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Multiple unemployment equilibria and asymmetric dynamics—Norwegian evidence☆

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Abstract

A high degree of persistence in European unemployment rates is often termed unemployment hysteresis, which is equated with the presence of a unit root in unemployment series using linear models. However, the implications of this practice are difficult to reconcile with existing evidence on long run unemployment behaviour and the asymmetric response of unemployment to positive and negative shocks. This paper investigates whether the persistence and asymmetric response to positive and negative shocks may be due to the existence of multiple equilibria. Using Norwegian unemployment data, which share important characteristics with other European unemployment series, this paper characterises unemployment behaviour by applying a logistic smooth transition autoregressive (LSTAR) model. It appears that unemployment is a stationary variable that has switched between two stable equilibria during the period 1972–2003. The paper demonstrates that a large shock, or a sequence of small shocks, may cause a transition from one equilibrium level to another and that unemployment responds asymmetrically to large positive and negative shocks, but symmetrically to small shocks.

JEL classification: E24; C22; C52

Keywords: Hysteresis; Multiple equilibria; Non-linear model; Asymmetric adjustment

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☆ The views expressed in this paper are those of the author and should not be interpreted as reflecting those of Norges Bank (the central bank of Norway).

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1. Introduction

In the last few decades, (western) European unemployment has displayed a weak tendency, if any, to revert to a unique stable equilibrium rate as prescribed by the natural rate or the NAIRU hypotheses (e.g. Ljungqvist and Sargent, 1998). The hypothesis of unemployment hysteresis has been launched to account for this apparently ratcheting behaviour of unemployment (e.g. Blanchard and Summers, 1986). Hysteresis denotes a situation in which the equilibrium state of a system depends on the past history of the system, in a linear or a non-linear way (e.g. Amable et al., 1995; Røed, 1997). Since the seminal work of Blanchard and Summers (1986), however, tests of unemployment hysteresis have been conducted as unit root tests in linear dynamic models of unemployment (e.g. Cross, 1995).

The practice of equating unemployment hysteresis with the presence of a unit root in unemployment using linear models may be questioned for a number of reasons (e.g. Amable et al., 1995; Cross, 1995; Røed, 1997). First, it compels every shock, irrespective of its size, sign, and persistence, to have a permanent effect on the level of unemployment, disregarding the existence of (endogenous) stabilising mechanisms. Second, long data series often show that unemployment does not wander around randomly, but reverts to its past levels sooner or later. Indeed, the bounded nature of the unemployment rate series prevents it from taking values outside the 0–1 range. Third, empirical evidence in favour of a unit root, or a high degree of persistence, in relatively small samples may be due to large shocks to the series. It is well known that standard unit root tests underreject the null hypothesis of unit root when there are breaks in a time series (e.g. Perron, 1989).

Furthermore, linear models imply symmetric response of unemployment to positive and negative shocks. However, a number of studies suggest that it may exhibit asymmetric dynamic response to positive and negative shocks. Such behaviour is often ascribed to asymmetric employment adjustment costs (e.g. Hamermesh and Pfann, 1996). Accordingly, one may observe that unemployment rises faster than it falls if firing costs are smaller than hiring costs, and the opposite if firing costs are larger than hiring costs. Moreover, the speed of unemployment adjustment may depend on the rate of unemployment, if, e.g. hiring costs decline when unemployment rises, as found by, e.g. Burgess (1992) and Palm and Pfann (1997).

Models of multiple equilibria seem to be capable of reconciling the empirical evidence from long and short unemployment time series and offer an additional explanation of asymmetries in unemployment behaviour. In these models, large shocks can shift the unemployment from one self-sustaining level, i.e. equilibrium level, to another, while small shocks just cause temporary deviations from a given equilibrium level (e.g. Diamond, 1982; Cooper and John, 1985; Murphy et al., 1989; Weitzman, 1982).[1] In addition, unemployment may respond asymmetrically to positive and negative shocks, since the sign of a shock determines whether it becomes attracted towards, e.g. both of the equilibria or just the initial one. In the former case it is attracted in opposite directions, in which case the adjustment towards the initial equilibrium may become more sluggish, if any, than in the latter case.

[1] A number of studies associate the presence of multiple equilibria with hysteresis, which is interpreted as a non-linear phenomenon (e.g. Røed, 1997, and the references therein). Accordingly, the multiple equilibria approach differs from the standard hysteresis approach only in the sense of regarding hysteresis as a non-linear phenomenon.

This also suggests that, if there are multiple equilibria, the unemployment response to a positive or a negative shock may depend on the rate of unemployment.

This paper tests for the presence of multiple equilibria and the possibility of asymmetric response to positive and negative shocks in Norwegian unemployment, which is similar to other European unemployment series in some aspects. In particular, formal tests using linear models suggest that it contains a unit root. The paper contrasts the merits of the multiple equilibria approach against those of the unique equilibrium and the standard hysteresis approaches in characterising unemployment. For this purpose, it employs the smooth transition autoregressive (STAR) model proposed by Teräsvirta (1994), in which models of unique equilibrium and hysteresis can be derived as special cases. The STAR model also allows one to characterise abrupt transitions between potential equilibria. Moreover, this model offers a convenient way to test for multiple equilibria and investigate the dynamic behaviour of unemployment.

Previously, Skalin and Teräsvirta (2002) have studied Norwegian unemployment using a smooth transition model, but with the *time* trend as a transition variable to allow for changes in the unemployment equilibrium over time. Their findings suggested a permanent autonomous shift to a higher equilibrium level of unemployment in 1988. This date fits well with the actual behaviour of Norwegian unemployment. However, their model precluded a reversion towards a lower equilibrium level at a later time period. This contrasts with a characteristic feature of theory models of multiple equilibria, which allow for back and forth movements between equilibria depending on the size and sign of shocks.[2]

Moreover, policy implications of a model that allows for a permanent autonomous transition to another equilibrium level differ greatly in the long run from a model that allows for back and forth movements between equilibria. The former model implies that, e.g. monetary policy is neutral in the long run. In contrast, the latter model implies that monetary policy can be non-neutral in the long run, albeit within a certain range of the unemployment rate. This is because large monetary policy shocks of appropriate sign can shift the unemployment rate between a finite number of equilibria. In this sense, the policy implications of the latter model are a compromise between the implications of natural rate models and linear hysteresis models, where every shock has a permanent effect on unemployment.

The paper is organised as follows. Section 2 formalises the multiple equilibria approach within the framework of a STAR model. Section 3 presents the unemployment data and highlights its features. This section also characterises the unemployment series using a linear AR model and points out its inadequacy in capturing the behaviour of unemployment over time. Section 4 presents results based on the STAR model and elaborates on the behaviour of unemployment out of disequilibria, especially when exposed to positive and negative shocks of different sizes. In addition, this section undertakes a comprehensive post-sample evaluation of the derived model. Section 5 presents our conclusions.

[2] In addition, the model in Skalin and Teräsvirta (2002) does not seem to provide reasonable estimates of the equilibria in Norwegian unemployment. Taken at face value, the proposed model implies the equilibrium level to be $0.0012/0.22 \approx 0.5\%$ before 1988 and $(0.0012 + 0.64)/0.22 \approx 2.9\%$ afterwards.

2. Formalising multiple unemployment equilibria

Multiple unemployment equilibria may be associated with high and low levels of self-sustaining activity and employment levels. They may arise when the activity level of one agent depends positively on the activity level of another agent and vice versa. Such reciprocal externalities can arise from trading and exchange opportunities and or due to spillovers of demand across markets.

