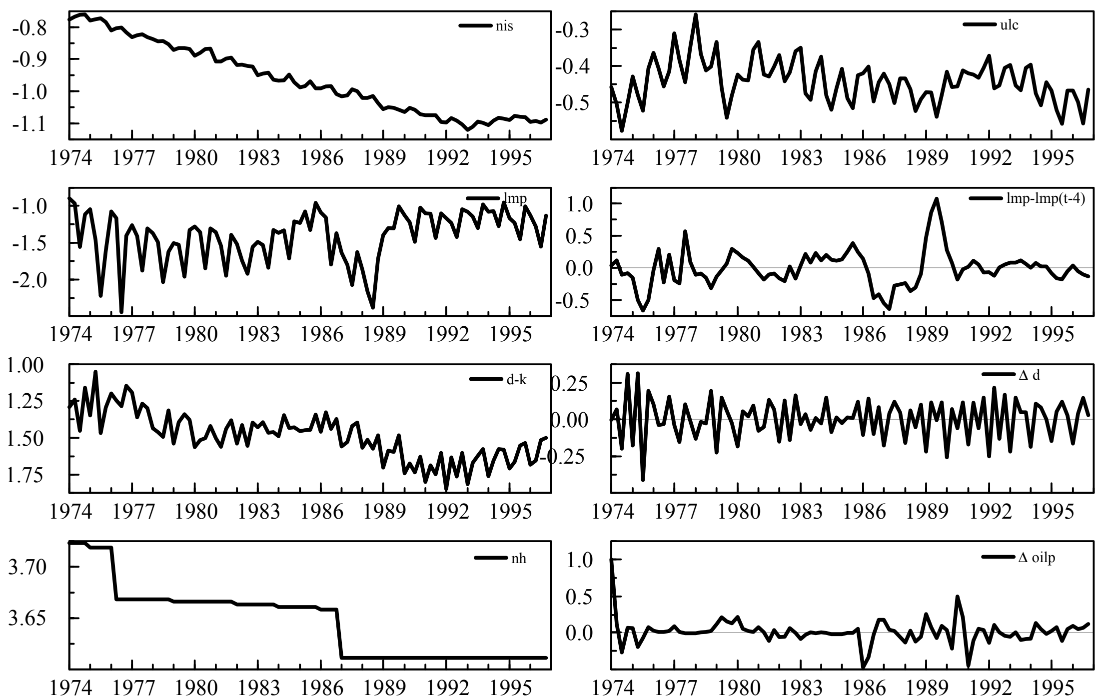

Figure 2: *Time series of the Z variables and their transformations over the period 1974(1)-1996(4). From left (in logs): Share of industry employment in total employment (nis), unit labour costs (ulc), the programme ratio (lmp)* and the annual growth in the *programme ratio* ($\Delta_4 lmp$), *indicator of capacity utilisation ($d-k$), quarterly growth in aggregate demand (Δd), normal working hours (nh) and finally, quarterly growth in crude oil prices ($\Delta oilp$).*

In line with the discussion in Section 2, the vector Z consists of variables that are assumed to determine the dynamics as well as the equilibrium level of Y, Y^*. Specifically, it contains unit labour costs (ulc), normal (institutional) working hours per week (nh), demand relative to the capital stock ($d-k$), the labour market program ratio (lmp), crude oil prices ($oilp$) and finally the share of industry employment in total employment (nis). Figure 2 shows a downward trend in nis over most of the sample period. This trend is negatively correlated with e.g., the secular rise in the female labour participation rate and in part-time work; with technological changes; with the tendency towards decentralisation of the wage bargaining; and with increase in social welfare programs. These structural developments may have contributed to a rise in the unemployment rate over time, see e.g., Dornbusch and Fischer (1994, pp. 511) and Layard et al. (1991).

Most elements of the Z and ΔZ vectors are displayed in Figure 2. The ADF tests indicate the presence of a unit root in the levels of all the series except lmp, which seems to be integrated of order zero, see the appendix.

The following subsection shows that the chosen set of variables enables us to derive data consistent and interpretable models of the endogenous variables.

5 A linear model

We estimated a 5th order VAR for $Y = (n, u, h)$ conditional on the vector Z. The following lags and transformations of the variables in Z were found to be statistically significant and provided a parsimonious representation of the effects of the Z variables: ulc_{t-1}, nh, nis_{t-1}, $(d-k)_{t-1}$, $\Delta_4 lmp_{t-1}$, Δd_t and $\Delta_4 d_t$. In addition, three centred seasonal dummies CS's, a *trend* and three impulse dummies, $i1981q1$, $i1986q1$ and $i1989q2$, were included to control for seasonal effects and to remedy violations of the (standard) assumptions about the residuals.

Table 1 reports the outcome of tests for residual misspecification. The results suggest that the empirical system is adequately specified.

