State dependent effects in labour and foreign exchange markets

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February 12, 2000

Contents

[0] This essay is a revised version of the one presented in the submitted thesis.

Preface

The impact of monetary policy in the short- and long-run, and over the business cycle are issues of great interest to a central bank. The response of nominal and real exchange rates to oil price fluctuations are also of particular interest to an economy that relies heavily on petroleum resources, such as the Norwegian economy. Based on Norwegian data, Qaisar Farooq Akram's dissertation addresses these issues by testing for multiple unemployment equilibria, by examining the employment behaviour in slack and tight labour markets, and by investigating the short- and long-run effects of oil prices on nominal and real exchange rates. The dissertation employs the latest techniques in econometric methodology for studying macroeconomic time series across different regimes and states.

The dissertation is part of the author's Dr. Polit exam at the University of Oslo, Department of Economics and it was defended on May 4, 2001. It is a great pleasure to mention that this dissertation has been awarded the first prize for the best Ph.D. thesis in macroeconomics submitted in Norway in the period May 1998–April 2001 by Norges Bank Fund for Economic Research. Norges Bank is pleased to make this dissertation accessible for a wider audience.

Oslo, February 12, 2002

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Acknowledgements

I have been fortunate to work on this thesis while being at the Research Department of Norges Bank and during my stay at the Department of Economics, University of Oxford in the academic year 1999-2000. I would like to express my gratitude to Norges Bank for its generous support of my studies and the research presented in this thesis, in particular. The views expressed in this thesis are however my sole responsibility. I am also grateful to the Department of Economics, University of Oxford for its hospitality and for providing such excellent research facilities; many thanks to David F. Hendry for his support in this regard.

Over the years I have become heavily indebted to numerous wonderful and nice persons, for their helpful comments, suggestions and encouraging remarks. I gratefully acknowledge my debt to especially Øyvind Eitrheim, David F. Hendry, Timo Teräsvirta, Bent Nielsen, Hans-Martin Krolzig, Karl Ove Moene, Christian Dahl and Gunnar Bårdsen.

I also owe many thanks to my colleagues at Norges Bank, particularly in the Economics and the Research Department for continuously suggesting issues to be addressed and offering constructive criticism. Parts of this thesis have been inspired by discussions and interaction with my colleagues, and comments from participants at workshops and seminars at Norges Bank have been invaluable during the writing of this thesis. Karsten Molværsmyr and Fredrik Wulfsberg have cheerfully helped me out of computer related problems and the staff of the Norges Bank Library have provided efficient help. Also, I wish to express my sincere appreciation of the great contribution made by Eilev Jansen in creating a stimulating and friendly environment at the Research Department. He has been very supportive, interested and encouraging towards the work undertaken in this thesis and has offered valuable suggestions and advice along the way.

Above all, I would like to thank Ragnar Nymoen, my thesis supervisor, for offering penetrating comments and useful suggestions, and for being a constant source of encouragement. His kind support on a professional and personal level, ever since he took me on as a research assistant in 1994 when I was a graduate student, is highly appreciated. He has also coauthored the second essay in this thesis, which has benefited greatly from suggestions by Karl Ove Moene.

Actually, working on this thesis has been great fun and a highly rewarding social activity, owing to the interaction with a large number of knowledgeable and kind persons, only some of whom are mentioned above. I have learnt a lot from them, but certainly not enough as I am a slow learner.

I would also like to thank my thesis committee, which consisted of Professors Svend Hylleberg, Ronald MacDonald and Asbjørn Rødseth, for insightful and constructive comments. The defence of this thesis took place on May 4, 2001.

Special thanks go to Steinar Strøm for arousing my interest in macroeconomics through his inspiring lectures at the University of Oslo and for steering me into courses in econometrics. It is also a great pleasure to thank Tone Enger at the Department of Economics, University of Oslo for her guidance on practical and formal matters.

Finally, I am grateful to my family, especially my parents Amna and Akram, my sisters and my wife Saadia for their immense support and encouragement throughout my studies.

QFA

February 12, 2002, Oslo

1. Introduction and overview

1. Introduction

This thesis consists of four essays on different topics primarily related to the labour and the foreign exchange markets. The essays contain empirical analyses of issues that in most cases require that we go beyond the use of linear econometric models and adopt non-linear models. Although the issues investigated are by no means specific to the Norwegian economy, the empirical analyses are based on macroeconomic time series pertaining to the Norwegian economy.

The different essays are concerned with the behaviour of both real and financial variables such as the unemployment rate, employment and the real and nominal exchange rates. In addition, the analyses are conducted by employing a wide range of techniques and models. The essays have, however, a large number of common features. In this overview we point out some of the main common features, particularly those associated with our theoretical and empirical approach. We also draw attention to similarities between the conclusions reached in the different essays.

A brief review of the theoretical and empirical literature on each of the issues is provided in the respective essays, which also contain brief outlines of the relevant econometric methods and techniques. These are well known in the econometric literature and are presented in inter alia Hendry (1995), Johansen (1995b), Teräsvirta (1998), Hamilton (1989) and Krolzig (1997). Thus in this overview we present only the main findings and characteristics of earlier studies and a general outline of the modelling approach. Also, we give some space to suggestions for future research in the light of the results and arguments presented in this thesis.

This overview is organised as follows. The next section (2) presents abstracts of the four essays. Against this background, Section 3 points out unifying features of these essays. Finally, Section 4 offers some suggestions for further research.

2. Overview

This overview provides a brief summary of the essays, with focus on their relationship to earlier work, our empirical approach and the main conclusions reached in the essays.

2.1. Essay 1: "Multiple unemployment equilibria and asymmetries in Norwegian unemployment"

Essay 1 characterises the Norwegian unemployment rate in a framework that allows for multiple equilibria and asymmetric responses to positive and negative shocks. The multiple equilibria approach is motivated by a number of shortcomings with the two common approaches to the study of European unemployment: (a) the unique equilibrium or the NAIRU approach and (b) the hysteresis approach that equates hysteresis with unemployment being a random walk process. In the latter approach every shock has a permanent effect on the level of unemployment while in the first approach only structural and institutional changes are allowed to bring about a change in the equilibrium level; every other shock brings about transitory deviations from the equilibrium level.

The European unemployment experience has, however, revealed slow if any tendency of actual unemployment to revert to a unique equilibrium. Moreover, estimates of the equilibrium level have been shown to track the actual rate, without matching structural and institutional changes. (see for instance Layard et al. (1991), Cromb (1993), Elmeskov and MacFarlan (1993). Apparently, the hysteresis approach has been able to account for the high degree of persistence in European unemployment rates as the presence of a unit root in the unemployment series is seldom rejected, empirically.

However, the practice of equating hysteresis with the presence of a unit root in a linear model has been questioned (see e.g. Amable et al. (1995), Cross (1995), Røed (1997). Firstly, it compels every shock, irrespective of its size and sign to have a permanent effect on the level of unemployment, disregarding the existence of (endogenous) stabilising mechanisms. Secondly, long series of unemployment rate data often show that it does not wander around randomly, but revert to its past levels, sooner or later, cf. Layard et al. (1991) and Bianchi and Zoega (1997). Indeed, the bounded nature of the unemployment rate series prevents it from taking values outside the 0-1 range. Thirdly, the empirical evidence in favour of a unit root, or high degree of persistence, in relatively smaller 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 the series, see e.g. Perron (1989) and Banerjee et al. (1993). Finally, the linear models imply symmetric dynamic behaviour when unemployment rises or falls. Observations of the unemployment behaviour, however, indicate that it rises faster than it declines. Such asymmetries are often explained by asymmetric adjustment costs or by the hiring and firing practices of firms, cf. Hamermesh and Pfann (1996), Johansen (1982). This criticism is also applicable to most studies within the unique equilibrium approach.

Models of multiple equilibria appear to be capable of reconciling the empirical evidence from long and short time series and allow for more flexibility with regard to the effects of shocks. In these models, a large shock may cause a movement from one equilibrium level to another while small shocks only cause a temporary deviation from a given equilibrium level. Models that display multiple equilibria are often based on the existence of reciprocal externalities in various guises. These can arise from trading and exchange opportunities as in Diamond (1982) and Cooper and John (1985), due to spillovers of demand across markets as in e.g. Weitzman (1982) and Murphy et al. (1989) or due to costs associated with layoffs and hirings as in Saint-Paul (1995) and Moene et al. (1997b), respectively. A characteristic feature of theoretical models of multiple equilibria is that they allow for movements between the equilibria depending on the size and sign of shocks. Furthermore, changes in possible multiple equilibria due to structural and institutional changes are allowed to affect a given set of equilibria.

Non-linear models are required to entertain the possibility of multiple equilibria, i.e. to allow for asymmetric responses to the size of a shock. Theories of multiple equilibria, as well as models of unique equilibrium, are often silent on whether the response towards shocks depends on their sign or not. This property, i.e. asymmetric response to the sign of a shock, can also be incorporated into non-linear models and tested for.