For example, a positive demand shock may increase trading opportunities and incentives to produce. The resulting higher production level may become self-sustaining through the positive feedback between production level and trading opportunities (Diamond, 1982; Cooper and John, 1985). The high production level may also be supported or boosted by the implied rise in labour income, which may further raise aggregate demand and profits. Thus it may become profitable for more firms to operate and demand labour (Weitzman, 1982). In addition, the increase in trading opportunities can make it profitable for firms to replace constant or decreasing returns technologies with increasing returns technologies supporting a high level of activity and employment (e.g. Murphy et al., 1989).[3]

The following subsection formalises the multiple equilibria approach within the framework of a STAR model. Its also reviews a test of the linear model against the STAR model, see e.g. Teräsvirta (1994) for more details.

2.1. Smooth transition autoregressive (STAR) model

A STAR model of unemployment can be formulated as follows:

$$U_t = \alpha + \sum_{i=1}^{q} \phi_i U_{t-i} + (\tilde{\alpha} + \sum_{i=1}^{q} \tilde{\phi}_i U_{t-i}) F(U_{t-d}) + \varepsilon_t, \quad \varepsilon_t \sim IIDN(0, \sigma^2) \tag{1}$$

where q denotes the number of lags. $F(U_{t-d})$ is a transition function that increases monotonically with the level of unemployment, lagged d periods, and takes on values in the 0–1 range. The transition function represents the phase or state of the economy. It can be specified as a logistic function:

$$F(U_{t-d}) = (1 + \exp[-\gamma\{U_{t-d} - c\}])^{-1}, \qquad \gamma > 0. \tag{2}$$

Here the transition parameter γ determines the speed of transition from 0 to 1, for a given deviation of U_{t-d} from a constant threshold value c. This logistic STAR (LSTAR) model allows both the intercept term and the autoregressive coefficients to change with the value of $F(U_{t-d})$. Thus unemployment can evolve around distinct rates with different dynamics in expansions $U_{t-d} < c$ and contractions $U_{t-d} > c$. The change in parameters occurs with some delay d, which can be determined together with the tests of a linear AR(q) model against a STAR model, see below.

Two unemployment equilibria correspond to values of $F(U_{t-d}) = 0$ and $F(U_{t-d}) = 1$ and may be defined as $\mu_1 = \alpha/(1 - \sum_{i=1}^{q} \phi_i)$ and $\mu_2 = (\alpha + \tilde{\alpha})/[1 - \sum_{i=1}^{q}(\phi_i + \tilde{\phi}_i)]$, re-

[3] A large number of other mechanisms can also lead to multiple equilibria, including costs associated with layoffs and hirings as in Saint-Paul (1995) and Moene et al. (1997), respectively.

spectively. The sums $\varrho_1 \equiv \sum_{i=1}^{q} \phi_i$ and $\varrho_2 \equiv \sum_{i=1}^{q}(\phi_i + \tilde{\phi}_i)$ measure the degree of persistence in each of the two equilibrium states.

Both a two regime autoregressive model with abrupt transitions and the linear AR(q) model are nested in this LSTAR model. The LSTAR model reduces to a self-exciting threshold autoregressive (SETAR) model with threshold value c if γ is extremely large. Then $F(U_{t-d}) = 0$ for $U_{t-d} \leq c$ but $F(U_{t-d}) = 1$ for $U_{t-d} > c$. The LSTAR model is reduced to an AR(q) model if $\gamma = 0$, since $F(U_{t-d}) = 1/2$ for all values of U_{t-d}. In this case, one can test the null hypothesis of unemployment hysteresis in the unit root sense against the alternative hypothesis of reversion towards a unique equilibrium. If we denote the measure of persistence as ϱ in the case of an AR model, unemployment hysteresis in the unit root sense corresponds to $\varrho = 1$.

However, the parameters defining the LSTAR model are not identified under the null hypothesis of an AR model. It follows from (1) and (2) that under the null hypothesis, i.e. under H_0: $\gamma = 0$, the $\tilde{\alpha}$, $\tilde{\phi}_i$s' and c may take on any value without changing the value of the likelihood function. These parameters are only identified under the alternative hypothesis of $\gamma \neq 0$. To deal with such cases, Teräsvirta (1994) suggests a sequence of tests to evaluate the null hypothesis of an AR model against the alternative of a STAR model. The tests are based on estimating the following auxiliary regression for a chosen value of the delay parameter d:

$$U_t = \beta_0 + \sum_{i=1}^{q} \beta_{1i} U_{t-i} + \sum_{i=1}^{q} \beta_{2i} U_{t-i} U_{t-d} + \sum_{i=1}^{q} \beta_{3i} U_{t-i} U_{t-d}^2 + \sum_{i=1}^{q} \beta_{4i} U_{t-i} U_{t-d}^3 + \nu_t, \quad (3)$$

where ν_t is the error term. The test of an AR(q) model against a STAR model (LSTAR or ESTAR) is equivalent to conducting a joint test of [4]

$$H_0 : \beta_{2i} = \beta_{3i} = \beta_{4i} = 0, \qquad i = 1, 2, \ldots q.$$

The value of d can be determined by conducting this test for different values of d in the range $1 \leq d \leq q$. If linearity is rejected for more than one value of d, then the value which causes the strongest rejection of the null is chosen, that is, the value corresponding to the lowest p-value of the joint test. If the AR(q) model is rejected, one needs to test the appropriateness of an LSTAR formulation against an ESTAR formulation. For this purpose, the following sequence of tests within the auxiliary regression is suggested:

$$H_{04} : \beta_{4i} = 0, \qquad i = 1, 2, \ldots q,$$
$$H_{03} : \beta_{3i} = 0 | \beta_{4i} = 0, \qquad i = 1, 2, \ldots q,$$
$$H_{02} : \beta_{2i} = 0 | \beta_{3i} = \beta_{4i} = 0, \qquad i = 1, 2, \ldots q.$$

[4] An exponential STAR (ESTAR) model is defined by the following exponential transition function $F(U_{t-d}) = 1 - \exp(-\gamma(U_{t-d} - c)^2)$ (e.g. Teräsvirta, 1994).

An LSTAR model is chosen if H_{04} or H_{02} is rejected for at least one value of i and an ESTAR model is chosen if H_{03} is rejected for at least one i.

3. Data and a linear AR(q) model

This section offers some informal evidence against the unique equilibrium approach. It is organised as follows. Section 3.1 describes the unemployment data and plots its density to check whether its shape is consistent with multiple equilibria or not. This subsection also discusses whether unemployment rises faster than it falls. Section 3.2 derives an AR model and explores its properties. A number of parameter constancy tests are also conducted to test for parameter changes over time. These tests can be helpful in dating possible changes in the parameters and indicate the timing of possible transitions between equilibria. The AR model also serves as a reference model for the non-linear model to be estimated later.

3.1. Data

The empirical analysis is based on quarterly data for the (open) unemployment rate in Norway, see Fig. 1.[5] We develop our models using the same observations for the period 1972:1–1997:1 as Skalin and Teräsvirta (2002). However, we extend the data set with observations for the period 1997:2–2004:1 to undertake a comprehensive post-sample evaluation of our preferred model.