5.1 Cointegration

We next tested for cointegration using the Johansen (1988) procedure, within a system that restricted ulc_{t-1}, nh, nis_{t-1}, $(d-k)_{t-1}$ and a deterministic trend to the cointegration space, while the constant term, $\Delta_4 lmp_{t-1}$, Δd_t, $\Delta_4 d$ and the dummy

Table 1: Diagnostics for 5. order conditional VAR for industry employment, working hours and aggregate unemployment rate (in logs); 1974(1)-1996(4); p-values in square brackets.

	n	u	h	VAR
$F_{ar,\,1-5}(5,53)$	0.57[0.72]	2.22[0.07]	2.05[0.09]	
$F_{arch,\,1-4}(4,50)$	0.19[0.94]	0.91[0.46]	0.69[0.60]	
$F_{het}(38,19)$	0.29[0.99]	0.68[0.85]	0.47[0.98]	
χ^2_{nd}	1.43[0.49]	2.54[0.28]	1.47[0.48]	
$F^v_{ar,\,1-5}(45,122)$				1.42[0.07]
$F^v_{het}(228,91)$				0.39[1.00]
$\chi^{2,v}_{nd}(6)$				6.51[0.38]

variables were entered unrestricted, cf. Harbo et al. (1998) and Doornik et al. (1998). The results are reported in Table 2. It contains the relevant eigenvalues and the associated trace (Tr) statistics employed in testing the hypothesis of $(r-1)$ versus r cointegration vectors. The critical values are from Table 2 in Harbo et al. (1998).

Numerically, all the three eigenvalues are well above zero suggesting three cointegration vectors. Statistically however, the Tr-statistic gives formal support to one cointegrating vector, $r = 1$. Since the test may lack power, we proceed with the assumption that there are three cointegration vectors, and investigate if we can interpret these statistical relationships within the framework of Section 2.

Table 2: Cointegration rank.

r	1	2	3
eigenvalue	0.46	0.21	0.12
Tr	90.47	33.53	11.42
95%	69.7	44.5	20.7

Table 3 therefore imposes relevant restrictions on the β and α vectors, which are jointly acceptable with $\chi^2(11) = 16.86$ [0.11].[3] Figure 3 shows the recursive estimates of the β-coefficients and their 95% confidence intervals denoted as $\pm 2SE$. The estimates of the unrestricted β-coefficients appear statistically significant and stable over the period 1985(1)–1996(4).

[3] The unrestricted system was first re-estimated without a deterministic trend, since testing test (based on $r = 3$) showed that the trend can be excluded from the system, with $\chi^2(3) = 4.6448[0.1997]$.

Table 3: Restricted cointegration analysis, identification of cointegration vectors

$$\begin{bmatrix} \widehat{\beta}' & n & u & h & ulc\_1 & nh & (d-k)\_1 & nis\_1 \\ 1 & -1 & -0.14 & -1 & -0.13 & 0 & 0.20 & 0 \\ & & (0.01) & & (0.05) & & (0.03) & \\ 2 & -1.81 & -1 & 0 & 0 & 0 & 0 & -3.97 \\ & (0.78) & & & & & & (0.64) \\ 3 & 0 & 0 & -1 & 0 & 1 & 0 & 0 \\ & & & & & & & \end{bmatrix},$$

$$\begin{bmatrix} \widehat{\alpha} & 1 & 2 & 3 \\ n & 0.42 & -0.039 & 0 \\ & (0.11) & (0.01) & \\ u & 2.05 & 0.03 & 0 \\ & (0.64) & (0.06) & \\ h & 0 & 0 & 0.30 \\ & & & (0.13) \end{bmatrix}$$

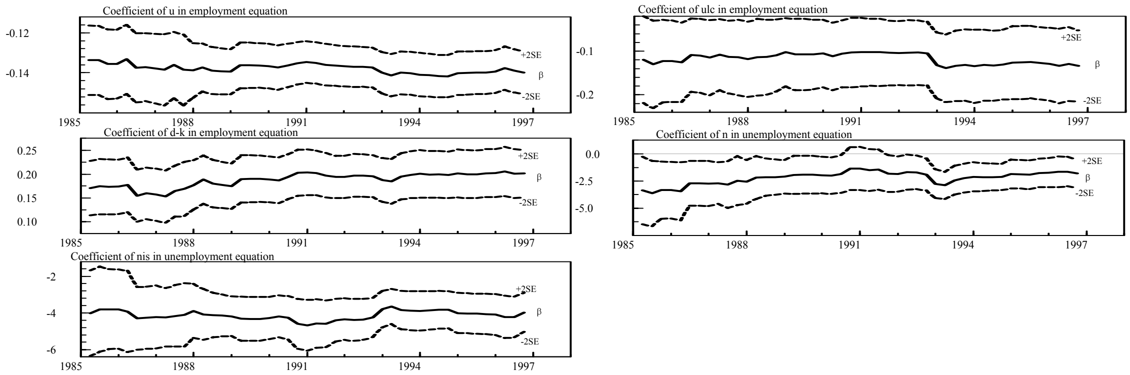

Figure 3: *Recursive estimates of the cointegration vectors with +/-2SE. Initial sample: 1974(1)-1984(4).*

The restricted cointegration vectors are interpretable. A rise in u reduces the equilibrium level of employment which may suggest a reduction in labour hoarding in the face of easier access to labour, see Section 2. h and n appear to be perfect substitutes in the long run, which is consistent with the "labour sharing view". A rise in ulc reduces the employment, consistent with a downward sloping demand curve for labour. The positive coefficient estimate of $(d-k)$ suggests that higher capacity utilisation raises employment, or alternatively, a rise in the capital stock (k) substitutes employment. The second vector implies a reduction in u following a rise in sectoral employment n, though the coefficient estimates are imprecise.

Furthermore, the proxy for structural changes e.g., in labour supply nis, is significant in the long run unemployment equation. The third vector suggests that average working hours follow the institutionally determined working hours nh.