The essay employs a univariate framework to test for multiple equilibria and asymmetric response to positive and negative shocks. The specific non-linear models considered are: the Markov regime switching model and a logistic smooth transition autoregressive model (LSTAR) model, see e.g. Hamilton (1989) and Teräsvirta (1998). In these models, the unemployment response to shocks is regime/state dependent. Models with a unique equilibrium and hysteresis in the unit root sense can be derived as special cases of these models. Possible changes in multiple equilibria are interpreted as shifts in the parameters defining a given equilibrium. The merits of the derived non-linear autoregressive models are compared with a linear autoregressive model. Furthermore, we undertake an extensive evaluation of the derived models in order to examine whether they are data consistent.

However, only a logistic smooth transition autoregressive (LSTAR) model turns out to be data consistent. Accordingly, unemployment is modelled as a non-integrated variable that has switched between two stable equilibria during the sample period. It is shown that a large shock, or a sequence of small shocks, can cause a transition from one equilibrium level to another and thereby have a permanent effect on the unemployment rate. The model also implies that unemployment displays asymmetric response to large positive and negative shocks, while the response is symmetric to small positive and negative shocks. In other words, unemployment recovers faster from a fall than a rise, only when the disturbances are large. These findings are consistent with the results for a number of OECD countries, see Bianchi and Zoega (1998) and Skalin and Teräsvirta (1999).

The univariate framework adapted in this essay and in general, does not shed light on the sources of shocks, of persistence and on mechanisms which may provide scope for multiple equilibria. However, it provides a convenient way to test for multiple equilibria and to draw out the characteristic features of the unemployment series. By this, it can provide stylised facts to be explained by theoretical and multivariate empirical models.

2.2. Essay 2: "Employment adjustment in slack and tight labour markets"

In Essay 2 we investigate whether shortage and abundance of labour affect employment's adjustment towards its equilibrium and its response to changes in forcing variables. Existing empirical evidence on state dependence in employment adjustment and in the effects of shocks seems ambiguous. Though a number of studies report state dependent employment adjustment or effects of shocks, evidence supporting joint occurrence of state dependent adjustment and effects of shocks remains scarce, see e.g. Smyth (1984), Burgess (1988), (1992a) and (1992b). Thus there seems to be some inconsistency in the results, as both implications of the hypothesis of state dependent employment response are not supported.

Also, we investigate whether anticipated difficulties in hiring can explain the presence of multiple equilibria in the labour market as suggested by e.g. Essay 1. Moene et al. (1997b) argue that perceived rationing in the labour market impinges on the firms employment decisions. Accordingly, states of high and low employment characterised by shortage and abundance of labour, respectively, may be self-sustaining and lead to multiple (un)employment equilibria. A number of recent studies presents evidence of multiple equilibria in unemployment, see Bianchi and Zoega (1998), Skalin and Teräsvirta (1999) and Akram (1999). It emerges that large shocks lead to movements between a given set of multiple equilibria while institutional changes shift the equilibria itself. However, the evidence is based on univariate models and hence it is not possible to identify the economic mechanisms giving rise to the multiple equilibria. That is to say that the evidence is not only consistent with the friction argument of Moene et al. (1997b), but also with numerous other arguments for multiple equilibria, as presented in e.g., Cooper and John (1985), Manning (1990), Murphy et al. (1989), Pagano (1990) and Saint-Paul (1995).

We model employment in a partial system framework, while most of the earlier studies are based on single equation models of employment. Here manufacturing employment, working hours and aggregate unemployment are treated as endogenous while their commonly assumed determinants are conditioned on. The latter in order to make allowance for a wide range of explanatory variables in the system. Most of the earlier empirical studies are based on single equation models of employment with untested weak exogeneity assumptions about employment determinants for e.g. the parameters defining the equilibrium level of employment and the adjustment of actual employment towards this level, see Engle et al. (1983) for definitions of the different exogeneity concepts. A few recent studies, however, model employment within a system framework but these are reduced form in their nature, usually based on bivariate vector autoregressive models, see e.g., Acemoglu and Scott (1994) and Krolzig and Toro (1999). Hence, they are not suited for providing information about possible state dependence in the effects of the wide range of variables that are usually considered in e.g. the single equation models of employment.

Furthermore, we control for possible asymmetric response to positive and negative changes in forcing variables when testing for state/cycle dependent employment response. Earlier studies seem to have tested state dependent employment adjustment without taking into account the influence on the results of such response, see e.g. Acemoglu and Scott (1994) and Burgess (1988), (1992a) and (1992b). The asymmetric response to shocks, or alternatively sign dependent response, could be due to asymmetric but state independent employment adjustment costs, see e.g. Hamermesh and Pfann (1996) and Escribano and Pfann (1998). If hiring costs are greater than firing costs, the sign dependent and state dependent responses may lead to observationally similar employment behaviour if a tight labour market coincides with predominantly positive shocks and and a slack labour market coincides with predominantly negative shocks. Intuitively, such coincidences seem to be a rule rather than exception, hence, the failure to disentangle sign dependent response from state dependent response appear to be an additional shortcoming with earlier studies.

Methodologically, this essay builds on Krolzig and Toro (1999) who employ a Markov regime switching vector equilibrium correcting model (MS-VEqCM) to allow for state dependence in the parameters, see Krolzig (1997) and Hamilton (1989) among others. A generalisation of an MS-EqCM of manufacturing employment is employed to control for possible asymmetric response to over- and undermanning (relative to the equilibrium employment) and to positive and negative changes in the forcing variables.

The empirical evidence in this essay suggests that the dynamic behaviour of employment alters with movements between a slack and tight labour market. 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 hirings due to labour shortage contribute to labour hoarding. These conclusions have appeared robust to allowance for asymmetric response to shocks. However, our evidence does not suggest that present and anticipated difficulties in hiring by themselves lead to multiple equilibria in the labour market.

The evidence of state dependent employment behaviour implies that a linear (constant parameter) characterisation of the employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions. Also, it follows that linear models may lead to overestimation of wage and price growth in recessions and underestimation in expansions. However, the possible implications for the wage and price growth are not investigated further.

2.3. Essay 3: "When does the oil price affect the Norwegian exchange rate?"

In Essay 3 we investigate whether linear exchange rate models lead to an underestimation of the Norwegian exchange rates response to oil price fluctuations and hence a failure to explain major changes in the exchange rate. The inquiry is motivated by the apparently puzzling results obtained using linear exchange rate models. They tend to show a numerically weak and/or statistically insignificant relation between the price of crude oil and the Norwegian nominal exchange rate, see e.g. Bjørvik et al. (1998) and Akram and Holter (1996). This is in contrast to the common perception that the price of crude oil has a significant influence on the Norwegian exchange rate. For example, the Norwegian currency crises in the 1990s, i.e. the appreciation pressure in 1996/97 and the depreciation pressure in 1998/1999, have been attributed to the rise and fall of oil prices, see e.g. Alexander et al. (1997), Haldane (1997) and Norges Bank (1998) for details. Likewise, the large devaluation of the krone in 1986 is often explained with reference to low oil prices in 1985/86, see e.g. Norges Bank (1987, pp. 17).[1] The assumed link between the oil price and the value of the krone is based on the size of the petroleum sector relative to GDP, 10-20 % since the mid 1970s, and its relatively large share in Norway's total export of goods and services (around 1/3), see Aslaksen and Bjerkholt (1986) and Statistics Norway (1998).

The possibility of a non-linear relation can be motivated by asymmetric costs of stabilising the exchange rate in the face of appreciation and depreciation pressure. Since 1972,

[1] In May 1986, the krone was devalued by 12 per cent relative to a trade weighted currency basket, mainly composed of (western) European currencies, see Norges Bank (1987, pp. 35-38) for details about the composition of the basket. The appreciation in the spring of 1997 and the depreciation in the autumn of 1998 were of around 10 per cent to the ECU.

the Norwegian monetary policy has been aimed at exchange rate stabilisation against (western) European currencies, see Alexander et al. (1997) and Norges Bank (1987) and (1995) for details and overview. In this monetary policy framework, the nominal exchange rate will display (excessive) fluctuations, due to appreciation or depreciation pressure arising from changes in e.g. oil prices, only if the central bank is unable or unwilling to ensure stability in the exchange rate. It follows that one is more likely to observe a negative relation between oil prices and the value of the krone when the authorities abandon the practice of currency stabilisation. Studies of currency crises suggest that central banks are often more willing to and capable of resisting pressure for currency appreciation than depreciation pressure, cf. Flood and Marion (1998) and the references therein. This asymmetry is explained by pointing to the higher costs of resisting depreciation pressure than appreciation pressure. The costs are usually measured in terms of sacrifices of objectives other than exchange rate stabilisation pursued by a central bank. These may be concerns for unemployment, competitiveness, economic growth, inflation and/or the viability of financial institutions due to its role as a lender of last resort, cf. Obstfeld (1990) and Calvo (1998).[2] It follows that a possible relation between the oil price and the value of the Norwegian krone is likely to be stronger in the face of low and falling oil prices compared than in the opposite case.