Fig. 1 shows that during the period 1972:1–1997:1, there are three noticeable upswings in unemployment relative to its low level in the early 1970s: in 1975, 1982/1983 and 1988. It moves down relatively slowly to its initial levels after the first two upswings except after the third upswing, which is the most striking one. Unemployment continues to rise after this strong upswing in 1988, but at a slower pace until 1992/1993. Thereafter it moves slowly downward, but is still close to its 1988-level, at the beginning of 1997. Extension of the data set shows that unemployment continues its downward movement until 1999 but moves upward thereafter, see the dashed line in Fig. 1.

Visually, the unemployment behaviour after the (first) three upswings indicates that unemployment rises faster than it falls. However, this impression is based mainly on upswings from relatively low levels of unemployment. If we consider the high unemployment level of 1988 as our reference point, subsequent upward movements around this level do not seem to be faster than downward movements. In addition, the impression of asymmetric adjustment is based primarily on the unemployment behaviour in the aftermath of relatively large upswings. A glance at the small movements in the unemployment rate does not capture obvious asymmetries in adjustment, even around the relatively low unemployment rates prior to 1988. Furthermore, we lack observations of marked downswings in the unemployment rate and of the unemployment behaviour afterwards. Thus it is not apparent whether upward

[5] The frequency and content of the Norwegian Labour Force Survey (LFS) was changed in 1996, see Section 4.3 for details. We employ the official data, which has not been adjusted to account for the change in LFS, and test explicitly for the effects of this change through dummy variables. We are grateful to an anonymous referee for pointing out the change in LFS in 1996 to us.

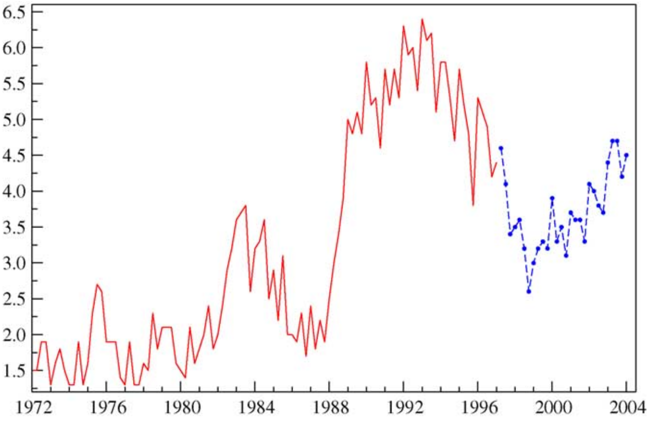

Fig. 1. Quarterly series of Norwegian unemployment in the period 1972:1–2004:1. The circled dashed line represents observations for the period 1997:2–2004:1. The unemployment figures are seasonally non-adjusted and based on the Norwegian labour force survey. *Source*: OECD Main Economic Indicators.

unemployment adjustment after a fall would be faster than the downward adjustment after a rise.

The observed unemployment behaviour may be interpreted in the light of the multiple equilibria approach. Fig. 1 indicates that unemployment evolved between two different equilibria in the sample. This impression is supported by the apparently bimodal frequency distribution of unemployment, with the modes centered at around 2 and 5%, respectively, see Fig. 2. Most of the observations in the period up to 1982 seem to have been generated by a low-equilibrium regime, while the observations for the period 1988–1997 are more likely to have been generated by a high-equilibrium regime. It is however not obvious whether the observations in the period 1982–1985 should be considered as extreme observations from a low- or a high-unemployment regime. Alternatively, they could have been generated from

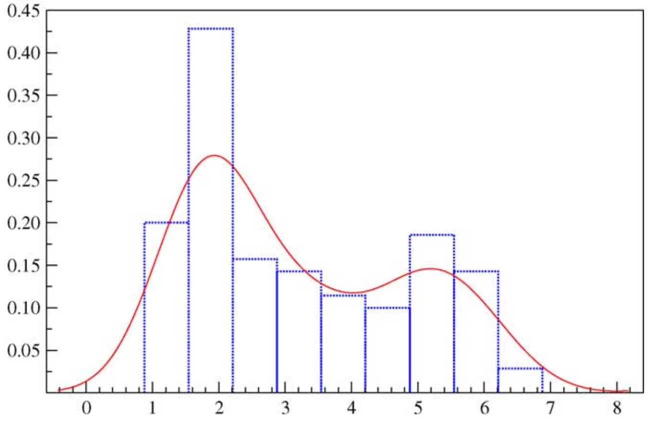

Fig. 2. Non-parametric density estimation of Norwegian unemployment (with histogram); 1972:1–1997:1.

Table 1
An AR(5) model of the Norwegian unemployment rate

$$\hat{U}_t = \underset{(0.099)}{0.17} + \underset{(0.092)}{0.76}\,U_{t-1} + \underset{(0.104)}{0.19}\,U_{t-2} - \underset{(0.105)}{0.12}U_{t-3} + \underset{(0.103)}{0.61}\,U_{t-4} - \underset{(0.093)}{0.48}U_{t-5}$$

$\hat{\varrho} \approx 0.96$, $t-\text{ADF} = -1.59$; DF-critical value at 5% $= -2.89$

Effective sample: 1973:2–1997:1

Diagnostics	
Log-likelihood value	35.8
Standard error of residuals, $\hat{\sigma}$	0.44
Autocorrelation 1–5, $F_{\mathrm{ar},1-5}$ (5,82)	2.85 [0.02]
ARCH 4, $F_{\mathrm{arch},1-4}$ (4,79)	0.66 [0.62]
Normality, χ^2_{nd} (2)	0.57 [0.75]
Heteroscedasticity, F_{XiXj} (20,69)	0.86 [0.64]
Heteroscedasticity, F_{Xi^2} (10, 79)	0.90 [0.54]
Model specification, RESET F(1, 89)	1.73 [0.19]

Note: The standard errors are in parentheses below the estimates and *p*-values are shown in square brackets. $F_{\mathrm{ar},1-5}$ (d.f.1, d.f.2) tests for autocorrelation in residuals up to five lags. d.f.1 and d.f.2 denote degrees of freedom. $F_{\mathrm{arch},1-4}$ (d.f.1, d.f.2) tests for autoregressive conditional heteroscedasticity (ARCH) up to order 4 (Engle, 1982). The normality test with chi-square distribution is that by Jarque and Bera (1980). F_{XiXj} (d.f.1, d.f.2) and F_{Xi^2} (d.f.1, d.f.2) are tests for residual heteroscedasticity due to omission of cross products of regressors and/or squares of regressors (White, 1980). RESET F(d.f.1, d.f.2) is the standard regression specification test (Ramsey, 1969). The results in this table are based on the implementation of these tests in PcGive 9.10 (Hendry and Doornik, 1996).

a third unemployment regime, with a mode between 2 and 5%. This also applies to many observations from the post-sample period 1997:2–2004:1.

3.2. An AR(5) model

Table 1 presents an estimated AR model with five lags. Starting with eight lags, the autoregressive order of five is determined by excluding the statistically insignificant lags of a higher order. The lag order of five is also preferred by Akaike's information criterion (AIC).

The diagnostic test statistics of the model indicate that, except for signs of autocorrelation in the residuals, there are no significant violations of the standard assumptions about residuals. The autocorrelation remained a feature of the residuals even when a lag length of eight was used. A significant value for the autocorrelation test may result when a linear functional form is imposed on a data-generating mechanism that is more properly characterised by non-linear functional forms. The regression specification test (RESET), however, does not indicate significant functional form misspecification, at least not at standard levels of significance. However, this test is constructed to have power against general forms of functional misspecification and may thus have low power against specific non-linear forms. We have tested for the effects of change in the Labour Force Survey in 1996 by adding a step dummy with a value of 1 from 1996 and onwards to the AR(5) model. However, it turned out to be numerically and statistically insignificant. The coefficient estimate was 0.07 and *t*-value equal to 0.35, which implied a *p*-value of 0.72.