The restricted $\widehat{\alpha}$ matrix in Table 3 shows that both n and u respond to deviations between the actual and the equilibrium values of n and u.[4] The test of joint restrictions on the α and β' matrices accepts the weak exogeneity of h for the long run parameters in the employment and unemployment equations. This seems inconsistent with the common finding that working hours act as a buffer against deviations between actual and equilibrium level of employment, cf. Jacobson and Ohlsson (2000) *inter alia*. However, the more restricted simultaneous equation model in the next subsection does not support the weak exogeneity of hours. This apparently contradictory result may be ascribed to low test power in the unrestricted VAR.

5.2 A simultaneous equation model with linear friction effects

The cointegration analysis implies that $Y - Y^*$ is a 3×1 vector defined as:

$$\begin{aligned} n - n^* &= n - \{0.20(d-k) - h - 0.13(u + ulc)\}, && (14)\\ u - u^* &= u - \{-1.81n - 3.96nis\}, && (15)\\ h - h^* &= h - nh. && (16) \end{aligned}$$

Using these equilibrium correction terms, the conditional VAR model was reformulated as a (conditional) VEqCM of order 4 in the differences. Thereafter, parsimony was sought through data consistent coefficient restrictions. Further, the parsimonious version of the model was reformulated as a structural VEqCM with contemporaneous effects between the endogenous variables, cf. Bårdsen and Fisher (1999) and Boswijk (1995). Accordingly, $(n - n^*)_{t-1}$ was restricted to the equation of Δn_t while $(u - u^*)_{t-1}$ was restricted to the equation of Δu_t. Table 4 presents the pre-

[4] Note that constant terms, which do not appear in the cointegration space may be a part of the equilibrium solutions of n, u and h, as assumed later.

ferred specification of the structural VEqCM which has been estimated by FIML. The diagnostics indicate that the standard assumptions regarding the residuals are not violated at the standard levels of significance. The test for overidentifying restrictions shows that it parsimoniously encompasses the initial VEqCM.

The short run effects of the explanatory variables are interpretable. In particular, a rise in unemployment reduces the growth in employment, which indicates dynamic labour hoarding effects. A rise in aggregate demand increases employment and hours while it reduces unemployment. The latter is also lowered by a rise in the program ratio, higher oil prices and by a reduction in normal working hours.

In this structural VEqCM, actual working hours act as a buffer against undermanning ($(n - n^*) < 0$) and overmanning ($(n - n^*) > 0$) in the short run. Thus the weak exogeneity of hours is rejected relative to the long run parameters in the employment equation. Also, working time adjusts faster towards its equilibrium level than employment and unemployment. However, there seems to be high degree of negative autoregression in hours, probably reflecting the pronounced seasonal variation in working hours.

The model in Table 4 is considered as an empirical counterpart to the theoretical model in Section 2, with linear specification of friction effects $f(U_{t-1})$. Section 2 shows that a non-linear $f(U_{t-1})$ implies multiple equilibria; more specifically, shifts in the long run means of $u - u^*$ and $n - n^*$, as u^* and n^* are interpreted as counterparts to $\dfrac{\theta' Z}{(1-\rho)}$ and $\Gamma_1 Z + \dfrac{\theta' Z}{(1-\rho)}$ in Section 2. To investigate this possibility we defined $f(U_{t-1})$ as a logistic function of U_{t-1}, as in equation (7). The value of the threshold parameter (c) was set to 0.04 and that of the steepness-parameter (ξ) to 100; since estimates of c and ξ were found to be quite imprecise when the method of Maximum Likelihood was applied to the employment equation in Table 4, cf. Teräsvirta (1998). Consequently, $f(U_{t-1})$ behaves as a step function with a value close to 1 (high friction) when $U_{t-1} < 0.04$ and close to 0 (low friction) when $U_{t-1} > 0.04$.

Notably, the joint test of the significance of the logistic $f(U_{t-1})$ when added

Table 4: Simultaneous equation model with linear friction effects

Industry employment

$$\widehat{\Delta n_t} = \underset{(0.139)}{0.543} - \underset{(0.010)}{0.033}\,\Delta u_t - \underset{(0.009)}{0.028}\,\Delta u_{t-2} - \underset{(0.026)}{0.152}\,\Delta h_t$$
$$+\underset{(0.089)}{0.183}\,\Delta n_{t-4} - \underset{(0.028)}{0.110}\,(n-n^*)_{t-1} + \underset{(0.011)}{0.032}\,\Delta d_t$$
$$-\underset{(0.011)}{0.032}\,i81q1_t + \underset{(0.011)}{0.022}\,i86q1_t + \underset{(0.005)}{0.012}\,CS_{t-1}$$
$$\hat{\sigma}_n = 1.089\%$$

Aggregate unemployment

$$\widehat{\Delta u_t} = \underset{(0.130)}{0.735} - \underset{(0.386)}{0.523}\,\Delta_3 n_t + \underset{(0.063)}{0.343}\,\Delta u_{t-1} - \underset{(0.069)}{0.150}\,\Delta u_{t-2}$$
$$+\underset{(0.050)}{0.171}\,\Delta u_{t-3} + \underset{(0.072)}{0.467}\,\Delta u_{t-4} - \underset{(0.034)}{0.192}\,(u-u^*)_{t-1}$$
$$-\underset{(0.044)}{0.156}\,\Delta op_{t-1} - \underset{(0.075)}{0.240}\,\Delta_4 d_t + \underset{(0.539)}{2.04}\,\Delta_4 nh_t$$
$$-\underset{(0.027)}{0.103}\,\Delta_4 lmp_{t-1} - \underset{(0.026)}{0.150}\,CS_{t-1}$$
$$\hat{\sigma}_u = 6.66\%$$