We examine daily observations of the oil price and the value of the Norwegian krone against European currencies to explore the possibility of a non-linear relation between these variables. The investigation reveals the existence of a strong non-linear relation between the ECU and the oil price. Specifically, it shows that the strength of the relation depends on whether the oil price is below, inside or above the range of 14-20 US dollars a barrel. Moreover, it depends on whether the oil price is displaying a falling or rising trend. The relation is relatively strong when oil prices are below 14 dollars and are falling. The relation is absent if oil prices fluctuates within the range of 14-20 US dollars, which is considered as the normal range of oil prices in the sample. At levels above the normal

[2] Generally a trade off will exist between realisation of these additional objectives where it may appear less costly for a central bank to e.g. lower interest rates in the face of appreciation pressure than raise interest rates in the face of depreciation pressure. Especially, if it is more concerned with the "side effects" on activity level than on inflation.

range, the relation reappears but is significantly weaker than the relation when oil prices are below the normal range.

These non-linear effects are tested and quantified within equilibrium correcting models of the NOK/ECU rate and the effective nominal exchange rate to control for the influence of other macroeconomic variables. The models are derived on monthly data from the 1990s, and on quarterly data since the end of the Bretton Woods system. Thus, the empirical non-linear effects are not artefacts of a specific model and a data sample.

It is shown that models with non-linear oil price effects outperform similar models with linear oil price effects. The latter models grossly underestimate the exchange rate response to oil price changes in a state of low oil prices. We undertake an extensive evaluation of the derived models to demonstrate the robustness of the results. It is noteworthy that the derived multivariate exchange rate models have remarkably stable parameter estimates and relatively high explanatory power. This is encouraging against the background of widespread pessimism in the literature regarding the possibility of deriving exchange rate models with such properties, with or without non-linear effects of macroeconomic variables, see e.g. Meese and Rose (1991), Meese (1990) and Frankel and Rose (1995). Therefore, the paper not only suggests that one takes a new look at studies that have reported unstable oil price effects on exchange rates, but also offers results that can be utilised in further theoretical and empirical research on exchange rates.

2.4. Essay 4: "PPP despite real shocks: An empirical analysis of the Norwegian real exchange rate"

In the last essay we analyse the long run behaviour of the Norwegian real exchange rate. In particular, we undertake explicit tests of the implications of the purchasing power parity (PPP) hypothesis using data from Norway and its trading partners.

The empirical analysis is undertaken in a linear framework and the sample consists of quarterly observation over the period 1972:1-1997:4. Specifically, the Norwegian effective real exchange rate is characterised in a univariate model. The equilibrium real exchange rate is conveniently estimated on the basis of this model. Recursive estimates of the equilibrium real exchange rate are used to examine whether it has changed or remained constant as implied by the PPP theory. Other implications of the PPP theory are tested using full and partial vector autoregressive models of the effective nominal exchange rate, domestic and foreign consumer prices and interest rates and the oil price. Here we employ the Johansen procedure as suggested in e.g. Johansen (1988). We test explicitly whether: (i) the nominal exchange rate only depends on domestic and foreign prices in the long run, (ii) domestic and foreign prices have symmetric and proportional effects on the nominal exchange rate and (iii) whether both the nominal exchange rate and domestic prices adjust to deviations from the purchasing power parity and contribute to convergence towards a constant equilibrium real exchange rate. The results also shed light on whether a shift in the monetary policy target from e.g. exchange rate stabilisation to inflation targeting may affect the process determining domestic inflation.

Our empirical results provide strong support for long run purchasing power parity between Norway and its trading partners. Moreover, the half life of a deviation from the equilibrium real exchange rate is only 6 quarters. These are novel results which stand out against the common findings in the vast literature on the PPP hypothesis. In particular son, since the Norwegian economy has been exposed to numerous real shocks in the sample period such as discoveries of petroleum resources and their revaluations through oil price fluctuations. It is well known that PPP is commonly rejected in data predominantly exposed to real shocks, see e.g. Patel (1990) and Cheung and Lai (2000). Often, PPP is only supported in studies that employ data samples from periods where monetary shocks dominate and/or that employ long time series, e.g. with time spans of more than a half century. The consensus estimates on the half life measure range from 2.5 to 6 years on data from industrialised countries, see e.g. Froot and Rogoff (1994), Rogoff (1996), Isard (1995) and MacDonald (1995) and the references therein.

We ascribe the strong support for the PPP and in particular the relatively fast convergence to the equilibrium real exchange rate to a number of factors. Firstly, the Norwegian economy is likely to be more exposed to arbitrage pressure than e.g., continental European economies, since it is small and relatively open. For example, the average of Norwegian exports and imports is more than 1/3 of GDP, which is almost the twice of that for e.g., France, Germany and Italy, see Haldane (1997). The economy's openness is also testified by the relatively large weight, about 40%, in the Norwegian consumer price index (CPI) of prices on imported goods and domestically produced goods exposed to foreign competition.

Secondly, Norway has a system of centralised wage bargaining which may speed up the adjustment of a real exchange rate towards its equilibrium rate, see e.g. Calmfors and Driffill (1988). For instance, the central wage bargainers may lower their wage claims to absorb adverse shocks to the profitability of the sector for tradables, and thereby restore its competitiveness relative to abroad and the domestic sector for non-tradables. The essay elaborates on this and shows that a centralised wage bargaining system exerts stabilising pressure on the real exchange rate process. Our empirical results show that domestic prices contribute to convergence towards PPP, as implied by this argument.

Thirdly, Norway has mainly pursued a policy of exchange rate stabilisation against western European countries since the end of Bretton Woods system in 1971. It is known that real exchange rates often follow random walks in samples from floating nominal exchange rate regimes, but tend to display mean reverting behaviour in samples from stable nominal exchange rate regimes, see e.g. Mussa (1986) and Stockman (1983). This non-neutrality property of nominal exchange rate regimes is often considered as an anomaly and rarely explained. However, we argue that one explanation for this observation could be that central banks often determine the central parity in the light of the PPP theory, see e.g. Cassel (1922), Officer (1976), Isard (1995) and Hinkle and Montiel (1999).

Furthermore, Norway has undertaken a large number of devaluations to correct for the weakening of the competitiveness of the economy due to excessive wage and price growth, see e.g. Norges Bank (1987), Skånland (1983) and Rødseth and Holden (1990). These devaluations may partly account for the rapid convergence of the real exchange rate towards the long run equilibrium.

In addition, concern for the viability of the sector for tradables seems to have been the guiding principle for the Norwegian governments income policies and for its attempts to influence the wage and price growth, especially since the discovery of North Sea oil. The government has even resorted to legal actions to contain growth in wages and prices in far excess of that in its trading partners, see e.g. Rødseth and Holden (1990) and

Alexander et al. (1997). It seems reasonable that such policies contribute to stabilise the real exchange rate.

We argue that such additional features of an economy need to be taken into account when explaining cross-country differences in the persistence of real exchange rates.

3. A common approach and similar results

The thesis consists of essays that investigate the behaviour of different economic variables using a wide range of techniques and models. Nevertheless, these essays are interrelated in a number of ways, e.g. in the motivation of non-linearities, modelling approach and results. This section relates the essays to each other. We start by outlining the modelling approach that forms the basis for the empirical analyses in the essays. Thereafter, we point out the common features of the essays in a more specific way.

3.1. Modelling approach

This subsection sketches the reductions which are implicitly or explicitly made when formulating empirical models of the data generation process (DGP) of a series. It will emerge that operational models are defined by a sequence of data reductions, where the order of reductions does not matter. The subsection also introduces the linear and non-linear models that have been employed in this thesis. In addition, the subsection presents criteria for evaluating the derived models. The criteria are also motivated by the reduction steps. The outline of the theory of reduction and evaluation criteria draws extensively on Hendry (1995).

Let $\{u_t\}$ denote a stochastic process where u_t is a vector of n measurable random variables relevant to the economy under investigation over $t = 1,..., T$. The data generation process of $\{u_t\}$ can be written as

$$D_U(U_T^1|U_0;\ \psi) \text{ with } \psi \in \Psi \subseteq \Re^k \tag{3.1}$$

where $D_U(.)$ denotes the conditional density function of $U_T^1 = (u_1, ..., u_T)$ conditional on U_0,which denotes the initial conditions/observations. $D_U(.)$ is defined by a k-dimensional vector of parameters $\psi = (\psi_1, ..., \psi_k)'$ with parameter space $\Psi \subseteq \Re^k$.

Since U_T^1 is unmanageably large, one may consider an aggregate $W_T^1 = (w_1, ..., w_T)'$ in order to reduce its dimension. The reduction entails no loss if there is one-one mapping between micro and macro economic variables $w_t : U_T^1 \longleftrightarrow W_T^1$. Then the DGP of W_T^1, that is, the conditional density function of W_T^1 on W_0 is:

$$D_W(W_T^1|W_0;\ \phi)\ = D_U(U_T^1|U_0;\ \psi) \text{ with } \phi \in \Phi \subseteq \Re^k \tag{3.2}$$

Still, the dimension of W_T^1 may be large. Marginalisation with respect to a part of W_T^1 can reduce its dimension to a manageable level. Accordingly, W_T^1 can be partitioned into two submatrices e.g. X_T^1 and $V_T^1 : W_T^1 = (X_T^1 : V_T^1)$, in the light of the purpose of analysis which could be to draw inference on μ. Without any loss of information, $D_W(.)$ can be factorised as follows:

$$D_W(W_T^1|W_0;\ \phi) = D_{V|X}(V_T^1|X_T^1,\ W_0;\ \Lambda_a)D_X(X_T^1|\ W_0;\ \Lambda_b) \tag{3.3}$$

Then, if everything about μ can be learnt from X_T^1, V_T^1 is superfluous to the analysis. In which case V_T^1 can be neglected by focusing on the marginal density of X_T^1, $D_X(.)$, i.e. the DGP of X_T^1.

