Econometric modelling of 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 firing 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 shortages lead to multiple equilibria in (un)employment. We develop both a linear equilibrium correction model (ECM) and a two- state Markov switching version of the ECM. The models are based on quarterly data for Norwegian industry employment and aggregate unemployment in the period 1974–1996. We find clear evidence of state-dependent adjustment and responses to changes in forcing variables. Yet equilibrium solutions for employment and unemployment appear invariant to cyclical and structural changes in the sample.

Key words: 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 responses to shocks, labour hoarding and asymmetric cycles in employment and GDP (e.g. Hamermesh and Pfann, 1996; 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 characterized as functions of labour shortage measures, e.g. the unemployment rate (e.g. Ball and Cyr, 1966; Peel and Walker, 1978; Burgess, 1992). Presumably, labour shortages raise hiring costs by increasing search costs for suitable workers and make 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 address the possible joint occurrence of all these aspects of cycle-dependent adjustment costs, i.e. cycle dependency of: (i) the adjustment process, (ii) effects of changes in forcing variables and (iii) multiple equilibria. Rather, the existing studies typically present evidence of (i) or (ii), but not of both (i) and (ii) occurring jointly (e.g. Smyth, 1984; Acemoglu and Scott, 1994; Burgess, 1992). Moreover, an increasing number of studies report evidence of multiple unemployment equilibria (e.g. Skalin

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and Teräsvirta, 2002; Bianchi and Zoega, 1998). 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 (e.g. Cooper and John, 1985; Manning, 1990; Murphy et al., 1989; Saint-Paul, 1995).

Compared to the existing studies, our study encompasses all three aspects of adjustment costs, using multivariate models of employment and unemployment that condition on relevant forcing variables. Econometrically, we build on Krolzig (2001) who employs a Markov regime switching vector equilibrium correcting model (MS-VECM) to allow for state dependence in the parameters (cf. Hamilton, 1989). In his two-step approach, cointegration between US employment and output is established by following the procedure developed by Johansen (1995). Thereafter, the vector autoregressive model (i.e. VAR) is reformulated as a vector equilibrium correction model (VECM) 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 VECM (cf. Bårdsen and Fisher, 1999). In the second step, we allow the parameters of the structural VECM to shift in the Markov way.

The rest of the paper is organized as follows: Section 2 presents the modelling framework while Section 3 presents the data set, which consists of seasonally unadjusted quarterly observations over the period 1974q1–2003q4. Estimation is conducted on data for the period 1974q1–1996q4 while the remaining observations are retained for post-sample evaluation of estimated models. Section 4 contains the structural VECM for industry employment, hours and aggregate unemployment. We test for friction-induced multiple equilibria within the context of this model. Section 5 presents the results for the models with state-dependent dynamics. Section 6 concludes while the appendix contains precise definitions of the variables, their source and tests of their time series properties.

2 The modelling framework

A large number of studies assume that present and anticipated labour shortages contribute to unemployment persistence by raising employment adjustment costs (e.g. Ball and Cyr, 1966; Peel and Walker, 1978; Burgess, 1992). Moreover, Moene et al. (1997) suggest that anticipated labour shortages, which is termed friction, may induce multiple unemployment equilibria. A common feature of all these models is that firms are assumed to respond directly (not only to changes in e.g. the real wage but) to "the state of"the current and future labour market. To test the various hypotheses about friction effects, an important step is to actually model labour demand, using for example a cross section of individual firms, or times series of a sub-sector of the economy. We would like to especially examine the relationship between labour demand and the overall unemployment rate in a closed system. We therefore choose the latter option and model sectoral labour demand, specifically in the Norwegian manufacturing and construction sector, henceforth also referred to as "industry"; see Section 3. Industry employment comprises 25% of all civilian employment in Norway.

We particularly investigate whether industry labour demand depends on a variable representing anticipated labour shortage. In the following:

$$f(U^e_{t+1}), \quad f' \geq 0,$$

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, a high U^e_{t+1} presumably leads to 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*. The further consequences of friction in a sector of the economy on the actual unemployment rate depends on the specification of the friction function. A linear friction function implies persistence in the rate of unemployment (hence it explains part of the observed autocorrelation of U_t), but does not by itself induce multiple equilibria in unemployment.

A non-linear f-function may imply two or more unemployment equilibria. For example, the perceived difficulty in hiring labour may only impinge on firms' hiring decisions when labour market tightness exceeds a threshold. We represent this possibility by the following logistic function:

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

which varies between 0 and 1, thus implying two equilibria. c is the threshold rate of unemployment and $\xi > 0$ is a steepness parameter, which reflects the strength of firms' responses to perceived labour shortages; ξ is likely to rise with the number of firms responding to perceived labour shortages. 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, respectively.

In the following sections we test for both linear and non-linear friction effects within an econometric model with three endogenous variables: industry employment, aggregate unemployment and the average number of working hours per employed wage earner in industry. Hours are included in the empirical model since there might be substitution between working hours and workers (e.g. Jacobson and Ohlsson, 2000).