Industry hours

$$\widehat{\Delta h_t} = \underset{(0.171)}{0.520}\,\Delta_4 nh_t - \underset{(0.151)}{0.660}\,\Delta h_{t-1} - \underset{(0.148)}{0.690}\,\Delta h_{t-2} - \underset{(0.122)}{0.557}\,\Delta h_{t-3}$$
$$-\underset{(0.08)}{0.230}\,\Delta h_{t-4} + \underset{(0.023)}{0.029}\,\Delta d_t - \underset{(0.160)}{0.287}\,\Delta n_{t-4} - \underset{(0.079)}{0.253}\,(h-h^*)_{t-1}$$
$$-\underset{(0.074)}{0.234}\,(n-n^*)_{t-1} - \underset{(0.021)}{0.065}\,i86q1_t - \underset{(0.022)}{0.033}\,i89q1_t$$
$$-\underset{(0.022)}{0.025}\,CS_{t-1} - \underset{(0.022)}{0.114}\,CS_{t-2}$$
$$\hat{\sigma}_h = 2.037\%$$

Diagnostics

AR $1-5$ $F(45,190)$	=	1.091[0.34]
Normality $\chi^2(6)$	=	3.544[0.74]
Heteroscedasticity $F(276,181)$	=	1.02[0.45]
Overidentification $\chi^2(46)$	=	56.89[0.13]

FIML estimates. The sample is 1974(1)–1996(4). Standard errors in parentheses below the coefficient estimates. p-values in square brackets.

to the employment and the unemployment equations in Table 4 yielded $\chi^2(2) = 0.02321[0.9885]$, lending no support to non-linear friction effects and the possibility of friction induced shifts in the long run means of $u - u^*$ and $n - n^*$. Furthermore, the recursive stability of the equilibrium means of $n - n^*$ and $u - u^*$ in Figure 4 suggests that possible changes in the marginal means of u and n should be attributed to the non-modelled variables and not to labour market friction effects. The figure displays recursive estimates of the means of $n-n^*$, $u-u^*$ and $h-h^*$ over the period 1977(1)–1996(4). The stability of the parameter estimates defining n^*, u^* and h^* is shown above, in Figure 3.

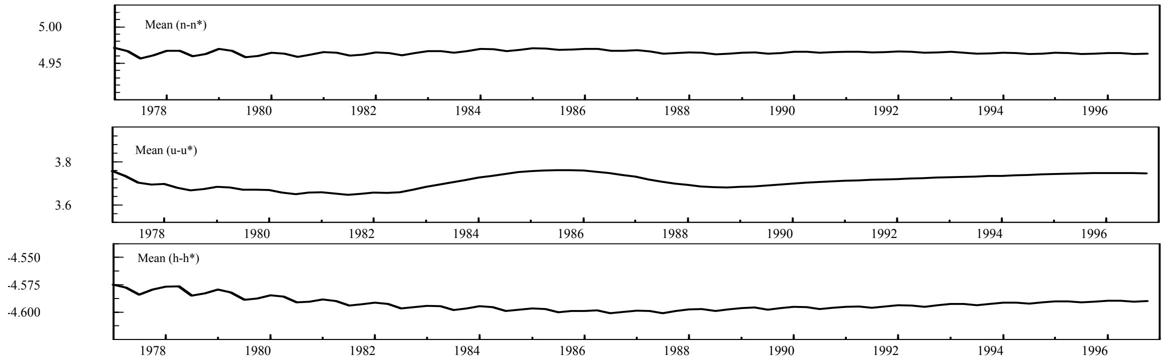

Figure 4: *Recursive estimates of the means of* $n - n^*$, $u - u^*$ *and of* $h - h^*$ *over the period 1977(1)–1996(4). The initial estimates are based on observations from the period 1974(1)–1976(4).*

Apparently, tests of the overall stability of the structural VEqCM in Figure 5 do not suggest non-constancies in the parameters. There are no outliers among the 1-step ahead residuals and none of the scaled Chow statistics exceeds the critical value of 1 over the period 1985(1)–1996(4).

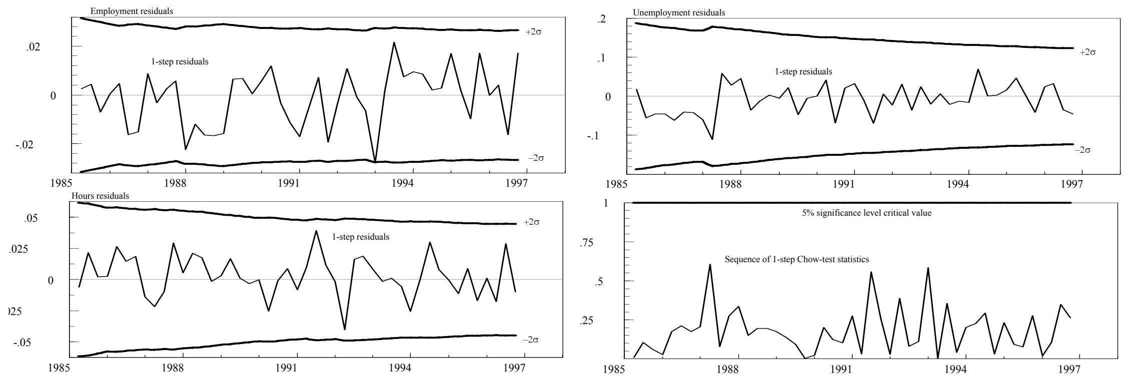

Figure 5: *1-step ahead residuals* ± 2 *estimated standard errors based on the equations of employment, unemployment and hours. Also, a sequence of 1-step Chow tests scaled by their critical values at the 5% level of significance.*