The linear model implies unemployment hysteresis in the unit root sense. The estimates of the five autoregressive coefficients sums up to 0.96 ($\approx\hat{\varrho}$) and the associated *t*-ADF value

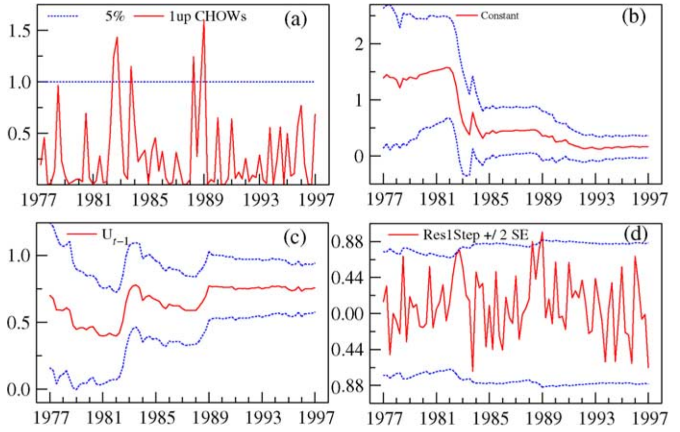

Fig. 3. (a) One-step ahead Chow tests with critical values at the 5%, (b) recursive OLS estimates of the constant term, (c) the first autoregressive coefficient and (d) the one-step ahead residuals, respectively. The recursive coefficient estimates and the one-step ahead residuals are presented with ±2S.E. Initial estimation period is 1973:2–1977:1.

becomes −1.59. Thus the null hypothesis of $\varrho = 1$ is not rejected at the 5% critical value of −2.89 in an Augmented Dickey–Fuller (ADF) test. Such outcomes are common when unit root tests are applied to European unemployment series. However, such outcomes can be due to model misspecifications and non-constant parameters. It is well known that non-constant parameters may reduce the power of an ADF test (e.g. Perron, 1989).

Fig. 3 indicates non-constancies in the model's parameters over time. One-step ahead Chow tests point to noticeable changes in the parameters in 1982/1983 and in 1988, see Fig. 3(a). These may be associated with non-constancies in the intercept term and the autoregressive coefficients at these dates, as indicated by the recursive OLS estimates of the intercept term and the first autoregressive coefficient in Fig. 3(b)–(c).

4. A smooth transition autoregressive model

This section analyses Norwegian unemployment within the STAR framework described in Section 2. Section 4.1 tests the AR(5) model against a STAR model and makes inference on the form of nonlinearity, i.e. whether to use an ESTAR or an LSTAR model. It turns out that the tests favour an LSTAR model, which is derived in the rest of this subsection. Section 4.3 conducts a number of tests to assess the data consistency of the obtained model and Section 4.2 explores its dynamic properties. In particular, it demonstrates that the effects of a shock depend on its sign, size and the (initial) rate of unemployment.

4.1. An LSTAR model

Formal tests reject the null hypothesis of an AR(5) model against a STAR model, and favour an LSTAR model against an ESTAR representation of the data, see Table 2. It shows

Table 2
Testing for the appropriate model specification

d	Testing linearity	p-value
1	$F(15, 75) = 1.00$	[0.46]
2	$F(15, 74) = 0.98$	[0.47]
3	$F(15, 73) = 0.77$	[0.70]
4	$F(15, 72) = 1.16$	[0.32]
5	$F(15, 71) = 1.79$	[0.05]
Testing the form of non-linearity		
H_{04}	$F(5, 71) = 1.64$	[0.16]
H_{03}	$F(5, 76) = 0.57$	[0.72]
H_{02}	$F(5, 81) = 3.08$	[0.01]

Note: Sample period 1973:2–1997:1. See Section 2.1 for details about the tests.

that the AR(5) model is rejected at about the 5% level of significance when $d = 5$. And for $d = 5$, the lower panel of the table shows rejection of H_{02} at the 1% level of significance. This indicates that an LSTAR model can be a more appropriate characterisation of the unemployment process than an ESTAR model. Stronger rejection of the linearity hypothesis and of H_{02} can be achieved if numerically small and statistically insignificant terms are excluded from the auxiliary regressions for values of $d \in [1, 5]$. In this case, the AR(5) model may be rejected in favour of a STAR model for all values of d at the 10% level of significance, but still most strongly for $d = 5$ with a p-value of 0.003. Moreover, H_{02} may be rejected at the p-value of 0.003 for $d = 5$. The F-tests may have more power in these parsimonious auxiliary regressions than in their general versions.

We estimated an LSTAR model with five autoregressive terms and $d = 5$ by the method of maximum likelihood. The outcome indicated that most of the lagged terms in the non-linear part as well of the linear part of the model were insignificant. One reason for this could be the high degree of correlation between the regressors: the lagged terms of U_t. Indeed, when the general model was sequentially reduced to a parsimonious model, several of the lagged terms in the non-linear part of the model became significant. This parsimonious LSTAR model is presented in Table 3.

The estimate of the transition parameter γ is fairly large (3.48). Thus even a one percentage point deviation of unemployment from $\hat{c} \approx 3.6$ is sufficient to bring the estimated transition function $\hat{F}(U)$ close to 0 or 1, see Fig. 4. The transition parameter is imprecisely

Table 3
The LSTAR model and its long run properties

$$\hat{U}_t = \underset{(0.23)}{0.50} + \underset{(0.09)}{0.73}U_{t-1} + \underset{(0.12)}{0.37}U_{t-2} - \underset{(0.13)}{0.31}U_{t-5} + (-\underset{(0.16)}{0.41}U_{t-2} + \underset{(0.15)}{0.95}U_{t-4} - \underset{(0.19)}{0.43}U_{t-5}) \times [1 + \exp\{-\underset{(2.28)}{3.48}(U_{t-5} - \underset{(0.26)}{3.57})\}]^{-1}$$

Sample: 1973:2–1997:1, Log-likelihood value = 42.37 and $\hat{\sigma} = 0.41$

Long run properties

$$\hat{F}(U_{t-5}) = 0 : \hat{\varrho}_1 = \sum_{i=1}^{5} \hat{\phi}_i = 0.78, \qquad \hat{\mu}_1 \equiv \frac{0.50}{1 - \hat{\varrho}_1} \approx 2.3$$

$$\hat{F}(U_{t-5}) = 1 : \hat{\varrho}_2 = \sum_{i=1}^{5} (\hat{\phi}_i + \hat{\tilde{\phi}}_i) = 0.91, \qquad \hat{\mu}_2 \equiv \frac{0.50}{1 - \hat{\varrho}_2} \approx 5.1$$