Consider the following VECM for the three variables, contained in the vector Y, conditional on a vector of non-modelled variables Z_t:

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} - \alpha (Y - Y^*)_{t-1} + \omega \Delta Z_t + \Omega \varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I). \quad (2)$$

Y^* represents the equilibrium level of Y which depends on the level of the Z variables. Specifically, the Y vector contains the (natural) logs of employment in Norwegian industry (n), the average working hours of industrial workers (h) and the economy-wide unemployment rate (u); The $Z-$variables include logs of wage costs, indicators of product demand and capital stock. In Section 4.1, we use cointegration analysis within the context of the corresponding VAR model to estimate the relationships that define Y^*; see Johansen (1995). Deviations between Y and Y^* in a given period are 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 VECM encompasses linear cyclical adjustment costs, termed linear friction effects above. 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 generalization of (2) that allows for shifts in e.g. the dynamics of Y and the short run effects of forcing variables is given by

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha(s_t)(Y-Y^*)_{t-1} + \omega(s_t)\Delta Z_t + \Omega(s_t)\varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I), \tag{3}$$

with parameters expressed as a 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.[1] 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 (Hamilton, 1989). 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. For example, the filtered probability can be expressed as:

$$P(s_t = j \mid Y_t,\ Z_t;\ \widehat{\Theta}), \qquad j = 1,\ 2,...S \text{ and } t = 1, 2, 3...T. \tag{4}$$

A potential shortcoming of model (3) 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 generalized version of the model with state-dependent effects. For example, one may use the following model, which allows for different responses to overstaffing $(Y-Y^*)^+$ and understaffing $(Y-Y^*)^-$ 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 $X > 0$.

$$\begin{aligned}\Delta Y_t = &\sum_{i=1}^{p} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha^+(s_t)(Y-Y^*)^+_{t-1} - \alpha^-(s_t)(Y-Y^*)^-_{t-1} + \\ &\omega^+(s_t)\Delta Z^+_t + \omega^-(s_t)\Delta Z^-_t + \Omega(s_t)\varepsilon_t.\end{aligned} \tag{5}$$

Given the large number of parameters to estimate, (5) obviously requires a relatively large number of observations to provide precise coefficient estimates and conclusive results.

3 Data

The empirical analysis is based on Norwegian seasonally unadjusted quarterly data over the period 1974q1–1996q4. Observations for the period 1997q1–2003q4 are retained for post-sample evaluation of the preferred models. 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.

[Fig. 1 about here.]

The number of persons employed in the manufacturing and construction sector display a 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 unemployment in the remaining sample period, it evolves at relatively high levels. A number of studies argue that Norwegian unemployment experienced a structural break in 1989/90 that led to a shift in its long run mean (e.g. Bianchi and Zoega, 1998; Skalin and Teräsvirta, 2002). Similarly, the downward shift in industry employment in the late 1980s can be interpreted as a shift in the long run mean of employment.

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

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

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 programme ratio (lmp), the share of industry employment in total employment (nis), and the change in real crude oil prices ($\Delta roilp$).

[Fig. 2 about here.]

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 for details. The roles of these variables in the identified labour demand and unemployment equations are discussed in Sections 4.1 and 4.2 below, but even at this stage, a certain partitioning of variables lies close at hand: For example, we would expect that industry labour demand depends on labour costs (represented by ulc) and demand (e.g. $d - k$) and on potential friction effects. Figure 2 suggests a downward trend in the variable 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 decentralization of wage bargaining; and with increases in social welfare programmes. This development may have contributed to a rise in the unemployment rate over time (e.g. Layard et al., 1991). Therefore, we expect it to be important mainly in the "unemployment equation". Crude oil-prices tend to have negative impact on industry employment in most economies. In Norway, however, the effects of oil prices on industry employment are not obvious, partly due to the relatively heavy reliance of Norwegian industry on hydroelectric power system. However, since tax revenues from North Sea oil are particularly important for fiscal policy, crude oil prices may have a positive relationship with the general level of activity. This may emerge as a negative relationship between oil prices and the overall rate of unemployment.

4 A linear model

We estimated a 5th order VAR for $Y = (n, u, h)$ conditional on the vector Z; small letters indicate logs of the variables. The following lags and transformations of the variables in Z seemed to be statistically significant and provided a parsimonious representation of the effects of the Z variables: ulc_{t-1}, nh, $(d - k)_{t-1}$, nis_{t-1}, Δulc_{t-1}, Δulc_{t-2}, Δulc_{t-3}, $\Delta_4 nh$, Δd_t, $\Delta_4 d_t$, $\Delta roilp_{t-1}$ and $\Delta_4 lmp_{t-1}$. In addition, three centred seasonal dummies CS's and two impulse dummies, $i81q1$ and $i86q1$, were included to control for seasonal effects and to remedy violations of the (standard) assumptions about the residuals. Furthermore, a deterministic trend was included to safeguard against invalid inference on the cointegrating rank (Harbo et al., 1998).

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

[Table 1 about here.]

4.1 Cointegration

We next tested for cointegration using the Johansen (1995) procedure, within a system that restricted ulc_{t-1}, nh, nis_{t-1}, $(d - k)_{t-1}$ and the deterministic trend to the cointegration space, while the short run effects of variables (listed above), constant term and the dummy variables were entered unrestricted (Harbo et al., 1998). Table 2 reports 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 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 about here.]

The linear trend was included in the system in order to make the inference on rank more robust. After deciding on $r = 3$, the system was re-estimated and analysed without the deterministic trend; a test showed that the trend could be excluded (the test statistic was $\chi^2(3) = 4.55[0.21]$).

[Table 3 about here.]

Table 3 shows three relationships, which are identified by restricting the β and α vectors. The (over) identifying restrictions are jointly acceptable with

$\chi^2(11) = 14.21$ [0.22]. Figure 3 shows the recursive estimates of the identified 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 1985q1–1996q4.

The identified cointegration vectors can be interpreted as follows. In the first row of $\widehat{\beta}'$, which gives the coefficients of the industry labour demand equation, 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. In the same row, h and n are modelled as perfect substitutes, which is consistent with the "labour sharing view". A rise in ulc reduces employment, consistent with a downward sloping demand curve for labour. The positive coefficient estimate of $(d - k)$ suggests that higher capacity utilization raises employment, or alternatively, a rise in the capital stock (k) substitutes employment. According to the second row of $\widehat{\beta}'$, a reduction in the overall unemployment rate u will follow from a rise in industry employment n, which amounts to only 1/4 of total employment. 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.