However, these tests may understate possible non-constancy in the short run parameters of the VEqCM because the long run parameters appear remarkably constant over the sample in Figures 3 and 4. Hendry (2000) shows that even large shifts in short run parameters, representing e.g., the dynamics, adjustment speeds and intercepts, are difficult to detect if parameters defining the long run equilibrium remain unaltered. Note that the full sample estimates of the long run means of $n-n^*$ and $u-u^*$ in Figure 4 are close to the derived long run estimates of the composite constant terms in Table 4, $0.543/0.11 \approx 4.94 \approx 5$ and $0.735/0.192 \approx 3.84 \approx 4$. This suggests that the composite constant terms in the employment and unemployment equations mainly consists of the evidently stable equilibrium means of $n-n^*$ and $u-u^*$, times the associated equilibrium correction coefficients; Implicitly, other components of the composite constant terms, including the autonomous growth rates in employment and unemployment, seem to be numerically small or to outweigh each others. In the equation for hours, the equilibrium mean of $n-n^*$ seems to be cancelled by the equilibrium mean of $h-h^*$, which may explain the insignificance and hence the exclusion of a constant term in the hours equation, see Figure 4.

Section 6 investigates whether the short run parameters of the VEqCM, characterising persistence in employment and unemployment and their response to changes in exogenous variables, depend on the phase of the economy. In line with the common practice, we assume that a model of hours (h) with state dependent parameters is not called for. Commonly, adjustment in working hours is modelled independently of the phase of the economy since costs in adjusting hours are small relative to the costs associated with adjusting persons, see e.g., Hamermesh and Pfann (1996) and Bosworth et al. (1996). The time series of H in Figure 1 lends support to this practise.

6 State dependent adjustment

The employment and unemployment equations in Table 4 were estimated separately with $S = 2$.[5] The estimation was conducted by Maximum Likelihood (ML) using a version of the Expectation Maximisation (EM) algorithm proposed by Hamilton (1990), see Krolzig (1997). The parameter estimates and the series of filtered and smoothed probabilities are obtained jointly by iterations between (preliminary) estimates of the parameters and those of the probabilities. The ML estimators are consistent and asymptotically normal under quite general regularity conditions, see e.g., Hamilton (1993) and (1996) Krolzig (1997).

The outcomes for the employment and the unemployment equations are presented in Table 5 where a recession corresponds to $s = 1$ while an expansion phase corresponds to $s = 2$. The classification of e.g., $s = 2$ as an expansion phase is based on the observed features of N and U in Figure 1 and the filtered and smoothed probabilities of $s_t = 2$ for the employment and unemployment in Figure 6.

[5] Results based on $S = 3$ turned out to be difficult to interpret. Also, estimation of both equations when all (short run) parameters in both equations were subjected to common shifts (i.e. imposing a common cycle) seemed infeasible; In particular, estimation of the reduced form of these equations subject to common shift led to failure of convergence for both $S = 3$ and $S = 2$.

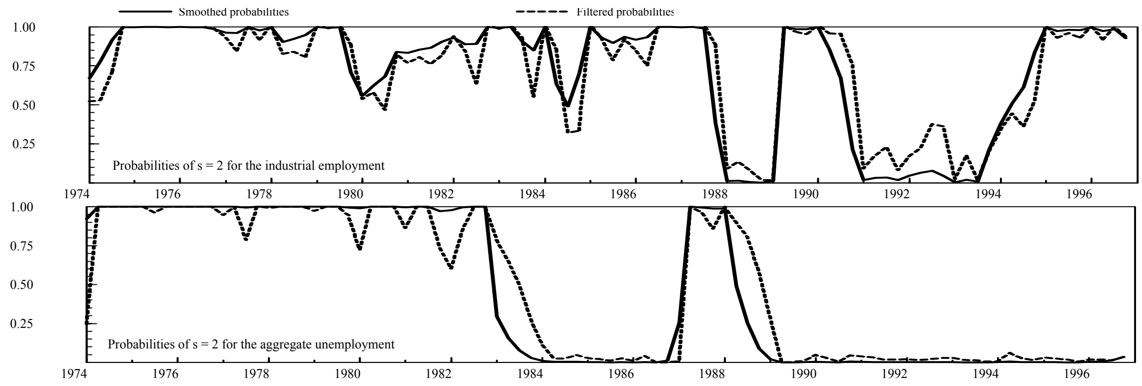

Figure 6: *The filtered and smoothed probabilities of industrial employment and aggregate unemployment being in state 2: the expansion phase.*

Figure 6 suggests some differences in the cycles of the industry employment and the aggregate unemployment rate. Notably, the dates of switches between the contraction and the expansion phases are different from about 1984. In particular, the probabilities related to the unemployment series suggest a contraction even after 1993, in contrast to the probabilities related to employment. This is not surprising given that the unemployment rate was still more than twice its size in the 1970s and the early 1980s. Also, the filtered and smoothed probabilities based on the unemployment behaviour offer a clearer classification into the two regimes than the corresponding probabilities for the employment behaviour.