The error term is derived by sequential factorisation of the DGP of X_T^1 :

$$D_X(X_T^1|\ W_0;\ \Lambda_b) = \prod_{t=1}^{T} D_x(x_t|X_{t-1}^1,\ W_0;\ \lambda). \tag{3.4}$$

The error $\epsilon_t \equiv x_t - E(x_t|X_{t-1}^1, W_0; \lambda)$ is a mean innovation process with respect to the information set, i.e. $E(\epsilon_t|X_{t-1}^1, W_0; \lambda) = 0\ \forall\ t$, by construction; E denotes the expected value of a variable. Hence, ϵ_t is also unpredictable from its own past values ϵ_{t-i}, where $i \geq 1$.

The information set can however be rather large as e.g. X_{t-1}^1 includes all lagged values of X down to x_1 and increases with t, since $X_{t-1}^1 = (x_1, ..., x_{t-1})'$. However, there is no loss of information under valid lag truncation:

$$D_x(x_t|X^1_{t-1};\ \lambda) = D_x(x_t|X^{t-p}_{t-1},\ W_0;\ \lambda), \qquad \forall\ t \geq p+1 \tag{3.5}$$

where p denotes the largest relevant lag and $X^{t-p}_{t-1} = (x_{t-p}, ..., \ x_{t-1})'$. Lag truncation reduces the information set, but does not entail loss of information if ϵ_t remains a mean innovation process.

The empirical analysis aimed at drawing inference on the parameters of interest μ may be simplified further by considering the conditional process of y_t on z_t, instead of the process of x_t where $x_t^{'} = (y_t' : z_t')$. The density of x_t can be factorised as:

$$D_x(x_t|X^{t-p}_{t-1},\ W_0;\ \lambda) = D_{y|z}(y_t|z_t,\ X^{t-p}_{t-1},\ W_0;\ \lambda_1)D_z(z_t|X^{t-p}_{t-1},\ W_0;\ \lambda_2), \qquad \forall t \tag{3.6}$$

For the purpose of drawing inference on the parameter(s) of interest contained in μ, there is no loss of information from analysing only $D_{y|z}(.)$ if: (1) μ is a function of $\lambda_1$alone and (2) $(\lambda_1,\ \lambda_2) \in \Lambda_1\times\ \Lambda_2$, that is, the parameter space (of λ_1) Λ_1 is independent of λ_2 and Λ_2 is independent of λ_1. z_t is said to be weakly exogenous with respect to μ when the conditions (1) and (2) are fulfilled.

In addition to reducing the complexity of the analysis, inference on the basis of $D_{y|z}(.)$ alone may be more robust to misspecifications of $D_z(.)$.

Often a transformation of x can have more desirable distributional properties than x, as being normal and homoscedastic. Let us map x_t into $x_t^* = J(x_t)$, then

$$D_x(x_t^*|X^{*\ t-p}_{t-1},\ W_0;\ \lambda) = D_x(x_t|X^{t-p}_{t-1},\ W_0;\ \lambda), \qquad \forall t \tag{3.7}$$

where $X^{*\ t-p}_{t-1} = (x^*_{t-p}, ..., \ x^*_{t-1})'$ and $x_t^{*\ '} = (y_t^{*\ '} : z_t^{*\ '})$. The derived error inherits the distributional properties of x_t^*. Thus e.g. $\epsilon_t^* \equiv x_t^* - E(x_t^*|X^{*\ t-p}_{t-1},\ W_0;\ \lambda)$ which will, in addition to being a mean innovation process, be approximately normal and homoscedastic:

$\epsilon_t^* \sim N(\mathbf{0}, \Sigma)$. Commonly $x_t^* = \ln(x_t)$, but each of the variables in x may be transformed differently; for example, $y_t^* = \mathbf{h}(y_t)$ while $z_t^* = \mathbf{g}(z_t)$.

More importantly, the aim of considering transformations of y_t and z_t could be to obtain a proper characterisation of the relationship between y and z. For example, a linear relation between y and z may seem to be inconsistent with economic theory and/or be rejected statistically against a non-linear relation. In such cases a linear relation between transformations of y_t and z_t may turn out to be more adequate. If we focus on $D_{y|z}(.)$, a linear relation between transformations of y and z, $\mathbf{h}(y_t)$ and $\mathbf{g}(z_t)$, can be presented in the model form as follows:

$$A(L)h(y_t) = B(L)g(z_t) + \varepsilon_t \tag{3.8}$$

where $A(L)$ and $B(L)$ are polynomial matrices of order p in the lag operator L and ε_t is the derived error. Here the polynomial matrices are defined by constant parameters and the model is linear in the transformations of y and z, which need to be specified in order to make the model operational.

Sometimes a linear relation between $\mathbf{h}(y_t)$ and $\mathbf{g}(z_t)$ can be formulated as a non-linear relation between y and z, e.g. as a relation between y and z that is characterised by state dependent parameters. Below, $A_{s_t}(L)$ and $B_{s_t}(L)$ are state dependent polynomial matrices of order p in the lag operator; subscript s_t denotes the state at time t.

$$A_{s_t}(L)y_t = B_{s_t}(L)z_t + \varepsilon_t \tag{3.9}$$

Achieving proper characterisation of the relationship between the endogenous and conditional variables (e.g. y and z) is among the main concerns of this thesis. In other words, we seek an adequate specification of e.g. $E(y_t|z_t, X_{t-1}^{t-p}, W_0; \lambda)$. In the following we introduce different models specifications used in the thesis. For this purpose, it is convenient to work with the distribution of x in model form. Even though the different essays use log transformation of the endogenous and conditional variables in most cases, here we assume that no transformation is needed to render normally distributed x.

A vector autoregressive (VAR) model of x, which can be considered as a vector of n variables, can be formulated as follows

$$\begin{aligned} x_t &= \pi_1 x_{t-1} + \pi_2 x_{t-2} + + + \pi_p x_{t-p} + \varepsilon_t, & (3.10) \\ \varepsilon_t &\equiv x_t - E(x_t | X_{t-1}^{t-p};\ \boldsymbol{\lambda}) & (3.11) \\ \varepsilon_t &\sim IIDN(\mathbf{0},\ \Sigma);\ t = 1,\ 2,\ ...T. & (3.12) \end{aligned}$$

Greek letters without subscript t denote parameter matrices; ε_t is a vector of n identically, independently distributed errors with normal distribution: IIDN($\mathbf{0}$, Σ). [3]Equation (3.10) specifies a linear model of x.

The VAR model can be reformulated as a vector equilibrium correction model (VEqCM):

$$\begin{aligned} \Delta x_t &= -(I_n - \sum_{i=1}^{p} \pi_i) x_{t-1} - \sum_{i=2}^{p} \pi_2 \Delta x_{t-1} - \sum_{i=3}^{p} \pi_i \Delta x_{t-2} - - - \pi_p \Delta x_{t-p-1} + \varepsilon_t \\ \Delta x_t &= -(I_n - \sum_{i=1}^{p} \pi_i) x_{t-1} - \sum_{j=1}^{p-1} \sum_{i=j+1}^{p} \pi_i \Delta x_{t-j} + \varepsilon_t \\ &= \pi x_{t-1} - \sum_{j=1}^{p-1} \sum_{i=j+1}^{p} \pi_i \Delta x_{t-j} + \varepsilon_t & (3.13) \end{aligned}$$

If the variables are integrated, π has to be of reduced rank (r), $r < n$, in order to define a (vector) equilibrium correction model.[4] The reduced rank matrix π can be decomposed into $\alpha\beta'$ where β defines the cointegration vector making $\beta' x$ a vector of r cointegrating

[3]The assumption of normally distributed errors is made for the analysis of the likelihood function, though it is not needed for the asymptotic analysis if the Central Limit Theorem holds.

[4]Intuitively, a variable w_t is called integrated of order zero, I(0), if it does not behaves as the cumulation of all past perturbations. A variable w_t is called integrated of order 1, I(1), if Δw_t is integrated of order zero, I(0).

w is called cointegrated if it is e.g. I(1) but a linear combination defined by $\beta' w$ is I(0). β is referred to as the cointegrating vector. See e.g. Banerjee et al. (1993) for details.

relations.

$$\Delta x_t = -\alpha\beta' x_{t-1} - \sum_{j=1}^{p-1} \Gamma_j \Delta x_{t-j} + \varepsilon_t \tag{3.14}$$

α is a n x r matrix containing weights of the r cointegrating relations in the n equations of the system. The number of r cointegrating relations can be determined by following the procedure developed in Johansen (1988), which suggests maximum likelihood criteria for determining r and estimates π for a given choice of r. The procedure also provides estimates of α and β that satisfy the restriction $\pi = \alpha\beta'$. However, this restriction does not lead to unique estimates of α and β as $\alpha\beta' = \alpha\xi\xi^{-1}\beta' = \pi$, where ξ is any non-singular r x r matrix. However, subject matter theory can be used to achieve an identified and interpretable β. The estimates of α are obtained conditional on the identified β. Furthermore, tests of zero restrictions on α can be conducted to draw inference on whether a given variable is weakly exogenous for β and their associated weights in a conditional model, i.e. parts of α.