[Fig. 3 about here.]

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.[2] 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 the actual and equilibrium level of employment (e.g. Jacobson and Ohlsson, 2000). 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.

4.2 A simultaneous equation model with linear friction effects

In sum, the cointegration analysis implies that $Y - Y^*$ is a 3×1 vector defined as:

$$n - n^* = n - \{0.20(d-k) - h - 0.13(u + ulc)\}, \tag{6}$$
$$u - u^* = u - \{-1.8n - 4.0nis\}, \tag{7}$$
$$h - h^* = h - nh. \tag{8}$$

Using these equilibrium correction terms, the conditional VAR model was reformulated as a (conditional) VECM of order 4 in the differences. Thereafter, parsimony was sought through data-consistent coefficient restrictions. That parsimonious version of the system was reformulated as a simultaneous equations system (cf. Bårdsen and Fisher, 1999). 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. This led to further simplification through an exclusion of some variables. Table 4 presents the preferred final specification of the structural VECM 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 joint insignificance of the 44 over-identifying restrictions shows that the final model parsimoniously encompasses the initial VECM.

The short run effects of the explanatory variables are interpretable. In particular, a rise in unemployment reduces the growth in industry 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 programme ratio, higher oil prices and by a reduction in normal working hours.

[Table 4 about here.]

In this structural VECM, actual working hours act as a buffer against understaffing ($(n - n^*) < 0$) and overstaffing ($(n - n^*) > 0$) 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 a high degree of negative autocorrelation in hours, probably reflecting the pronounced seasonal variation in working hours.

As noted above, the model in Table 4 is consistent with linear friction effects. To investigate whether there are any discernible traits of a non-linear $f(U_{t-1})$, which would imply multiple equilibria, we defined $f(U_{t-1})$ as a logistic function of U_{t-1}, as in equation (1). The value of the threshold parameter (c) was set at 0.04 and that of the steepness-parameter (ξ) at 100; since estimates of c and ξ were found to be quite imprecise when the Maximum Likelihood method 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 0 (high friction) when $U_{t-1} < 0.04$ and close to 1 (low friction) when $U_{t-1} > 0.04$.

Notably, the joint test of the significance of the logistic $f(U_{t-1})$ when added to the employment and the unemployment equations in Table 4 yielded $\chi^2(2) =$

1.60[0.45], 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 1977q1–1996q4. The stability of the parameter estimates defining n^*, u^* and h^* is shown above, in Figure 3.

[Fig. 4 about here.]

[Fig. 5 about here.]

Tests of the overall stability of the structural VECM 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 1985q1–1996q4.

However, one can not firmly conclude that the parameterization of the multi-equation model is unaffected by regime shifts. This is because some parameter changes can remain undetected by a vector constancy test. As shown and discussed in Hendry (2000), even quite large shifts in the parameters representing dynamics, such as adjustment speeds, can be difficult to detect if the parameters defining the long run equilibrium remain unaltered, which seems to be the case here, see Figure 4 (and Figure 3). One explanation is that subject to stability of the cointegrating vectors, even large shifts in the adjustment speeds do not induce mean-shifts in the errors, thus giving rise to low statistical test power; see Hendry (2000) and the references therein.

In the next section, we therefore directly investigate whether the short run parameters of the VECM, characterizing persistence in employment and unemployment and their response to changes in exogenous variables, depend on the phase of the economy. In line with common practice, we assume that a model of hours (h) with state-dependent parameters is not required. 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 (e.g. Hamermesh and Pfann, 1996). The time series of H in Figure 1 lends support to this practice.

5 Modelling with state-dependent adjustment

The employment and unemployment equations in Table 4 were estimated separately with $S = 2$.[3]

The estimation was conducted by Maximum Likelihood (ML) using a version of the Expectation Maximization (EM) algorithm (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 (Krolzig, 1997).

Table 5 presents the outcomes for the employment and the unemployment equations 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 employment and unemployment in Figure

6. Table 6 reports the results of an extensive evaluation of the models.

[Fig. 6 about here.]

[Table 5 about here.]

Figure 6 suggests some differences in the cycles of industry employment and the aggregate unemployment rate. Notably, the series for the unemployment rate clearly indicate two recessionary periods: 1983–1985 and 1988–1996q3; unemployment seems to enter an expansionary phase at the end of 1996. These finding are consistent with those from univariate studies of unemployment (e.g. Akram, 2005). The series for employment largely suggest an expansionary period until 1988 and a recessionary period until about 1994, except for a brief expansionary phase in the period 1989–1990. There are also some signs of a recession in the period 1983–1985, but the filtered and smoothed probabilities for employment are not as clear as the corresponding probabilities for unemployment in classifying the cycle. The probabilities also suggest that industry employment, in contrast to aggregate unemployment, moves out earlier and gradually from the recession in the latter part of the sample.

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; see Table 5. There is also a substantial improvement in the fit of the unemployment equation in the state of recession, but with a slight deterioration in the state of expansion; $\widehat{\sigma}_{u,1}$ is 2.33% and $\widehat{\sigma}_{u,2}$ is 7.71% against $\widehat{\sigma}_u = 6.60\%$; see Table 5.

[Table 6 about here.]

Moreover, the linear models for employment and unemployment can be rejected against the models with state-dependent parameters at the 10% and 1% levels of significance, respectively; see Tables 5 and 6. The diagnostics of both models suggest that the models remain relatively well specified, except for signs of autocorrelation in the residuals from the model of unemployment, see Table 6. This invites a more careful specification of the model within the state-dependent framework. However, to facilitate comparison with results from the linear models, we leave the specification of the state-dependent models equal to those of their linear versions in Table 5.