The explanatory power of the models has increased substantially by making allowance for state dependent parameters, especially, in the state of recession. In the case of the employment equations, the standard deviations of the residuals have declined by 1/3 and 1/4 in the state of recession and expansion, respectively, relative to the size of the standard error in the model with constant parameters. There is also a substantial improvement in the fit of the unemployment equation in the state of recession, though a slight deterioration in the state of expansion; $\widehat{\sigma}_{u,\ 1}$ is 2.63% and $\widehat{\sigma}_{u,\ 2}$ is 7.55% against $\widehat{\sigma}_u = 6.66\%$.

Table 5 shows that employment adjustment is highly state dependent; it adjusts much faster towards its equilibrium value and is more responsive to shocks

Table 5: Models with state dependent parameters.

Industry employment

In recession:

$$\widehat{\Delta n_t} = \underset{(0.297)}{1.410} - \underset{(0.015)}{0.062}\,\Delta u_t - \underset{(0.020)}{0.071}\,\Delta u_{t-2} - \underset{(0.041)}{0.308}\,\Delta h_t$$
$$- \underset{(0.148)}{0.081}\,\Delta n_{t-4} - \underset{(0.060)}{0.285}\,(n-n^*)_{t-1} + \underset{(0.023)}{0.050}\,\Delta d_t$$
$$- \underset{(0.019)}{0.007}\,i81q1_t + \underset{(0.0251)}{0.020}\,i86q1_t + \underset{(0.010)}{0.021}\,CS_{t-1}$$
$$\hat{\sigma}_{n,\,1} = 0.697\%$$

In expansion:

$$\widehat{\Delta n_t} = \underset{(0.130)}{0.414} - \underset{(0.008)}{0.016}\,\Delta u_t - \underset{(0.008)}{0.015}\,\Delta u_{t-2} - \underset{(0.024)}{0.123}\,\Delta h_t$$
$$+ \underset{(0.084)}{0.209}\,\Delta n_{t-4} - \underset{(0.026)}{0.083}\,(n-n^*)_{t-1} + \underset{(0.010)}{0.022}\,\Delta d_t$$
$$- \underset{(0.010)}{0.039}\,i81q1_t + \underset{(0.009)}{0.021}\,i86q1_t + \underset{(0.005)}{0.012}\,CS_{t-1}$$
$$\hat{\sigma}_{n,\,2} = 0.846\%$$

Aggregate unemployment

In recession:

$$\widehat{\Delta u_t} = \underset{(0.102)}{0.458} - \underset{(0.226)}{0.535}\,\Delta_3 n_t + \underset{(0.087)}{0.365}\,\Delta u_{t-1} - \underset{(0.048)}{0.006}\,\Delta u_{t-2}$$
$$+ \underset{(0.043)}{0.162}\,\Delta u_{t-3} + \underset{(0.082)}{0.440}\,\Delta u_{t-4} - \underset{(0.027)}{0.121}\,(u-u^*)_{t-1}$$
$$- \underset{(0.029)}{0.038}\,\Delta op_{t-1} - \underset{(0.076)}{0.032}\,\Delta_4 d_t + \underset{(0.812)}{1.53}\,\Delta_4 nh_t$$
$$- \underset{(0.025)}{0.132}\,\Delta_4 lmp_{t-1} - \underset{(0.023)}{0.132}\,CS_{t-1}$$
$$\hat{\sigma}_{u,\,1} = 2.63\%$$

In expansion:

$$\widehat{\Delta u_t} = \underset{(0.102)}{0.458} - \underset{(0.766)}{0.228}\,\Delta_3 n_t + \underset{(0.101)}{0.244}\,\Delta u_{t-1} - \underset{(0.110)}{0.258}\,\Delta u_{t-2}$$
$$+ \underset{(0.086)}{0.090}\,\Delta u_{t-3} + \underset{(0.110)}{0.343}\,\Delta u_{t-4} - \underset{(0.029)}{0.112}\,(u-u^*)_{t-1}$$
$$- \underset{(0.085)}{0.227}\,\Delta op_{t-1} - \underset{(0.104)}{0.246}\,\Delta_4 d_t + \underset{(0.732)}{2.489}\,\Delta_4 nh_t$$
$$- \underset{(0.052)}{0.133}\,\Delta_4 lmp_{t-1} - \underset{(0.051)}{0.205}\,CS_{t-1}$$
$$\hat{\sigma}_{u,\,2} = 7.55\%$$

The sample is 1974(1) to 1996(4), 92 observations. Asymptotic standard errors in parentheses. Estimation by the EM algorithm.

in its determinants during a recession than in an expansion, which tends to be characterised by shortage of labour. In recession, the autoregressive coefficient is insignificantly different from zero and the absolute value of the estimated equilibrium correction coefficient is more than three times its size than in the expansion phase of the economy, 0.285 versus 0.083. Furthermore, the coefficient estimates of all the other regressors (except the impulse dummies) tend to double, at least, when there is a switch from expansion to recession. The corresponding coefficient estimates in Table 4 are largely in-between the state dependent coefficient estimates. This implies that a linear (constant parameter) characterisation of the employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions.

However, despite the clear differences in the employment response across the two states, the equilibrium solution of the employment remains the same across the two states and close to that found in the case of the linear model. Note that the constant term in the equilibrium solution, i.e., the ratio between the state dependent intercept and the equilibrium correction coefficient, is the same across the two states: $1.410/0.285 \approx 0.414/0.083 \approx 5$. The stability of the estimated equilibrium is consistent with the outcome of the test about the significance of the logistic $f(U_{t-1})$ and the demonstrated stability of the sample mean of $n - n^*$ in Section 5.2.