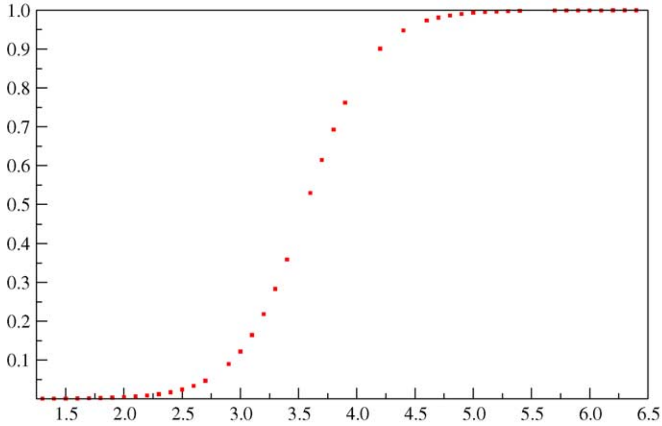

Fig. 4. A cross plot of the transition function $\hat{F}(U)$ (vertical axis) against the transition variable (U). One dot represents at least one observation.

estimated, which is not unexpected given the clustering of most of the observations around levels of 2 and 5%, cf. Fig. 2. In general, precise estimates of large transition parameters require many observations in the neighbourhood of c (Teräsvirta, 1994). This is attributed to the problem of accurate estimation of γ when the transition variable U_{t-d} is close to the threshold value c and the transition function is steep. In this case the transition function $F(U_{t-d})$ rises rapidly for small deviations between U_{t-d} and c, and its shape becomes consistent with a broad range of values of γ. The high standard deviation of $\hat{\gamma}$ is assumed to reflect this feature.

Fig. 5 shows that $\hat{F}(U)$ moves rapidly from 0 to 1 during 1988 and stays close to 1 in the subsequent periods of the sample, i.e. until 1997:1. The transition to the high unemployment

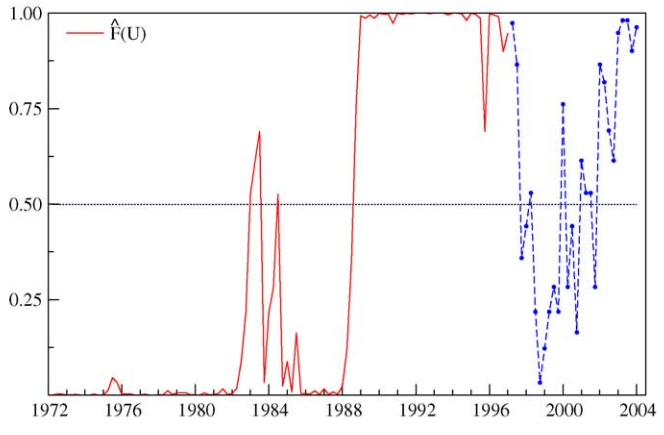

Fig. 5. Values of the estimated transition function $\hat{F}(U)$ over the estimation period 1973:2–1997:1 and out-of-sample 1997:2–2004:1. To date the transition of unemployment process from one regime to another, the transition function $\hat{F}(U)$ is derived by using U_t rather than U_{t-5}.

regime during 1988 is consistent with the findings of Skalin and Teräsvirta (2002). In the period 1982/83, $\hat{F}(U)$ displays a transitory increase from zero but falls short of reaching 1. The unemployment process in this period may be interpreted as being in-between the two regimes. The implied parameter changes in the early 1980s and in 1988 are in line with the results from the parameter constancy tests in Fig. 3.

Calculations of $\hat{F}(U)$ using new unemployment data for the period 1997:2–2004:1 suggest that the LSTAR model, which allows for back-and-forth smooth shifts between regimes, can easily account for the unemployment behaviour in this period, see Figs. 1 and 5. Values of $\hat{F}(U)$ indicate a shift to the lower unemployment regime in the second half of 1997 and a rather gradual reversal to the higher equilibrium regime from 1999:1 and onwards. If we let values of the transition function below 0.5 indicate the prevalence of the lower unemployment regime, and of the higher unemployment regime otherwise, then unemployment was probably in the lower regime in the period 1997:4–2000:4. Thus it seems unlikely that a model which would imply a permanent transition of Norwegian unemployment to a higher equilibrium rate in the sample period, can account for the apparent reversion to the lower unemployment regime in the post-sample period (cf. Skalin and Teräsvirta, 2002). Nor would a SETAR model, which would imply abrupt transitions between the two regimes, be consistent with the gradual transition to the higher unemployment regime from 1999 and onwards.

The LSTAR model has a higher explanatory power than the AR(5) model. The log likelihood value of this model is 42.37 compared with 35.8 for the AR(5) model, see Tables 1 and 3. The ratio between the standard deviations of residuals is about $0.93 = 0.41/0.44$.

The results in Table 4 suggest that the LSTAR model is consistent with the in-sample unemployment behaviour. None of the tests regarding the assumptions about residuals is rejected at the 5% levels of significance. The following null hypotheses are tested: absence of autocorrelation up to five lags, no heteroscedasticity, including ARCH type up to order five, and that the residuals have a normal distribution. The tests for no remaining nonlinearity of

Table 4
Testing the adequacy of the LSTAR model

	Maximum lag q or d_2				
	1	2	3	4	5
AR(q), $F(q, T-q-9)$	[0.92]	[0.86]	[0.93]	[0.98]	[0.99]
ARCH(q), $F(q, T-2q-9)$	[0.68]	[0.29]	[0.46]	[0.48]	[0.57]
Nonlinearity (d_2), $F(15, 70)$	[0.20]	[0.47]	[0.34]	[0.27]	[0.27]

Heteroscedasticity: $F_{\mathrm{Xi}^2}(16, 70) = 0.87$ [0.60]
Heteroscedasticity: $F_{\mathrm{XiXj}}(41, 45) = 0.77$ [0.80]
Normality: $\chi^2(2) = 1.15$ [0.56]
Parameter constancy: $F(21, 66) = 0.85$ [0.65]

Note: The tests for no autocorrelation, no remaining nonlinearity of STAR type and parameter constancy are those suggested by Eitrheim and Teräsvirta (1996). Large brackets [·] contain the p-values of the test statistics. In testing for no-autocorrelation, the (initial) missing values of residuals were set at zero, as recommended by Teräsvirta (1998). The tests for no remaining non-linearity of STAR type were conducted for U_{t-d_2} as the transition variable. The other tests for no heteroscedastisity and normality of the errors are the standard tests, as used in linear models, see Table 1 for details.

STAR type do not indicate any remaining nonlinearity in the model. The test of parameter constancy suggests that the possible non-constancies in the parameters have been modelled adequately. The non-rejection of this test also implies that there have not been significant changes in the two equilibria over time. This test has power against the alternative of both smooth and abrupt changes in the parameters (Eitrheim and Teräsvirta, 1996).

4.2. Dynamic properties of the model

The LSTAR model suggests that Norwegian unemployment is stationary within both regimes, though quite persistent. The equilibria associated with the low and high unemployment regimes, i.e. with $\hat{F}(U) = 0$ and $\hat{F}(U) = 1$, are about 2.3 and 5.1%, respectively, see Table 3. These estimates are consistent with the observed behaviour of unemployment in Figs. 1 and 2.

The lower unemployment equilibrium seems to be more stable, or attractive, than the higher one. This is suggested by the relatively lower degree of unemployment persistence around the lower equilibrium rate. It is 0.78 around the lower equilibrium rate and 0.91 around the higher equilibrium rate, see Table 3. This may be a reflection of the Norwegian government's policies, which were aimed at low unemployment during the sample period (cf. Manning, 1990). In addition, the wage bargaining institutions are relatively centralised making it more difficult to leave the lower unemployment regime than the higher one (cf. Calmfors and Driffill, 1988).