Furthermore, a post-sample evaluation of both models suggests that the state-dependent models have better forecasting properties than the corresponding linear models. Specifically, means of 1-step prediction errors over the period 1997q1–2003q4 suggest that employment and unemployment growth are underpredicted by 0.4% and 3%, respectively, when using the linear models. The corresponding figures when using the state-dependent models are 0.06% and 0.8%, respectively. The state-dependent models are also preferred by the RMSE-statistics. The ratio of he RMSE of the state dependent models to the linear models is 0.87 and 0.69 for the employment and unemployment models, respectively.

Table 5 shows that employment adjustment is highly state-dependent; it adjusts much faster towards its equilibrium value and is more responsive to shocks in its determinants during a recession than in an expansion, which tends to be characterized by a shortage of labour. In a 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) characterization of employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions.

It should be noted that despite the clear differences in the employment response across the two states, the equilibrium solution of employment remains the same across the two states and close to that found in the case of the linear model. We note that the constant term in the equilibrium solution, i.e., the ratio of the state-dependent intercept to 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 Figure 4.

The above results on short run and long run employment behaviour remained robust, qualitatively, to an extension of the model where we allowed for asymmetric response to positive and negative changes in the regressors, as in equation (5). Asymmetric response may occur if e.g. hiring costs are larger than firing costs, as observed by e.g. Hamermesh and Pfann (1996). Specifically, in each of the two states, we allowed the employment response to vary with over- and understaffing, $(n - n^*)^+_{t-1}$ and $(n - n^*)^-_{t-1}$, and to positive and negative changes in the other regressors, except the autoregressive and deterministic terms; see Section 2.[4]

Interestingly, the results for the unemployment rate suggest that the degree of persistence and speed of adjustment towards equilibrium is largely state-independent. Thus, the equilibrium solution of unemployment is the same across the two states; the derived estimates of the constant terms in the equilibrium solution are $0.659\ /0.177 \approx 0.659\ /0.180 \approx 3.7$, as in the case of the linear model. This adds to the evidence of the stability of the long run mean of $u - u^*$. Also, the effects of (infrequent) institutional changes in working hours ($\Delta_4 nh$) and changes in the share of unemployed enrolled in labour market programmes ($\Delta_4 lmp$) seem to be the same across both states.

Furthermore, unemployment responds more strongly to demand shocks in a tight labour market than in a slack market. In particular, the effects of changes in demand and real oil prices are much stronger in an expansion than in a recession. The relatively sluggish response of unemployment in a slack labour market may be an implication of the "encouraged workers effect"(Pencavel, 1986). In a slack labour market, positive impulses from e.g. oil prices and aggregate demand may 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.

6 Conclusions

The empirical evidence in this paper suggests 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 shortages contribute to labour hoarding and employment persistence. These conclusions have appeared robust to allowance for asymmetric responses to shocks.

The derived equilibrium solutions of 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) characterization of 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 macroeconometric model of 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].

iyyq1: Impulse dummy, 1 in period 19yyq1 and zero elsewhere.

K: Stock of physical capital in manufacturing and construction. Mill. 1991 NOK. [KIBA].

LMP: Number of unemployed on labour market programmes 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/NWPF].

ROILP: Real spot price of Brent Blend crude oil in NOK; derived by deflating the nominal oil price in NOK by the Norwegian CPI.

U: Total unemployment rate as a fraction of total labour force. [UTOT2].

ULC: Unit labour costs (including payroll tax) in manufacturing and construction deflated by the producer price index. 1991 NOK. [WCIBA/PYIBA $\times$ ZYIBA].

[Table 7 about here.]

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Notes

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

[2]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.

[3]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 shifts led to failure of convergence for both $S = 3$ and $S = 2$.

Furthermore, the choice of modelling the two equations separately owes to lack of technology for estimating Markov switching models with contemporaneous explanatory variables. However, a comparison of FIML estimates in Table 4 with the corresponding OLS estimates showed virtually no differences between the coefficient estimates, indicating that potential simultaneity bias is likely to be negligible.

[4]Due to the large numbers of parameters to be estimated, many coefficient estimates became more uncertain, providing inconclusive results. Apparently, the coefficient estimates offered mixed support for an asymmetric response to positive and negative shocks. However, we found clear indications of state-dependent employment behaviour. The explanatory variables generally had a bigger impact on employment in a recession than in an expansion. The details are available from the authors on request.