A non-linear cointegrated VEqCM can be formulated as follows:

$$\Delta x_t = -\alpha(s_t)\beta' x_{t-1} - \sum_{j=1}^{p-1} \Gamma_j(s_t) \Delta x_{t-j} + \varepsilon_t(s_t). \tag{3.15}$$

Here parameter matrices are made functions of state s_t. That is to say that their values depend on the realised value of s at time t.[5] Regarding s, it can be specified in numerous ways: be unobservable or observable and take on values in a discrete or continuous space. Also, its value can be made dependent on stochastic or deterministic variables as e.g., time. Furthermore, the form of the function relating s_t to the stochastic or deterministic variables can be specified in many ways. The choice between the numerous specifications of s_t can be narrowed down by relying on economic theory, graphical analyses of the relevant series and/or formal statistical tests where the null hypothesis of a linear specification is

[5]Cointegration entails Granger causality in at least one direction, i.e. α is different from zero fully or partially, see Granger (1986). However, when α is made state dependent, Granger causality may change from one state to another. Krolzig (1997) extends the VEqCM to Markov Switching VEqCM and contains a discussion of the issues involved.

tested against the alternative of a specific non-linear form.

In this thesis, s_t is specified either as a unobservable state variable that takes on values in a discrete space, 1 or 2, or as logistic function of a chosen observable variable, which is denoted as the transition variable, see Hamilton (1989), Krolzig (1997) and Teräsvirta (1994) and (1998). In the former case, the evolution of the unobservable s_t is governed by a first order Markov chain with transition probabilities $\{p_{ij}\}_{i,j=1,\,2}$.[6] Accordingly, the probability of being in a certain state at time t depends on the value of s at time *t-1*:

$$P\{s_t = j \mid s_{t-1} = i,\ s_{t-2} = k, ...\} = P\{s_t = j \mid s_{t-1} = i\} \equiv p_{ij}, \tag{3.16}$$

with $p_{i1} + p_{i2} = 1$ for $i = 1,\, 2$. The transition probability p_{ij} denotes the probability of state j conditional on the economy being in state i in the previous period.

Since s is unobservable, probabilistic measures are used to draw inference on its value in a given period. The probabilistic inference about the value of s_t is made conditional on the history of x and the estimated value of the parameter vector Θ, where: $\Theta = (\alpha_1',\ \alpha_2',\ \beta',\ \Sigma_1,\ \Sigma_2,\ \Gamma_{1,1}\,,\ \Gamma_{1,2}\,,\ \Gamma_{2,1}\,,\ \Gamma_{p-1,1}\,,\ \Gamma_{p-1,2}\,,\ p_{11},\ p_{12},\ p_{21},\ p_{22})'$, i.e.,

$$P(s_t = j \mid x_t,\ x_{t-1},\ x_{t-2}, ...,\ x_1;\ \widehat{\Theta}) \qquad ; j = 1, 2. \tag{3.17}$$

The probability of being in regime j at time t given the observed data and the estimated value of Θ, $\widehat{\Theta}$, is called *filtered probability.* In contrast to filtered probabilities, *smoothed probabilities* are calculated by using the whole sample, as indicated below:

$$P(s_t = j \mid x_T,\ x_{T-1},\ x_{T-2}, ...,\ x_1;\ \widehat{\Theta}) \qquad ; j = 1, 2. \tag{3.18}$$

Both filtered and smoothed probabilities are calculated for every date in the sample and are useful in dating the transition(s) between the regimes in the series. Maximum

[6] Explicit conditioning on s_t is needed in e.g. (3.7), if it is not included in the set of conditioning variables X_{t-1}^{t-p}.

likelihood estimates of Θ and the probabilities are obtained by iterations between (preliminary) estimates of Θ and those of the probabilities, using the Expectation Maximisation (EM) algorithm, see e.g. Hamilton (1990) and Krolzig (1997) for details.

Alternatively, s_t can be specified as a logistic function of a transition variable k_{t-d} in the following way:

$$s_t = (1 + \exp[-\gamma\{k_{t-d} - c\}])^{-1}, \qquad \gamma > 0;\ d \geq 0. \tag{3.19}$$

Here γ is denoted as the transition parameter and it determines the speed of transition between 0 and 1, which are the two extreme values in the continuous 0-1 space, see e.g. Teräsvirta (1994) and Teräsvirta (1998) for details. c is referred to as the threshold value. For sufficiently large values of γ, s makes a rapid transition from 0 to 1 or vice versus on $k_{t-d} - c \neq 0$. In that case, s resembles a step function that is either 0 or 1.

In order to allow for asymmetric response to positive and negative changes in the forcing variables and particularly to positive and negative deviations from equilibria, represented by $\beta' x$, one may consider the following generalisation of the non-linear VEqCM with state dependent effects, (3.15).

$$\Delta x_t = -\alpha^+(s_t)\beta' x^+_{t-1} - \alpha^-(s_t)\beta' x^-_{t-1} - \sum_{j=1}^{p-1}\{\Gamma^+_j(s_t)\Delta x^+_{t-j} + \Gamma^-_j(s_t)\Delta x^-_{t-j}\} + \varepsilon_t(s_t). \tag{3.20}$$

Here superscript "+" denotes that $x^+ = x$ iff $x \geq 0$ while $x^+ = 0$ iff $x < 0$; similarly, x^- $= x$ iff $x \leq 0$ while $x^- = 0$ iff $x > 0$. The model allows for both state dependent and sign dependent effects of the right hand side variables on Δx_t.

Model evaluation Essentially, a model is evaluated through retracing the reductions and checking whether a given reduction represents a loss of information against a more general model formulation. Specifically, it may include testing for the null hypotheses of errors being IIDN(.) against the available information, correct functional form and parameter constancy over time. A model is considered as misspecified if some of the null hypotheses are rejected.

Model misspecification is dealt with by extension of the information set, by e.g. expanding the dimension of x; increasing the lag length p; taking account of relevant structural and institutional factors; and/or by altering the way different variables enter the model, i.e. reconsidering the mapping of some elements of x into the corresponding elements of x^*. Even a reduction in the dimension of the vector x may be considered if some of the variables in this vector turns out to be difficult to characterise parsimoniously. In that case the problematic variable is conditioned on, that is, one consider a conditional model of y_t on z_t, where (vector) z_t consists of the problematic variable(s) and y_t of the remaining variables in x. Statistically, however, conditioning is only warranted if the problematic variables are weakly exogenous with respect to parameters of interest. An explicit test of the weak exogeneity of a conditional variable requires that the marginal model of the conditional variable is derived.

Also, models are evaluated against economic theory. It is required the results are interpretable in the light of economic theory. Furthermore, when there are rival models of the same process, it is required that the preferred model encompasses the rival models. Intuitively, a model M_1 encompasses another model M_2 if M_1 is able to explain the results/properties of M_2, see e.g. Hendry (1995) and the references therein.

3.2. Time series of macro economic variables

The thesis employs time series of macro economic variables and models the behaviour of aggregate variables of the Norwegian economy. At the relatively most disaggregated level, Essay 2 models the employment in Norwegian manufacturing and construction. At the most aggregated level, Essay 3 and 4 use all foreign net financial investment in Norway as an explanatory variable when analysing the behaviour of the Norwegian trade weighted nominal exchange rate. The Norwegian exchange rate and the total unemployment rate employed in e.g. Essay 1 indicates the level of aggregation in-between the two extreme levels.

All of the four essays use seasonally non-adjusted quarterly observations and the samples span the period from early 1970s to 1996/97. In addition to a quarterly data set, Essay 3 employs daily observations over the period 1986-1998 and monthly observations over the period 1990-1998. Furthermore, Essay 3 and 4 utilise the same quarterly data set when analysing the Norwegian nominal exchange rate.

Given the apparently non-stationary behaviour of most of the macro economic variables, formal tests of their degree of integration are conducted to avoid spurious results due to unbalanced models, i.e. models where the left hand side variables and the right hand side variables (or terms) are integrated of different orders, see e.g. Banerjee et al. (1993). The order of the integration of right hand side variables and terms are matched with that of the endogenous variables through differencing or cointegration. When examining the degree of integration, we employ standard augmented Dickey-Fuller (ADF) tests, cf. (3.13). The variables are either assumed to be integrated of order 1 or of order 0. Inference on whether or not variables cointegrate and on the cointegration vector is either based on single equation models with a theory based cointegration vector or on the system based procedure suggested in Johansen (1988). In the former case, we test the statistical significance of e.g. βx_{t-1} in a single equation equilibrium correction model under the null hypothesis of no cointegration, using the non-standard critical values in MacKinnon (1991).

The empirical analyses are conducted using PcGive 9.10, PcFIML 9.10, MSVAR 0.99 for Ox2.10 and Gauss code kindly provided by James Hamilton, see Hendry and Doornik (1996), Doornik and Hendry (1997) and Krolzig (1998). All graphs are produced using GiveWin 1.30, see Doornik and Hendry (1996).