The greater attraction of the lower equilibrium, compared with that of the higher equilibrium, implies that unemployment requires a stronger impetus to leave the lower equilibrium than the higher equilibrium. Figs. 6 and 7 display the response of unemployment when it is exposed to shocks of different sizes when in the lower and the higher equilibrium, respectively. In both figures the shocks hit in period 10. Fig. 6 shows that unemployment returns to the low equilibrium when exposed to positive single shocks ($\varepsilon > 0$) of sizes up to 1.83, which is about 4.5 times the standard error of residuals ($\hat{\sigma}$). A larger shock, however, makes

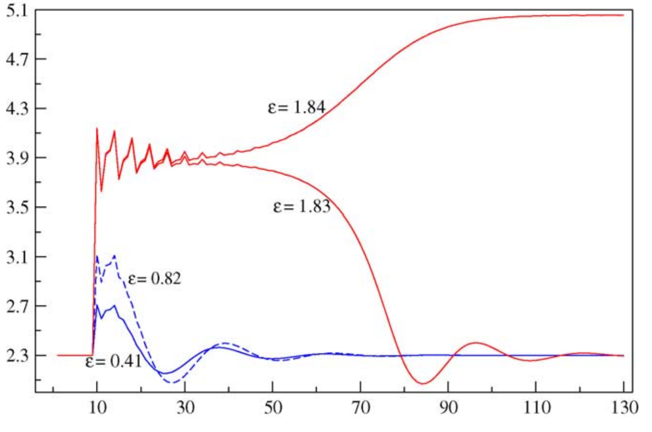

Fig. 6. Large and small shocks (ε) to the unemployment rate when it is at the low equilibrium rate of 2.3%. The horizontal axis shows number of periods (quarters).

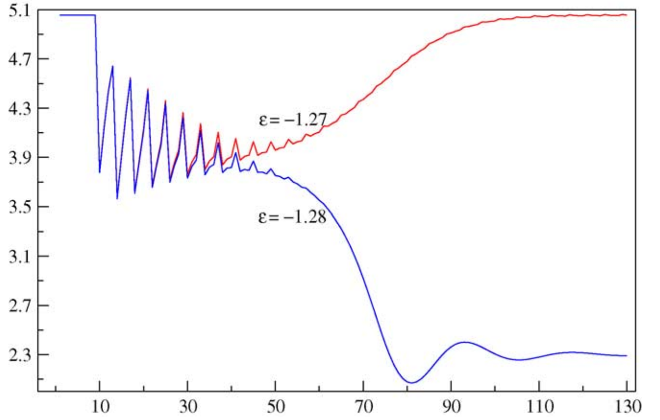

Fig. 7. Shocks (ε) of different sizes to the unemployment rate when it is at the high unemployment equilibrium of 5.1%.

it converge towards the high unemployment equilibrium, cf. the shift in the unemployment rate in 1988, see Fig. 1. From there, it requires a shock of size −1.28, i.e. of about $-3\hat{\sigma}$ to revert to the low equilibrium rate, see Fig. 7. Otherwise, it will get stuck at the high unemployment equilibrium. Needless to show, unemployment returns to the initial equilibrium level after a positive (negative) shock in the vicinity of the higher (lower) equilibrium rate, irrespective of the size of the shock.

The model suggests that the unemployment process does not require a large single shock to switch from one regime to another. A sequence of small shocks of the same sign may ultimately lead to a sufficient deviation of unemployment from its threshold value (3.57) to cause a transition from one regime to another. However, the small shocks have to be larger

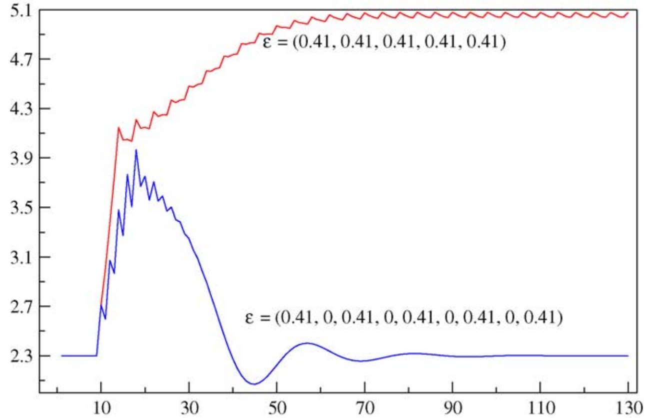

Fig. 8. A continuous sequence of five shocks, each equal to $\hat{\sigma} = 0.41$, and a discontinuous sequence of the same shocks to the unemployment rate when it is at a low equilibrium rate.

in sum than a large single shock, because of slight reversions towards the initial equilibrium during the intervals between the shocks. Fig. 8 displays the effects of a continuous sequence of five small shocks, each of size $\hat{\sigma} = 0.41$, when unemployment is initially at the lower equilibrium. Unemployment converges to the higher equilibrium even in this case.[6] The figure also shows the response to a discontinuous sequence of the five small shocks. In this case unemployment reverts to the initial equilibrium. This occurs because unemployment moves slightly back towards the initial equilibrium during the pause before the arrival of the next shock in the sequence, see Fig. 8. Therefore, it requires larger shocks than 0.41 in the periods afterwards to converge towards the higher unemployment regime.

The gradual reversion towards the lower unemployment equilibrium after 1992/1993 and towards the higher equilibrium since 1999 may be ascribed to sequences of predominantly negative and positive small shocks, respectively (see Figs. 1 and 5). The latter period has been characterised by relatively high wage growth, tight monetary policy, appreciation of the Norwegian exchange rate and recessionary impulses from abroad (e.g. Norges Bank, 2002). The relatively high wage growth (about 6% per annum) in 1998, the fall in oil prices (below 13 USD per barrel) and a sharp rise in interest rates (by 2 percentage points) in the autumn of 1998 may also explain why unemployment did not fall more than it did in 1998 and why it started climbing upwards thereafter.

The model implies that, within the vicinity of a given equilibrium, unemployment responds asymmetrically to large positive and negative shocks, but symmetrically to small positive and negative shocks. Moreover, the asymmetric response depends on the rate of unemployment. The latter is because the value of the transition function $\hat{F}(U)$, and hence the degree of persistence, which determines unemployment dynamics after a shock, depends on the rate of unemployment, cf. Table 3 and Fig. 4. This also explains the symmetric unemployment response around a given equilibrium when it is exposed to small positive or negative shocks. Fig. 4 shows that shocks within the range of ca. ± 0.41 do not alter the value of the transition function and hence unemployment dynamics. However, in the face of large shocks, i.e. outside the suggested range, the value of the transition function and the degree of persistence change with the size of shocks (depending on their sign), and thereby contribute to asymmetric unemployment dynamics.

For example, if unemployment is at the lower equilibrium level, a sufficiently large positive shock will shift the value of $\hat{F}(U)$ from zero to one, and the degree of persistence from 0.78 to 0.91. In the case of a negative shock of the same size, the value of $\hat{F}(U)$ and hence the degree of persistence will remain unaltered. The opposite happens when unemployment is at the higher equilibrium. In that case, the degree of persistence remains at 0.91 if unemployment is exposed to a positive shock of any size, but falls towards 0.78 if the shock is sufficiently large and negative. Intuitively, unemployment is generally attracted by both of the equilibria if a shock brings it in-between them, but by just the initial one if the shock is small, or if the shock moves unemployment even further away from the other equilibrium.

Accordingly, if unemployment is initially low, it rises faster than it falls after large negative and positive shocks, respectively. However, if unemployment is initially high, it

[6] For small shocks a set of values seem to repeat themselves when approaching the high unemployment equilibrium, cf. Fig. 8.