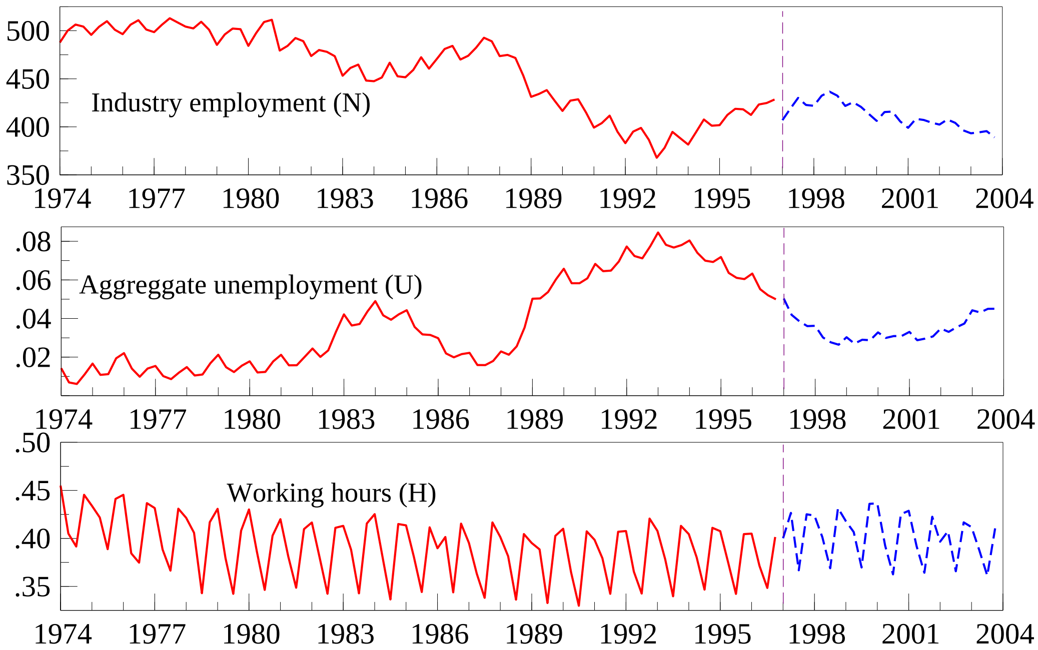


Fig. 1. *Time series of the Y variables (in levels) over the (effective) full sample period: 1974q1–2003q4. These are persons employed in manufacturing and construction in thousands (N), the aggregate unemployment rate (U) and average working hours in manufacturing and construction (H) in thousands. Solid lines represent the estimation period, while the dashed lines mark the post-sample period: 1997q1–2003q4.*

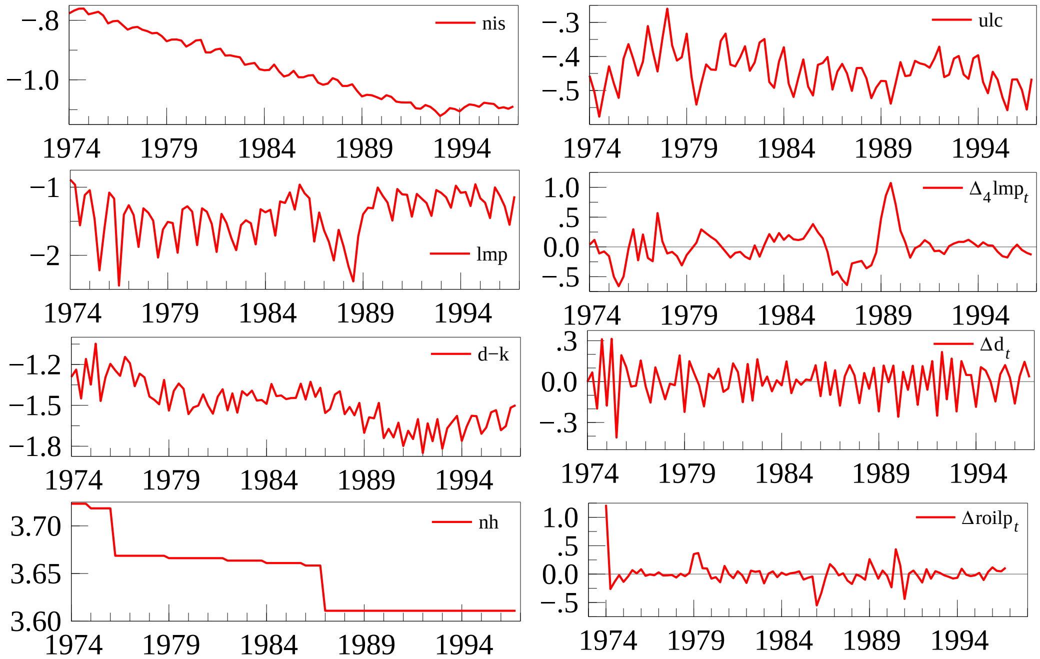


Fig. 2. *Time series of the Z variables and their transformations over the period 1974q1–1996q4. From left (in logs): Share of industry employment in total employment (nis), unit labour costs (ulc), the programme ratio (lmp) and annual growth in the programme ratio ($\Delta_4 lmp$), indicator of capacity utilization ($d-k$), quarterly growth in aggregate demand (Δd), normal working hours (nh) and finally, quarterly growth in real crude oil prices ($\Delta roilp$).*