3.3. Models employed in the different essays

The models employed in the different essays can be derived as special cases of the models presented above. The model choice is based on the purpose of the analysis and on the objective of deriving data consistent models, as suggested below.

In Essay 1 we employ a univariate framework when testing for multiple equilibria in Norwegian unemployment, i.e. $x_t = U_t$ where U_t is the unemployment rate. We start out with an equation similar to (3.10) implying unique unemployment equilibrium or hysteresis in the unit root sense, depending on whether or not $|(\sum_{i=1}^{p} \pi_i - 1)| < 0$ or $= 0$.

To allow for a discrete number of equilibria, we generalise the initial model with s_t defined as an unobservable state variable governed by a first order Markov process in a binary space. However, the implied model turns out to be too sensitive to model specification. Therefore, we consider the case where s_t is formulated as a logistic function of U_{t-1}, i.e. we apply a logistic smooth transition autoregressive (LSTAR) model. The LSTAR turns out to have more desirable statistical properties than the linear and the Markov regime switching models. Moreover, it provides easily interpretable results.

In Essay 2, we employ a conditional model of y on z. Here y is a vector of Norwegian total unemployment rate and employment and hours in Norwegian manufacturing and construction. z is a vector of conditional variables as indicators of aggregate demand, unit labour costs and factors affecting labour supply. In this essay we employ the whole range of models presented in equations (3.10)−(3.20). First we define equilibrium employment, unemployment and hours by following Johansen (1988), and derive VEqCMs, cf. (3.10). Thereafter, we allow for state dependent dynamic adjustment towards the equilibrium employment and unemployment and to changes in the explanatory variables. To test whether perceived shortage of labour can lead to self enforcing equilibria, the constant term in the employment and unemployment equations are made dependent on s_t, which is defined as a logistic function of the unemployment rate, cf. (3.19). Thereafter allowance for state dependent parameters is made by making all the parameters dependent on an unobservable s, which follows a first order Markov process, cf. (3.16). The implied model is a generalisation of Hamilton (1989) to the case of a vector equilibrium correction model conditional on a number of explanatory variables. However, we are only able to estimate the model equation by equation, and not as a system.

In Essay 3, we employ single equation equilibrium correction models for the Norwegian nominal exchange rate. The equilibrium correction term is defined in the light of the PPP theory and the empirical evidence in Essay 4. The focus of the study is whether a linear representation of oil price effects in the exchange rate model can lead to underestimation of the oil price effects on the exchange rate and thereby a failure to explain major fluctuations in the exchange rate. Given the focus of the study, only the oil price effects are made dependent on the state variable s_t. In this essay we only consider logistic functions of oil prices in USD. The suggested specifications of the logistic functions are implied by a close look at the daily observations of oil prices and the nominal exchange rate

In Essay 4, we commence the analysis with a univariate model of the real exchange rate similar to (3.13) when testing whether it is mean reverting or not. Thereafter a VAR model with x as 7 x 1 vector is used to analyse the long run properties of the nominal exchange rate, domestic and foreign prices and interest rates and net financial investment relative to GDP. Since the focus of the study is to highlight the long run properties, only models with constant parameters are considered. That is we do not need to formulate models similar with state dependent parameters, or trivially $s_t = 1 \; \forall \; t$ in Essay 4.

3.4. Derivation of models

We follow a general to specific modelling strategy when deriving our preferred linear models. To be more specific, we start out with quite general linear models containing several lags of variables, which are commonly considered to be of relevance on the basis of previous empirical and theoretical studies. The general models, which are of the linear form as (3.10) or (3.14), are thereafter simplified on the basis of data consistent exclusion restrictions, in particular.

Tests for possible state dependence in the parameters are based on the parsimonious linear models. In the light of the outcomes of tests, the parsimonious linear models are generalised to allow for state and/or sign dependent effects. Thus one could consider this as a specific to general modelling strategy.

The main reason for adopting a specific to general modelling strategy when deriving non-linear models is that the number of parameters increase substantially when they are allowed to be e.g. state dependent. Thus given the typical sample size of around 100 monthly or quarterly observations employed in this thesis, the uncertainty associated with the results may rise and lead to less powerful tests of null hypotheses. Moreover, since the non-linear models are estimated using numerical methods, at least initially, it becomes difficult to obtain rapid convergence, if at all. Nevertheless, when possible, we test for the relevance of variables that appear statistically insignificant en route to the parsimonious linear models. However, they turn out to be statistically insignificant in the non-linear models. On the other hand, some of the variables that are statistically significant in the linear models become statistically insignificant in the non-linear models, at least in one of the states. To ease comparison across states and with the linear models, the insignificant terms and variables are retained in the non-linear models. Thus, one can achieve more parsimonious versions of the non-linear models than presented in the essays.

3.5. Evaluation of models

The thesis places great emphasise on the evaluation of the derived models, particularly on testing the assumptions made about the error properties, functional form and the stability of parameters. Tests of the assumptions regarding the errors are undertaken systematically. Specifically, we test the null hypotheses of no-autocorrelation, no-heteroscedasticity and that of no violations of the normality assumption. We employ the RESET test to whether the functional form is adequate, see Ramsey (1969). Upon rejection of the null hypothesis of correct functional form, we test whether the rejection could be due to misrepresentation of certain explanatory variables in the model, cf. Teräsvirta (1998). The functional form is revised in the light of the outcome of these tests. Furthermore, the properties of the linear models are compared with those of the preferred non-linear models, cf. encompassing. The assumption of parameter stability over time is tested by recursive OLS estimation of the derived models. Moreover, by one-step ahead residuals and Chow tests at different horizons, see Chow (1960).

Also, the main conclusions are examined within alternative model specifications. The purpose of the extensive evaluation of the models and the implied results is to convince ourselves that the results presented in this thesis are not based on transient correlations between variables but represent relatively stable features of the economy.

For example, Essay 1 employs both the Markov regime switching framework and an LSTAR model to test for multiple equilibria. Moreover, the conclusions from the Markov regime switching model are evaluated within different specifications of the model. As they turn out to be sensitive to model specification we consider both an ESTAR and an LSTAR model to characterise the unemployment behaviour. However, formal tests prefer the LSTAR model and it turns out to be well specified providing data consistent and easily interpretable results.

In Essay 2, we consider whether cycle independent but asymmetric costs of hirings and firings could explain the obtained results, which seem to provide strong evidence for cycle dependent employment response to shocks. Still the evidence clearly suggests cycle dependent employment response to shocks.

In Essay 3, we analyse daily, monthly and quarterly observations covering different time periods and policy regimes to convince ourselves that the obtained results regarding the non-linearity of oil prices are not just artefacts of a given sample, model and/or regime specific. Furthermore, we evaluate the preferred model against variables not used in the derivation of the model; reestimate the model on extended data set and compare its predictive power against an alternative model with a deterministic account of the non-linearity. The results are supported across the different samples and the models.

In Essays 1-3, we evaluate the results derived on the basis of non-linear models against corresponding linear models. It is shown that the preferred non-linear models outperform the linear models in terms of model properties and interpretability.

In Essay 4, we employ a full VAR model to test the results that follow from univariate models of the real exchange rate. Moreover, the main results are re-examined within 10 different specifications of the full VAR model. Additionally, the main results are tested within relatively parsimonious conditional VAR models. The results turn out to be consistent across the whole range of models.

3.6. Market imperfections and institutional factors

In this thesis non-linearities and delays in adjustment towards a given equilibrium are motivated by market imperfections and/or institutional behaviour. Essay 1 refers to coordination failures, imperfect information and asymmetric adjustment costs to explain the presence of multiple equilibria and asymmetric response of unemployment to positive and negative shocks. Essay 2 ascribe state dependent employment adjustment and multiple equilibria to actual and perceived costs of hirings and firings. Also, possible asymmetric response of employment to positive and negative shocks is motivated by asymmetric costs of hirings and firings. Furthermore, a gradual decentralisation of wage bargainings is presented as a possible explanation of a rise in the unemployment rate in equilibrium.

Essay 3 offers an argument for a possible non-linear relation between oil prices and the nominal exchange rate based on the behaviour of a central bank in the face of higher costs of resisting depreciation pressure than appreciation pressure, especially if it is more concerned with the "side effects" of e.g. high interest rates on the activity level than on inflation. Essay 4 ascribes the relatively rapid elimination of deviations from purchasing power parity between Norway and its trading partners to (i) a relatively centralised system of wage bargaining and (ii) to fiscal and monetary policies in Norway that seem to have been guided by the concern for the competitiveness of the exposed sector. It is argued that such institutional factors should be taken into account when explaining the speed of adjustment towards the equilibrium level of the real exchange rate, rather than interpreting it as merely a reflection of arbitrage pressure between trading countries.

3.7. Non-neutrality of monetary shocks

The thesis does not restrict monetary or demand shocks to have transitory effects on real economic variables as the unemployment rate and employment. In Essay 1, a sufficiently large shock or a sequence of small shocks, irrespective of their source, are allowed to shift the unemployment from one equilibrium to another and thus have permanent effects on the level of unemployment. Similarly, in Essay 2 transitory demand shocks may raise the perceived difficulty in hirings and thus make firms more reluctant to lay off workers, in which case a high employment equilibrium may become self-sustaining. In Essay 4, however, the equilibrium real exchange rate depends on the production technology and consumer preferences, which determine the weight of tradable goods and services in e.g. the consumer prices. Here monetary shocks are only allowed to bring about transitory changes in the real exchange rate. Moreover, monetary policy regime only affects the short term fluctuations in the real exchange rate. For example, a stable exchange rate policy is believed to entail a relative more stable real exchange rate than a policy of floating nominal exchange rates. To summarise, monetary shocks are allowed to affect both the short run and the long run behaviour of real economic variables.