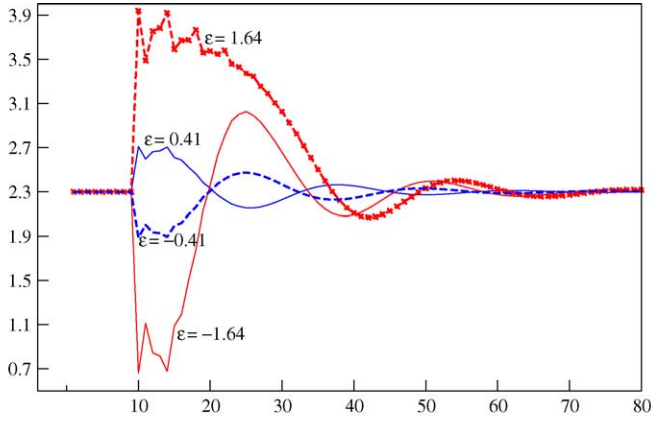

Fig. 9. The unemployment response to large positive and negative shocks of size 1.64 and −1.64, and the response to small shocks of size 0.41 and −0.41.

falls faster than it rises after large positive and negative shocks, respectively. Fig. 9 illustrates the case when unemployment at the lower equilibrium is disturbed by shocks of size −1.64 and 1.64. The figure also shows that unemployment adjustment is symmetric if the shocks are small, i.e. equal to −0.41 and 0.41. The importance of the sign of a shock for the unemployment adjustment becomes obvious when the shock is sufficiently large to shift unemployment from one regime to the other. In that case, large positive (negative) shocks can cause a switch from the lower (higher) regime, while large negative (positive) shocks only case a transitory deviation from the initial equilibrium rate, cf. Figs. 6 and 7.

The observed asymmetric dynamic response to positive and negative shocks is also consistent with the implications of studies reporting pro-cyclical hiring and or countercyclical firing costs (e.g. Burgess, 1992; Palm and Pfann, 1997). Accordingly, unemployment rises faster than it falls when at low levels, but falls faster than it rises when it is at relatively high levels. Moreover, the observed symmetric responses to small positive and negative shocks in Fig. 9 can be alternatively ascribed to convex employment adjustment costs, where small increases or decreases in the labour stock can be made with about the same costs (e.g. Hamermesh and Pfann, 1996, and the references therein).

4.3. Out-of-sample evaluation of the LSTAR model

In the following we test the stability of the LSTAR model by exposing it to data from seven additional years, 1997:2–2004:1. First, we reestimate the model on the extended sample 1973:2–2004:1 to investigate possible changes in the unemployment process. In particular, we examine whether estimates of persistence in each regime and of equilibria are robust to the extension of the data set and a break in the Norwegian Labour Force Survey (LFS) in 1996.

The LFS became a continuous survey in 1996:1, i.e. the survey frequency was increased from one week a month to every week. Secondly, the LFS questionnaire was changed: it relaxed the availability-for-work-requirement for job-seekers and started classifying persons on government funded qualification programmes to promote employment as unemployed, if they satisfied other requirements concerning job-seeking and availability-for-work. The latter group of persons was previously classified as employed.[7] It is not clear how increased monitoring of the labour market through more frequent surveys might affect the unemployment process. However, the changes in the LFS questionnaire may have raised the unemployment rate and the implied unemployment equilibria. Statistics Norway (2003) suggests that, all in all, changes in the 1996 raised the unemployment rate by about 0.50 percentage point as an annual average for 1996.

For the following years, 1997–2004, we are not aware of any evidence of the partial effect of all or specific changes in the LFS on the unemployment rate. The impact of relaxing the availability-for-work requirement on the unemployment rate is difficult to pin down. However, the substantial decline in the number of persons attending the qualification programmes, from 37 000 in 1996 to 25 000 in 1997 and 13 000 on average for the years 1998–2003, suggests that the reclassification itself is unlikely to contribute as much to raising the unemployment rate in the latter years as it did in 1996.

Formal tests, however, suggest that the changes in the LFS from 1996:1 did not have any significant and/or systematic effect on the unemployment rate. If the LSTAR model is estimated with a step dummy, which takes on a value of 1 from 1996:1 and onwards to represent the effect of changes in the LFS, the coefficient estimate of the step dummy would suggest an increase in the unemployment rate of 0.03 percentage point only and the null hypothesis of no effect from the step dummy would not be rejected with a p-value of 0.89. If the LSTAR model is reestimated on the extended sample (1973:2–2004:1) with the step dummy, its coefficient estimate becomes 0.06% and the associated p-value becomes 0.54.

Moreover, even if we allow the effects of the changes in the LFS to differ from one year to another through year-specific step dummies for the period 1996–2003, the effects are statistically insignificant at the 5% level of significance in all but one case; the step dummy for 1997 indicates a significant reduction in the unemployment rate by 0.40 (percentage point) at the 5% level, see Table 5. Numerically, the coefficient estimates of the year-specific dummies take on values in the range of −0.40 to 0.27. For the year 1996, however, the coefficient estimate of the dummy is 0.23, which is statistically close to the impact of the change in the LFS for the year 1996 suggested by Statistics Norway (2003). As suggested below, one cannot rule out the possibility that these dummy variables are partly capturing effects of economic shocks to the unemployment rate in addition to those of the changes in the LFS. We leave it to future studies, however, to disentangle possible effects of the changes in the LFS while controlling for other economic shocks.

Table 5 also suggests that the LSTAR model is quite robust to extension of the information set. It shows that estimates of the LSTAR model based on the extended sample are numerically quite close to those based on the initial sample period 1973:2–1997:1, cf. Table 3. In particular, the implied degrees of persistence $\hat{\varrho}_1$ and $\hat{\varrho}_2$ are 0.82 and 0.91, respectively, while the implied estimates of the two equilibria are 2.3 and 4.7%, respectively. The latter

[7] Previously, it was required that an out-of-work job-seeker was available for work in the week the survey was conducted for him/her to be considered unemployed. From 1996 it suffices that he/she is available for work within the next two weeks.

Table 5
The LSTAR model reestimated on the extended sample

$$\hat{U}_t = \underset{(0.25)}{0.42} + \underset{(0.08)}{0.69}U_{t-1} + \underset{(0.10)}{0.43}U_{t-2} - \underset{(0.13)}{0.30}U_{t-5} + (-\underset{(0.13)}{0.50}U_{t-2} + \underset{(0.12)}{0.98}U_{t-4} - \underset{(0.15)}{0.38}U_{t-5}) \times [1 + \exp\{-\underset{(3.06)}{3.70}(U_{t-5} - \underset{(0.22)}{3.37})\}]^{-1} + \underset{(0.19)}{0.23}d96 - \underset{(0.20)}{0.40}d97 - \underset{(0.21)}{0.35}d98 + \underset{(0.21)}{0.19}d99 + \underset{(0.20)}{0.02}d00 + \underset{(0.19)}{0.14}d01 + \underset{(0.19)}{0.19}d02 + \underset{(0.19)}{0.27}d03$$

Sample: 1973:2–2004:1

Note: The LSTAR model of Table 3 has been reestimated by the ML-method on the extended sample with year-specific step dummies. The parentheses contain estimates of standard errors.

estimate indicates a lower equilibrium rate of unemployment in the higher unemployment regime than suggested in Table 3. Statistically, however, null hypotheses of equality between estimates across Tables 3 and 5 would not be rejected at standard levels of significance.