The above results on the short run and the long run employment behaviour remained robust to an extension of the model where we allowed for asymmetric response to positive and negative changes in the regressors, as reported below in Subsection 7. Asymmetric response may occur if e.g., hiring costs are larger than firing costs, as observed by Hamermesh and Pfann (1996) inter alia.

Interestingly, the results for the unemployment rate suggest that it responds more strongly to shocks in a tight labour market than in a slack market. Firstly, the degree of persistence is much higher in a slack labour market compared with when it is tight, though the equilibrium correction coefficients appears as state independent. Secondly, the effects of most of the other determinants are found to be stronger in an expansion than in a recession. In particular, the effects of changes in demand and oil prices are much stronger in an expansion than in recession. However, the equilibrium solution of unemployment is almost the same across the two states; the derived estimates of the constant terms in the equilibrium solution are 0.458/0.112 $\approx$ 0.458/0.121 $\approx$ 4, as in the case of the linear model. This adds to the evidence of the stability of the long run mean of $u - u^*$.

The relatively sluggish response of unemployment in a slack labour market may be an implication of the "discouraged workers effect", see e.g., Pencavel (1986) and Bosworth et al. (1996). In a slack labour market, positive impulses from e.g., oil prices, aggregate demand or a reduction in working hours raise participation rates, in addition to employment opportunities. This may dampen their effects on the unemployment rate. In a tight labour market, however, the reserves of labour supply are (relatively) exhausted, i.e. the labour supply curve is inelastic, hence the rate of unemployment falls rapidly in response to an increase in employment opportunities.

7 Asymmetric response to shocks?

The employment response may depend on the sign of a shock since hiring costs are believed to be higher than firing costs, see Hamermesh and Pfann (1996) inter alia. Table 6 presents a generalised version of the employment equation in Table 5 where the employment is allowed to respond asymmetrically to positive and negative shocks, as in equation (13). Specifically, in each of the two states, the response is allowed to vary with positive and negative deviations from the equilibrium employment and to positive and negative changes in the other regressors, except the autoregressive and the deterministic terms.

The increased flexibility of this model has led to a large reduction in the standard errors of the residuals in both states. However, the coefficient estimates are less precise than in the previous models. Also, the coefficient estimate of Δn_{t-4}

Table 6: Model with sign and state dependent parameters.

Industry employment

In recession:

$$\begin{aligned}
\widehat{\Delta n}_t =\ & \underset{(0.139)}{0.874} + \underset{(0.080)}{0.416}\,\Delta n_{t-4} - \underset{(0.010)}{0.067}\,\Delta u_t^+ + \underset{(0.009)}{0.002}\,\Delta u_t^- \\
& + \underset{(0.010)}{0.026}\,\Delta u_{t-2}^+ - \underset{(0.016)}{0.108}\,\Delta u_{t-2}^- - \underset{(0.027)}{0.196}\,\Delta h_t^+ - \underset{(0.035)}{0.088}\,\Delta h_t^- \\
& - \underset{(0.028)}{0.176}\,(n-n^*)_{t-1}^+ - \underset{(0.028)}{0.179}\,(n-n^*)_{t-1}^- + \underset{(0.019)}{0.055}\,\Delta d_t^+ + \underset{(0.020)}{0.056}\,\Delta d_t^- \\
& - \underset{(0.005)}{0.0050}\,i81q1_t + \underset{(0.007)}{0.022}\,i86q1_t + \underset{(0.005)}{0.003}\,CS_{t-1}
\end{aligned}$$

$$\hat{\sigma}_{n,\,1} = 0.380\%$$

In expansion:

$$\begin{aligned}
\widehat{\Delta n}_t =\ & \underset{(0.236)}{0.331} + \underset{(0.075)}{0.193}\,\Delta n_{t-4} - \underset{(0.013)}{0.010}\,\Delta u_t^+ + \underset{(0.015)}{0.018}\,\Delta u_t^- \\
& + \underset{(0.012)}{0.004}\,\Delta u_{t-2}^+ - \underset{(0.012)}{0.002}\,\Delta u_{t-2}^- - \underset{(0.043)}{0.199}\,\Delta h_t^+ - \underset{(0.034)}{0.066}\,\Delta h_t^- \\
& - \underset{(0.047)}{0.065}\,(n-n^*)_{t-1}^+ - \underset{(0.048)}{0.062}\,(n-n^*)_{t-1}^- + \underset{(0.022)}{0.000}\,\Delta d_t^+ + \underset{(0.018)}{0.062}\,\Delta d_t^- \\
& \underset{(0.012)}{0.042}\,i81q1_t + \underset{(0.009)}{0.025}\,i86q1_t + \underset{(0.006)}{0.020}\,CS_{t-1}
\end{aligned}$$

$$\hat{\sigma}_{n,\,2} = 0.687\%$$

The sample is 1974(1) to 1996(4), 92 observations. Asymptotic standard errors in parentheses. Estimation by the EM algorithm.

has become larger in recession than in expansion, relative to the estimates in Table 5; The opposite has happened in the case of CS_{t-1}. These changes possibly call for a more adequate representation of seasonal effects in the model.