3.8. Evidence of non-linear effects in the short run

Since non-linearities and speed of adjustment towards equilibrium are motivated by market imperfections, only the dynamic adjustment and the short term effects of forcing variables are expected to be non-linear. The long run behaviour is expected to be adequately characterised by linear static relations. Implicitly, economic agents are assumed to achieve their desired objects in the long run.

The empirical evidence across the essays turn out to be as expected. Accordingly, we do not find evidence of non-linear relations between variables in the long run, except in Essay 1. In Essay 2 and 3, there is no evidence of non-linear long run effects even though allowance is made for such effects; state dependence in the parameters only appears as short term phenomenon. Partly because of the absence of non-linear long run effects, Essay 4 employs only linear models as it is mainly concerned with the long run behaviour of the Norwegian nominal and real exchange rates.

Essay 1, 2 and 3 deal with different forms of non-linear effects but are also concerned with essentially the same type of non-linear effects. The non-linear effects can be classified in three forms: (i) Size dependent effects, (ii) State dependent effects and (iii) Sign dependent effects.

Size dependent effects of shocks are present in Essay 1. Here, whether or not a transitory shock has a permanent effect on the unemployment rate depends on its size. That is small shocks turn out to bring about temporary deviations from a given equilibrium while large shocks may entail a shift from one equilibrium to the other. Moreover, the dynamic behaviour of unemployment in the vicinity of a given equilibrium also depends on the size of the shock.

State dependent effects of shocks appear in all the three essays, though "state dependence" is defined differently. In Essay 1, the size of the shock required to make a transition and the persistence of a deviation from the equilibrium varies with the equilibrium level itself. Thus, the equilibrium level can be defined as the state of the labour market. In Essay 2, the dynamics of employment and unemployment and their response to forcing variables differ across a slack and tight labour market. In this essay, the state can defined in terms of the tightness of the labour market. In Essay 3, the state can be defined in terms of the level of oil prices, i.e. whether they are below 14 USD, above 20 USD or in-between these levels.

Sign dependent effects also appear in all the three essays. In Essay 1, negative and positive shocks are shown to bring about different dynamic behaviour of unemployment. In particular, only positive shocks in the low unemployment equilibrium may lead to shift to the high unemployment equilibrium while negative shocks only entail a transitory deviation from the low unemployment equilibrium. The case is the opposite if shocks occur when unemployment is at the higher equilibrium level. Moreover, unemployment displays asymmetric response to large positive and negative shocks, while the response is symmetric to small positive and negative shocks. In other words, unemployment recovers faster from a fall than a rise, only when the disturbances are large. In Essay 2, we test whether employment responds asymmetrically to positive and negative shocks in forcing variables and to over- and undermanning. In most cases, the response to over- and undermanning and in a number of forcing variables is symmetric to positive and negative changes. The exceptions are the response to changes in working hours and in aggregate demand which appear to be asymmetric, at least in the state of a slack labour market. The findings suggest that a reduction in employment is easier than expansion. In Essay 3, the strength of the relation between the Norwegian nominal exchange rate and the oil prices depend additionally on whether the oil price is displaying a falling or rising trend. The relation is relatively strong when oil prices are falling, but is absent when oil prices are rising. In that case, the relation is only dependent on the level of oil prices, i.e. state dependent.

3.9. Concern with possible shifts in equilibrium levels

In addition to the investigation of non-linear effects, the thesis tests for possible changes in the equilibrium levels of the endogenous variables. Considerable attention is paid to this issue in Essay 1, 2 and 4.

Essay 1 not only tests for possible shifts in the unemployment equilibrium due to transitory shocks, but also examines whether the implied equilibria have changed gradually or abruptly over time due to structural and institutional changes. Gradual or abrupt shifts in the implied unemployment equilibria are assessed by employing the test suggested in e.g. Teräsvirta (1998).

Essay 2 undertakes recursive OLS estimation of the cointegration vectors that define the equilibrium levels of employment in manufacturing and construction and of the aggregate unemployment rate to detect possible changes in the equilibrium levels. As noted above, Essay 2 also tests whether perceived ease or difficulty in obtaining the desired level of employment can lead to shift from one equilibrium level of employment to another.

Essay 4 is especially concerned with possible shifts in the equilibrium real exchange rate over time. The constancy of the equilibrium real exchange rate is examined by deriving recursive estimates of the equilibrium rate together with ± 2 times the associated asymptotic standard errors, which are calculated by applying the formula in Bårdsen (1989). Furthermore, Essay 4 discusses whether possible changes in the process determining Norwegian consumer prices can alter e.g. the equilibrium level of the nominal exchange rate process by affecting the cointegration vector defining the equilibrium rate. This is tested for by examining whether domestic prices and/or interest rates are super exogenous for the cointegration vector defining the equilibrium real exchange rate, see Engle et al. (1983). Since weak exogeneity is a prerequisite for super exogeneity, it suffices to show that the domestic prices can not be considered as weakly exogenous and hence super exogenous for the parameters defining the equilibrium nominal exchange rate.

3.10. Multiple explanations

Our findings are explainable in the light of several theories and hypotheses. For example, in Essay 1 the evidence of multiple unemployment equilibria is consistent with a number of models relying on different mechanisms that lead to multiple equilibria.

In Essay 2, we note that both cycle dependent and cycle independent hiring and firing costs can lead to observationally similar employment behaviour. The preconditions for this are that slack and tight labour markets coincide with predominance of negative and positive shocks in the employment determinants, respectively, and that the cycle independent hiring costs are greater than the firing costs. The empirical evidence however suggests that the state dependent employment response may be mainly ascribed to cycle dependent adjustment costs.

In Essay 3, the non-linear effects can be explained by central bank behaviour in the face of asymmetric costs of stabilising the nominal exchange rate in the face of depreciation and appreciation pressure. The non-linearity can also be explained by downward stickiness in prices which can lead to larger fluctuations in the nominal exchange rate in the face of depreciation pressure on the *real* exchange rate than in the face of appreciation pressure. Note that a real exchange rate appreciation can also be brought about by a rise in prices, which would alleviate the pressure on the nominal exchange rate. In the face of a depreciation pressure, however, prices are sticky downward hence a real exchange rate depreciation is mainly brought out by nominal depreciation.

In Essay 4, the relatively fast convergence of the real exchange rate to its equilibrium level can be attributed to a number of factors. Such as, relatively high arbitrage pressure due to the relatively high degree of openness of the Norwegian economy, income policies, a stable exchange rate policy with devaluations and to a relatively centralised system of wage bargaining. The thesis however does not attempt to isolate the contribution of each of the factors in explaining the evidence presented in the different essays. This is left to future research along with a number of other issues as suggested below.

4. Topics for further research

As noted above, most of the empirical evidence in this paper is explainable in the light of alternative theories and hypotheses about market imperfections and the institutional behaviour. In most cases the exact source of a given non-linear relation or of the dynamic behaviour is not identified. Thus a closer look at the possible mechanism behind the empirical regularities presented in this thesis can provide more insight into the working of an economy. Some suggestions for further research are offered below.

Essay 1 presents evidence of multiple equilibria using a univariate model, which is not capable of identifying the sources of shocks, of persistence and mechanisms which may provide scope for multiple equilibria. Studies based on data samples from other countries reach also presents evidence of multiple equilibria, albeit using univariate models, see Bianchi and Zoega (1998) and Skalin and Teräsvirta (1999). Given the plethora of explanations that can be presented for these findings, there is a need to identify the most plausible explanation. Essay 2 investigates whether perceived rationing in the labour market can explain the apparent evidence for multiple equilibria, but concludes that an explanation has to be sought in other mechanisms.

Essay 2 also suggests that one investigates the implications of state dependent employment response for the inflation process. The evidence of state dependent employment response in Essay 2 is consistent with a strongly convex upward sloping supply curve. Accordingly, the employment response to changes in forcing variables is much weaker in a tight labour market than in a slack labour market; especially when the changes lead to an upward adjustment in the employment. The evidence suggests that linear models of the employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions. Thus one could argue that linear models of inflation may overestimate the inflationary pressure of e.g. positive demand shocks in a contraction and underestimate the pressure in an expansion. This provides an avenue of further research on the inflation process using e.g. the econometric framework applied in this essay.

As argued in Essay 3, a number of studies using linear models to estimate the relation between oil prices and nominal exchange rates have suggested statistically insignificant and/or numerically weak effects of oil prices on an exchange rate. Moreover, the relation has been found to be unstable over time. In contrast to these findings, Essay 3 provides empirical evidence of relatively strong and stable effects of oil prices on the Norwegian exchange rate when a non-linear relation between oil prices and the exchange rate is allowed for. Thus one could re-examine the data sets employed in earlier studies by using non-linear models and test whether the earlier findings could be explained by misrepresentation of oil price effects on the exchange rate.