Table 6 tests the stability of the parameters in both regimes jointly and separately, i.e. conditional on each other's stability. Column 2 shows that the null hypotheses of stability in the parameters defining both regimes in the reestimated model are not rejected at the standard levels of significance. Column 3 undertakes a more demanding test of parameter stability. It tests the stability of the parameters in the initial model using the extended data set. It appears that even in this case, the null hypotheses of parameter stability defining both regimes are not rejected at the standard levels of significance. This lends support to the above findings that changes in the LFS have not had any lasting effect on the unemployment process, especially because the LSTAR model has not been reestimated and/or altered to fit the post-sample period.

Fig. 10 shows that both the initial model and the reestimated model are largely in agreement on the dating of the switches between the two unemployment regimes. $\hat{F}(U)$ is based on the initial model but is calculated using new unemployment data for the period 1997:2–2004:1, while $\tilde{F}(U)$ is based on the reestimated model in Table 5. We note that the latter function is more conclusive regarding the dating of switches between the two regimes, mainly because of the relatively larger estimated value of the transition parameter, 3.70 versus 3.48. Both transition functions indicate a shift towards the lower unemployment regime

Table 6
Tests of parameter constancy

LSTAR model in	Table 5	Table 3
Estimation period	1973:2–2004:1	1973:2–1997:1
Evaluation period	1973:2–2004:1	1973:2–2004:1
Constancy test:		
(a) Regime 1 and 2: F_1 (21, d.f.)	0.83 [0.68]	0.95 [0.55]
(b) Regime 1: F_2 (12, d.f.)	0.42 [0.95]	1.23 [0.27]
(c) Regime 2: F_3 (9, d.f.)	0.36 [0.95]	0.67 [0.73]

Note: Tests of parameter constancy are those suggested by Eitrheim and Teräsvirta (1996). In rows (a)–(c) we test the following null hypotheses: "all parameters characterising both regimes are constant", "all parameters defining the lower regime, including the intercept, are constant" and "all parameters characterising the higher regime, except the intercept, are constant". In rows (b)–(c), the parameters characterising the higher and the lower regimes, respectively, are assumed constant a priori. d.f. denotes degrees of freedom. Large brackets [·] contain the p-values of the test statistics.

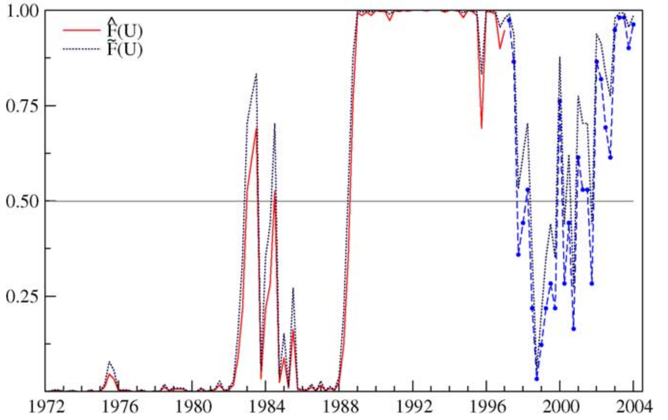

Fig. 10. Values of the estimated transition functions $\hat{F}(U)$ and $\tilde{F}(U)$ over the extended sample 1973:2–2004:1. $\tilde{F}(U)$ is defined as $[1+\exp\{-3.70(U_{t-5}-3.37)\}]^{-1}$, see Table 5.

in the second half of 1997 and a rather gradual reversal to the higher regime from 1999:1 and onwards. If we let a transition function value of less than 0.5 indicate the prevalence of the lower unemployment regime, and of the higher unemployment regime otherwise, then both transition functions indicate the lower regime in the period 1998:3–2000:4 and the higher regime afterwards. However, strict application of this rule makes the two transition functions conflict with each other regarding the regime prevailing in 1997:4–1998:1, in particular. $\hat{F}(U)$ suggests the lower unemployment regime in this periods, while $\tilde{F}(U)$ indicates continuation of the higher unemployment regime until 1998:3.

Interestingly, the transition towards the lower unemployment regime in 1997–1998 and the subsequent gradual reversion towards the higher regime seem to be consistent with the signs and size of the year-specific step dummies in Table 5. The step dummies can be interpreted as unidentified transitory shocks to unemployment. It is beyond the scope of this study to identify these shocks, but the dummy variables for the years 1997–1998 could be partly representing effects of the relaxation of Norwegian monetary policy in 1997 and the first half of 1998 (e.g. Norges Bank, 2002). The dummy variables for the following years could be partly associated with relatively high wage costs, tight monetary policy, appreciation of the Norwegian exchange rate and recessionary impulses from abroad during the latter years. However, owing to time lags in the transmission of shocks to unemployment which may even depend on the nature of the shocks, a more extensive investigation is required to associate particular movements in the unemployment rate with specific shocks.

5. Conclusions

This paper aimed to derive a data-consistent univariate model to test for the possibility of multiple equilibria in Norwegian unemployment and investigate whether it adjusts asymmetrically when exposed to positive and negative shocks of different sizes. To this end an LSTAR model of the unemployment rate has been derived. It has been shown that this model outperforms a linear AR model in explanatory power and appears to be consistent with the in-sample (1973:2–1997:1) and out-of-sample (1997:2–2004:1) behaviour of the unemployment rate.

The LSTAR model provides evidence in support of two stable unemployment equilibria at 2.3 and 5.1%, respectively. The lower equilibrium is relatively more stable than the higher one. The unemployment process tends to exhibit equilibrium reversion towards 2.3% before 1988 but thereafter towards 5.1% until the end of 1997. In the period 1998–2000, it seems that unemployment was in the lower unemployment regime, and in the higher regime afterwards, i.e. until the end of our sample 2004:1. Our findings indicate a rather abrupt transition from the low to the high unemployment regime in 1988, but rather gradual shifts between the two regimes in the period afterwards.

The model implies that a large transitory shock or a sequence of small shocks may cause a transition between equilibria. However, the small shocks must be larger in sum than the single large shock due to the equilibrium reversion property. These findings suggest that demand shocks such as monetary policy shocks can be non-neutral in the long run. However, this applies only within a limited range of the unemployment rate, as even particularly large shocks are unable to shift unemployment permanently below the lower equilibrium or above the higher one.

The model also implies that unemployment displays an asymmetric response to large positive and negative shocks, while the response is symmetric to small positive and negative shocks. In addition, the pace of adjustment depends on the rate of unemployment. It follows that unemployment rises faster than it falls towards the lower equilibrium rate only when the disturbance are large and unemployment is initially low. If unemployment is initially high, the opposite may be the case. These observations are consistent with empirical studies reporting faster employment adjustments in recessions than in booms. Our observation of a symmetric response to small positive and negative shocks also fits with the presence of convex employment adjustment costs, when small increases or decreases in the labour stock can be made with about the same costs.

The support for the LSTAR model against the linear AR model suggests that it is more fruitful to treat hysteresis as a non-linear phenomenon rather than a linear phenomenon. This enables one to reconcile the empirical evidence from long and short unemployment time series and allows for more flexibility regarding the effects of shocks. However, empirical analyses based on univariate time series need to be supplemented by multivariate studies in order to identify mechanisms that may explain the presence of multiple equilibria and asymmetric adjustment, and to point out factors that may explain transitions between equilibria.

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