Table 6 offers mixed evidence of asymmetric response to positive and negative changes in the explanatory variables. In particular, the response to over- and undermanning, $(n-n^*)_{t-1}^+$ and $(n-n^*)_{t-1}^-$, is symmetric across the two states.[6] The exceptions are the response to changes in working hours (Δh) and in aggregate demand (Δd) which appear asymmetric. The coefficient estimates of Δh_t^+ are more than twice the size of the coefficient estimates of Δh_t^- in both states, suggesting that a reduction in employment can be achieved faster than an expansion. As regards the demand shocks, the coefficient estimate of Δd_t^+ is zero while that of Δd_t^- is

[6] When deriving the series of $(n-n^*)^+$ and $(n-n^*)^-$, the sample mean of $(n-n^*)$ was subtracted from $(n-n^*)$.

0.062 in an expansion. This finding also suggests that a reduction is easier than an expansion. However, in a recession, this asymmetry seems to disappear as the coefficient estimate of Δd_t^+ and of Δd_t^- are almost identical.

However, Table 6 substantiates the evidence in favour of state dependent employment response to shocks. The explanatory variables have generally a bigger impact on employment in a recession than in an expansion. Particularly, the response to over- and undermanning is almost three times bigger in a recession than in an expansion. Furthermore, a positive shift in aggregate demand leaves employment unaffected if it occurs in an expansion. The response to changes in working hours, however, seems to depend more on the sign of a change than on the state of the labour market.

The table also supports the relevance of the dynamic friction effects, at least if we look at the case of a recession. (In the state of expansion, the estimates of the unemployment terms become small relative to those in the state of recession and statistically insignificant at the 5% level). Also, in this highly non-linear model, the implied equilibrium solution of employment is the same in both states and equal to that implied by the models in Table 4 and 5.

To summarise, the results supports state dependence in the employment response even when one allows for asymmetric response to positive and negative changes in the explanatory variables. The results also suggest that in general there are not considerable differences in the employment response to positive and negative changes. Hence one could argue that, for the sake of parsimony it suffices to make allowance for just state dependence in the parameters.

8 Conclusions

The empirical evidence in this paper shows that the dynamic behaviour of Norwegian industry employment alters with shifts between slack and tight labour markets. Specifically, employment adjusts more rapidly towards its equilibrium level and responds more strongly to changes in exogenous variables in a slack labour market than in a tight labour market. Moreover, anticipated difficulties in hiring due to labour shortage contribute to labour hoarding and employment persistence. These conclusions have appeared robust to allowance for asymmetric response to shocks.

The derived equilibrium solutions of the industry employment and the aggregate unemployment rate have, however, been found to be invariant to cyclical and structural changes over the sample period. Thus our evidence does not support the view that hiring difficulties alone can lead to multiple equilibria. Instead, shifts in the long run means of the variables are shown to depend on other factors, product demand relative to capacity and unit labour costs in particular. In sum, we find that adjustment costs affect the dynamic adjustment and not the long run equilibrium.

The evidence of cycle dependent employment behaviour implies that a linear (constant parameter) characterisation of the employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions. Our results demonstrate that such shortcomings of linear models may be overlooked by conventional tests of parameter non-constancy in samples of typical size.

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Appendix: Data definitions and properties

The data set has been extracted from the database of RIMINI: the quarterly macro-econometric model used in Norges Bank (The Central Bank of Norway). Square brackets include the variable name in the RIMINI data base.

- CS: Centred seasonal for the first quarter in a year.
- D: Indicator of aggregate demand. [DEMIBA.2].
- H: Average working hours per employed wage earner in manufacturing and construction. Thousand hours. [FHIBA].
- i19yy:q1: Impulse dummy, 1 in 19yy:1 and zero elsewhere.
- K: Stock of physical capital in manufacturing and construction. Mill. 1993 NOK. [KIBA].
- LMP: Number of unemployed on labour market programs divided by total unemployment. [AMUN].

- NH: Normal weekly working hours in Norway. Hours. [NH]
- N: Employment in manufacturing and construction. 1000 persons. [NWIBA].
- NIS: Total employment in manufacturing and construction relative to total employment in mainland Norway. Rate. [NWIBA/NWF].
- OILP: Spot price of Brent Blend crude oil in US $, indexed. [OLJEPIND].
- U: Total unemployment rate as a fraction of total labour force. [UTOT2].
- ULC: Unit labour costs (inclusive pay roll tax) in manufacturing and construction deflated by the producer price index. 1993 NOK. [WCIBA/PYIBA.ZYIBA].

Table 7: ADF tests of unit roots; 1974(1)-1996(4)

Variables	$\widehat{\alpha}$	t-ADF	ADF(k)
Δn	0.370	-2.966*	5
n	0.866	-3.035	8
Δu	0.438	-3.394*	8
u	0.831	-3.344	12
Δh	-3.092	-3.293*	8
h	0.774	-1.795	11
Δulc	-0.444	-4.198**	6
ulc	0.714	-2.930	8
$\Delta(d-k)$	-1.593	-3.527**	7
$d-k$	0.492	-2.892	12
Δnis	-0.018	-3.846**	3
nis	0.963	-0.704	4
Δd	-1.726	-3.656**	7
d	0.667	-2.322	8
lmp	0.619	-3.714*	8
$\Delta oilp$	0.081	-6.570**	2
$oilp$	0.938	-1.798	3

Note: Initially, 12 lags ($\approx$12(T/100)$^{1/4}$; T = 92, see Schwert (1989)), were allowed for in each of the ADF-models, which contained a constant when testing for a unit root in the 1. difference of a variable and both a constant and trend when testing for in the level of a variable. k denotes the largest significant lag at the 5% level. Lags of order $>k$ were excluded from the models. 5% DF-critical value when a constant and trend: -3.456; when a constant, the 5% and the 1% DF-values are -2.893 and -3.503.