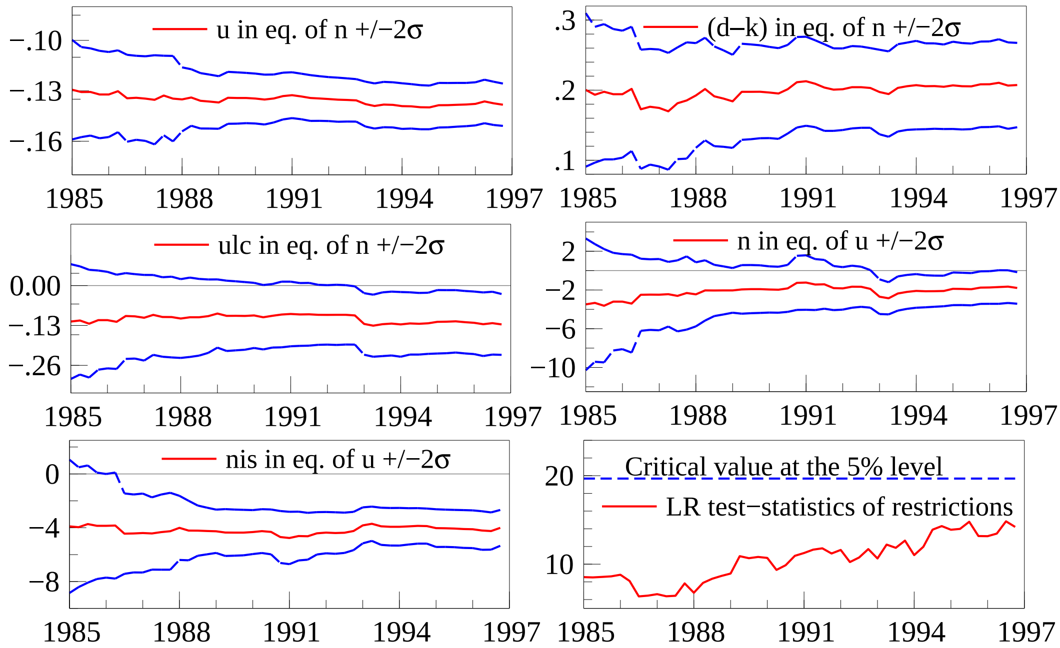
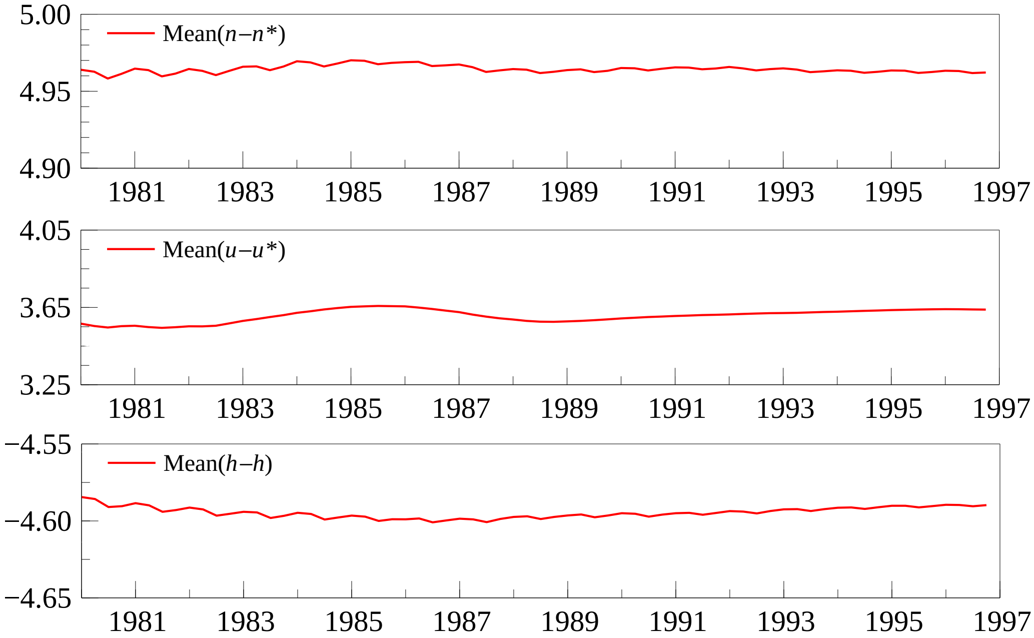


Fig. 3. *Recursive estimates of the cointegration vectors with +/-2SE and of the LR test-statistics for 11 joint restrictions on the* β *and* α *matrices. Initial sample: 1974q1–1984q4.*



Mean(n−n*)
5.00
4.95
4.90
1981
1983
1985
1987
1989
1991
1993
1995
1997
Mean(u−u*)
4.05
3.65
3.25
Mean(h−h)
−4.55
−4.60
−4.65

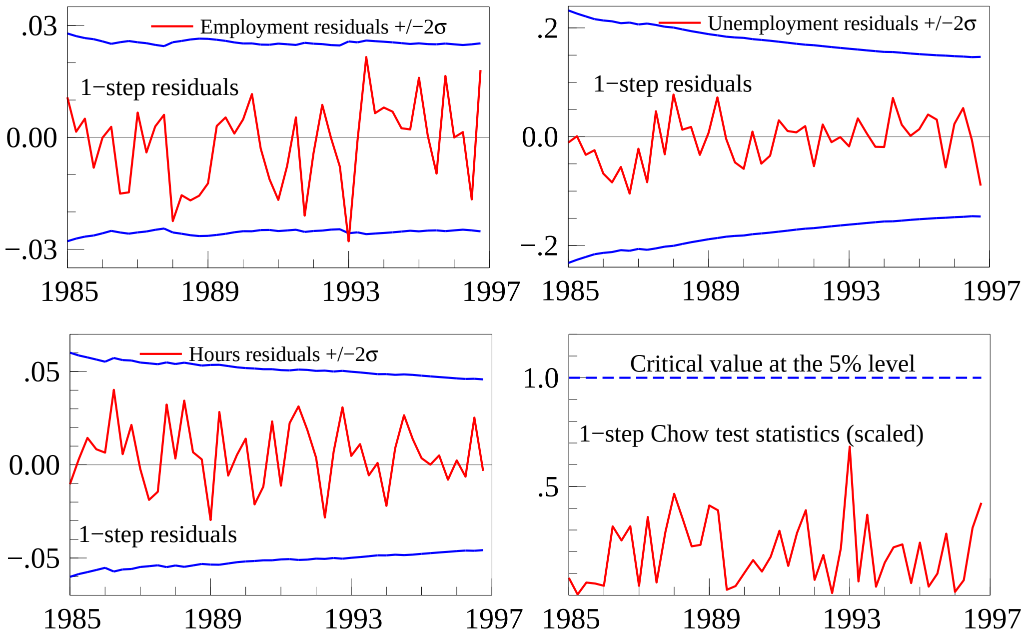


Fig. 5. *1-step ahead residuals ± 2 estimated standard errors based on the equations of employment, unemployment and hours. Also, a sequence of 1-step Chow tests scaled by their critical values at the 5% level of significance.*

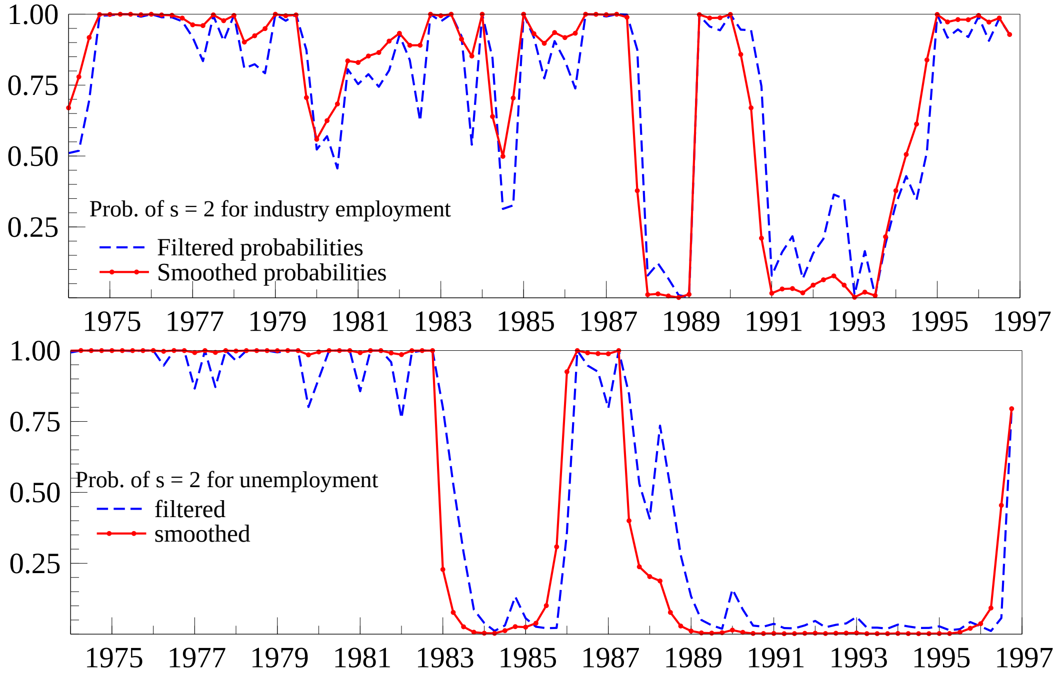


Fig. 6. *The filtered and smoothed probabilities of industry employment and aggregate unemployment being in state 2: the expansion phase.*

Table 1
Diagnostics for the conditional VAR model

Equation for:	n	u	h	VAR
AR: $F_{ar,\ 1-5}(5,53)$	0.91[0.48]	1.97[0.10]	2.08[0.09]	
ARCH: $F_{arch,\ 1-4}(4,50)$	0.21[0.93]	0.90[0.47]	0.65[0.63]	
Heterosced.: $\mathrm{F}_{het}(40,17)$	0.26[1.00]	0.75[0.77]	0.41[0.99]	
Normality: χ^2_{nd}	0.83[0.66]	4.56[0.10]	1.08[0.58]	
AR: $\mathrm{F}^v_{ar,\ 1-5}(45,122)$				1.47[0.05]
Heterosced.: $F^v_{het}(240,80)$				0.41[1.00]
Normality: $\chi^{2,v}_{nd}(6)$				5.90[0.43]

Note: The 5th order conditional VAR model is estimated on data for the period 1974q1–1996q4. We refer to Hendry and Doornik (1996) and references therein for details about the tests. The last column reports the results for the whole system. The square brackets contain p-values here and elsewhere in this paper.

Table 2
Cointegration rank.

r	1	2	3
eigenvalue	0.46	0.22	0.12
Tr	90.68	34.06	11.35
95%	69.7	44.5	20.7

Table 3
Restricted cointegration analysis, identification of cointegration vectors

$\widehat{\beta}'$	n	u	h	ulc_1	nh	$(d-k)\_1$	nis_1
1	-1	-0.138	-1	-0.126	0	0.207	0
		(0.006)		(0.050)		(0.030)	
2	-1.797	-1	0	0	0	0	-4.008
	(0.814)						(0.665)
3	0	0	-1	0	1	0	0

,

$\widehat{\alpha}$	1	2	3
n	0.417	-0.038	0
	(0.104)	(0.01)	
u	1.961	0.033	0
	(0.631)	(0.058)	
h	0	0	0.263
			(0.124)

Table 4
A simultaneous equation model with linear friction effects

Industry employment

$$\widehat{\Delta n_t} = \underset{(0.141)}{0.555} - \underset{(0.010)}{0.032}\,\Delta u_t - \underset{(0.009)}{0.028}\,\Delta u_{t-2} - \underset{(0.026)}{0.155}\,\Delta h_t$$

$$+ \underset{(0.089)}{0.176}\,\Delta n_{t-4} - \underset{(0.028)}{0.112}\,(n-n^*)_{t-1} + \underset{(0.011)}{0.033}\,\Delta d_t$$

$$- \underset{(0.011)}{0.032}\,i81q1_t + \underset{(0.011)}{0.022}\,i86q1_t + \underset{(0.005)}{0.012}\,CS_{t-1}$$

$$\hat{\sigma}_n = 1.088\%$$

Aggregate unemployment

$$\widehat{\Delta u_t} = \underset{(0.126)}{0.678} - \underset{(0.386)}{0.547}\,\Delta_3 n_t + \underset{(0.063)}{0.342}\,\Delta u_{t-1} - \underset{(0.070)}{0.144}\,\Delta u_{t-2}$$

$$+ \underset{(0.050)}{0.177}\,\Delta u_{t-3} + \underset{(0.072)}{0.460}\,\Delta u_{t-4} - \underset{(0.035)}{0.183}\,(u-u^*)_{t-1}$$

$$- \underset{(0.038)}{0.138}\,\Delta roilp_{t-1} - \underset{(0.075)}{0.204}\,\Delta_4 d_t + \underset{(0.536)}{2.08}\,\Delta_4 nh_t$$

$$- \underset{(0.027)}{0.113}\,\Delta_4 lmp_{t-1} - \underset{(0.026)}{0.146}\,CS_{t-1}$$

$$\hat{\sigma}_u = 6.598\%$$

Industry hours

$$\widehat{\Delta h_t} = \underset{(0.173)}{0.499}\,\Delta_4 nh_t - \underset{(0.152)}{0.633}\,\Delta h_{t-1} - \underset{(0.149)}{0.656}\,\Delta h_{t-2} - \underset{(0.124)}{0.545}\,\Delta h_{t-3}$$

$$- \underset{(0.080)}{0.201}\,\Delta h_{t-4} + \underset{(0.022)}{0.032}\,\Delta d_t - \underset{(0.161)}{0.260}\,\Delta n_{t-4} - \underset{(0.081)}{0.258}\,(h-h^*)_{t-1}$$

$$- \underset{(0.075)}{0.239}\,(n-n^*)_{t-1} - \underset{(0.021)}{0.065}\,i86q1_t - \underset{(0.022)}{0.028}\,CS_{t-1}$$

$$- \underset{(0.023)}{0.112}\,CS_{t-2}$$

$$\hat{\sigma}_h = 2.066\%$$

Diagnostics

$$\begin{array}{lcl} AR\ F^{v}_{ar,\,1-5}(45,190) & = & 1.20[0.20] \\ Normality\ \chi^2(6) & = & 3.59[0.73] \\ Heterosced.{:}\ F^{v}_{het}(270,187) & = & 1.04[0.39] \\ Over\text{-}identification\ \chi^2(44) & = & 58.13[0.08] \end{array}$$

Note: FIML estimates. The sample is 1974q1–1996q4. Standard errors are reported in parentheses below the coefficient estimates. Square brackets include the p-values.

Table 5
Models with state-dependent parameters

Industry employment

Variables	Recession (s = 1) Coeff	Std error	Expansion (s =2) Coeff	Std error
$Const$	1.410	(0.298)	0.414	(0.130)
Δu_t	−0.062	(0.016)	−0.016	(0.008)
Δu_{t-2}	−0.071	(0.020)	−0.015	(0.008)
Δh_t	−0.307	(0.041)	−0.123	(0.024)
Δn_{t-4}	−0.078	(0.148)	0.209	(0.084)
$(n - n^*)_{t-1}$	−0.285	(0.060)	−0.083	(0.026)
Δd_t	0.049	(0.023)	0.022	(0.010)
$i81q1_t$	−0.007	(0.019)	−0.039	(0.010)
$i86q1_t$	0.020	(0.025)	0.021	(0.009)
CS_{t-1}	0.021	(0.010)	0.012	(0.005)
$\hat{\sigma}_n$	0.698%		0.846%	

Aggregate unemployment

	Recession (s = 1) Coeff	Std error	Expansion (s =2) Coeff	Std error
$Const$	0.659	(0.093)	0.659	(0.093)
$\Delta_3 n_t$	−0.441	(0.183)	−0.393	(0.713)
Δu_{t-1}	0.341	(0.051)	0.394	(0.095)
Δu_{t-2}	−0.010	(0.045)	−0.147	(0.100)
Δu_{t-3}	0.103	(0.038)	0.214	(0.080)
Δu_{t-4}	0.485	(0.072)	0.465	(0.102)
$(u - u^*)_{t-1}$	−0.177	(0.024)	−0.180	(0.027)
$\Delta roilp_{t-1}$	−0.058	(0.030)	−0.117	(0.055)
$\Delta_4 d_t$	0.046	(0.069)	−0.274	(0.101)
$\Delta_4 nh_t$	1.719	(0.661)	1.774	(0.727)
$\Delta_4 lmp_{t-1}$	−0.121	(0.020)	−0.104	(0.049)
CS_{t-1}	−0.117	(0.018)	−0.197	(0.046)
$\hat{\sigma}_u$	2.335%		7.709%	

Note: The sample is 1974q1–1996q4. Asymptotic standard errors are reported in parentheses beside the coefficient estimates. The EM algorithm has been employed in the estimation using MSVAR 0.99 for Ox 2.10 (Krolzig, 1998).

Table 6
Evaluation of the state-dependent models

Diagnostics

Tests	MS-ECM of Δn		MS-ECM of Δu	
LR linearity test	$\chi^2(11)$:	17.44[0.09]	$\chi^2(12)$:	47.48[0.00]
AR portm. (12)	$\chi^2(8)$:	15.04[0.06]	$\chi^2(8)$:	21.39[0.01]
Normality	$\chi^2(2)$:	1.81[0.41]	$\chi^2(2)$:	0.323[0.85]
Heterosced.	$\chi^2(15)$:	10.45[0.79]	$\chi^2(21)$:	24.55[0.27]

Post-sample evaluation (1997q1–2003q4)

	Mean of pred. errors	RMSE	
ECM of Δn	−0.004	0.86%	
MS-ECM of Δn	−0.0006	0.75%	Ratio: 0.87
ECM of Δu	−0.030	5.17%	
MS-ECM of Δu	−0.008	3.59%	Ratio: 0.69

Note: p-values are reported in brackets. *MS-ECM* refers to the Markov switching models in Table 5, while *ECM* refers to their linear versions; see Table 4. Due to a break in the time series for employment (n), an impulse dummy for 1997q1 was included in the equations of Δn; see Figure 1.

Table 7
ADF tests of unit roots; 1974q1–1996q4

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
$\Delta roilp$	0.022	-9.270^{**}	2
$roilp$	0.918	-1.980	3

Note: Initially, 12 lags were allowed for in each of the ADF-models, which contained a constant when testing for a unit root in the first difference of a variable. Both a constant and a trend were included when we tested for unit root 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 a trend: –3.46; when a constant, the 5% and the 1% DF-values are –2.89 and –3.50. One raised star indicates rejection of the null hypothesis at the 5% level while two raised stars indicate rejection at the 1% level. Sample 1974q1–1996q4.