Another topic of further research is to identify the source behind the observed non-linear relation between oil prices and the nominal exchange rate. Essay 3 offers an explanation on the basis of the findings from the currency crises literature, which suggests that one is more likely to observe a relation between oil prices and the exchange rate in the face of a depreciation pressure initiated by lower oil prices than in the face of appreciation pressure triggered by higher oil prices. This asymmetric relation is likely to be present when a central bank aims to stabilise the exchange rate but is also concerned with the side effects on the activity level of policy instruments employed to stabilise the exchange rate. To substantiate this explanation one could employ data samples from countries that do not stabilise the exchange rate and/or where the central banks are commonly believed to be unresponsive to other concerns than the explicit object of the monetary policy.

Essay 4 argues that one needs to take into account institutional features of an economy in order to assess the partial effect of e.g. a nominal exchange rate regime, or centralised wage setting, on the real exchange rate behaviour. One could assess the empirical validity of this argument by using a cross country data set. One possible way to proceed would be to model e.g. the half life of a deviation from a real exchange rate equilibrium using data on the degree of centralisation and/or coordinations of wage setting, degree of openness and by taking into account the nature of fiscal and exchange rate policies of each country; the latter, possibly by taking into account the degree of central bank independence, which usually makes a central bank less disposed to undertake devaluations in order to make up for deterioration in the competitiveness of an economy. A cross country study along the sketched lines might add weight to the interpretation of results in this paper and throw more light on the empirical regularities encountered in the PPP literature.

Finally, though several extensions of the empirical results presented in this thesis have been suggested, we need to evaluate the obtained results in the light of new data accruing regularly. Moreover, new data can shed light on aspects that may have been unduly neglected and thus enable us to obtain more adequate models of the processess considered in this thesis; as Professor David Hendry often reminds us:

> "...econometric modelling is a continuing activity, not a 'one-off ' forging of empirical laws..."

2. Multiple equilibria and asymmetries in Norwegian unemployment

Abstract

This essay tests for multiple equilibria in the Norwegian unemployment rate and investigates whether it displays asymmetric response to positive and negative shocks. Linear and nonlinear univariate models are employed to account for the unemployment behaviour over the period 1972–1997. Among these, only a logistic smooth transition autoregressive (LSTAR) model is data consistent. Accordingly, unemployment is a non-integrated variable that has switched between two stable equilibria during the sample period. It is shown that a large shock, or a sequence of small shocks, can cause a transition from one equilibrium level to another and thereby have a permanent effect on the unemployment rate. Moreover, unemployment recovers faster from a fall, relative to a given equilibrium level, than from a rise.

1. Introduction

This essay characterises the Norwegian unemployment rate in a framework that allows for multiple equilibria and asymmetric responses to positive and negative shocks. This approach may explain the unprecedented rise in the unemployment rate at the end of 1980s, and shed light on its dynamic behaviour, in particular its persistence. A study of unemployment from this perspective can also address the issue of monetary neutrality. If transitory shocks e.g. due to changes in monetary policy cause a movement from one unemployment equilibrium to another, then monetary policy can have permanent effects on the level of unemployment and activity. This is in contrast to unique equilibrium models where nominal shocks only cause temporary deviations from the unique equilibrium level, see e.g. Friedman (1968) and Layard et al. (1991). The latter is assumed to be affected only by e.g. structural and institutional changes in the economy.

The essay proceeds as follows. Section 2 motivates our approach by providing the background facts and findings of previous studies. Section 3 summarises the main characteristics of multiple equilibria models and their implications for the actual (un)employment behaviour. Section 4 formalises the multiple equilibria approach in the Markov regime switching model of Hamilton (1989) and in a smooth transition autoregressive (STAR) model, see e.g. Teräsvirta (1994). These models are sufficiently general to encompass linear models that imply a unique unemployment equilibrium or hysteresis in the unit root sense. The empirical analysis is based on quarterly observations of seasonally non-adjusted data for the unemployment rate over the period 1972-1997. The series is described in Section 5.1 that lays out its the main features. These appear to be easier to reconcile with the presence of multiple equilibria than with a unique equilibrium. Subsection 5.2 derives a linear autoregressive model of unemployment and explores its properties. This model serves as a reference model and is used to evaluate the nonlinear models developed in the following sections. Subsection 5.3 models the series in the Markov regime switching framework and show the sensitivity of the results to model specification. Section 6 is devoted to the specification, estimation and evaluation of a logistic STAR (LSTAR) model. It is found that the LSTAR model is quite successful in characterising the unemployment behaviour and hence provides a sound basis to address the issues of interest. The conclusions follow in the last section, Section 7.

2. Background

The unemployment experience of (western) European countries in the last decades has revealed slow, if any, tendencies of actual unemployment to revert to a unique equilibrium rate, as prescribed by the natural rate or the NAIRU hypothesis. As noted by several authors, instead of providing an anchor for the actual unemployment rate, the estimated equilibrium level appears to track the actual rate, see for instance Layard et al. (1991), Cromb (1993) and Elmeskov and MacFarlan (1993). However, this occurs without matching structural and institutional changes. In addition, approximately the same rate or even a constant rate of price and wage inflation is observed at significantly different levels of unemployment.

To account for these findings, and in particular the apparently ratcheting behaviour of unemployment, the hypothesis of unemployment hysteresis has been launched, see Blanchard and Summers (1986) and Lindbeck and Snower (1988). The concept of hysteresis is associated with dynamic models. It denotes a situation in which the equilibrium state of a system depends on the past history of the system, see Amable et al. (1995) andRøed (1997). Since the seminal work of Blanchard and Summers (1986), however, unemployment hysteresis has been commonly equated with unemployment following a random walk. In this case, any shock, irrespective of its size and whether it is transitory or permanent, has a permanent effect on the level of unemployment. Following this notion, formal tests of unemployment hysteresis have been conducted as unit root tests in linear dynamic models of unemployment, see e.g. Alogoskoufis and Manning (1988), Cromb (1993) and Cross (1995). Most of the empirical studies of European unemployment find estimates of a root close to one, which is consistent with a null hypothesis of a unit root but also with the null hypothesis of near unit root. The latter case, termed *unemployment persistence,* is consistent with a unique equilibrium but suggests weak equilibrium reversion, cf. Bean (1994).

The practice of equating hysteresis with the presence of a unit root in a linear model has been questioned, see e.g. Amable et al. (1995), Cross (1995) and Røed (1997). Firstly, it compels every shock, irrespective of its size and sign to have a permanent effect on the level of unemployment, disregarding the existence of (endogenous) stabilising mechanisms. Secondly, long series of unemployment rate data often show that it does not wander around randomly, but revert to its past levels, sooner or later, cf. Layard et al. (1991) and Bianchi and Zoega (1997). Indeed, the bounded nature of the unemployment rate series prevents it from taking values outside the 0-1 range. Thirdly, the empirical evidence in favour of a unit root, or high degree of persistence, in relatively smaller 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 the series, see e.g. Banerjee et al. (1993). Finally, the linear models imply symmetric dynamic behaviour when unemployment rises or falls. Observations of the unemployment behaviour, however, indicate that it rises faster than it declines. Such asymmetries are often explained by asymmetric adjustment costs or by the hiring and firing practices of firms, cf. Hamermesh and Pfann (1996) and Johansen (1982).

Models of multiple equilibria appear to be capable of reconciling the empirical evidence from long and short time series and allow for more flexibility with regard to the effects of shocks, see Section 3. In these models, a large shock may cause a movement from one equilibrium level to another while small shocks only cause a temporary deviation from a given equilibrium level. These properties seem to be consistent with the findings of a number of studies, which ascribe the appearance of high degree of persistence in European unemployment series to infrequent shifts in their mean levels, see e.g. Blanchard and Summers (1987), Bianchi and Zoega (1996), (1997) and (1998).

Nonlinear models are required to entertain the possibility of multiple equilibria, i.e. to allow for asymmetric responses to the size of a shock. Theories of multiple equilibria, as well as models of unique equilibrium, are often silent on whether the response towards shocks depends on their sign or not. This property, i.e. asymmetric response to the sign of a shock, can also be incorporated into nonlinear models and tested for. A number of studies associate multiple equilibria with hysteresis, interpreted as a nonlinear phenomenon, see 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 it as a nonlinear phenomenon.

The possibility of multiple equilibria in the Norwegian unemployment rate and in a number of other OECD unemployment series has previously been tested for by Bianchi and Zoega (1996) and (1998), hereafter BZ (1996, 1998), and Skalin and Teräsvirta (1999). The findings in BZ (1996, 1998) are based on annual and quarterly data, respectively. These are consistent with the notion of unemployment swings between multiple equilibria and contradict the hypothesis of a unique equilibrium rate of unemployment for most of the OECD countries, including Norway. Indeed, the high degree of persistence in these unemployment series is accounted for by infrequent changes in the equilibrium (mean) rates of unemployment. The switching mean rates are estimated by using Markov regime switching models but with no autoregressive terms, see Hamilton (1989). The degree of persistence in each series is measured once the series are adjusted for the estimated mean rates. However, one may argue for three shortcoming with this modelling approach and the results that follow.

Firstly, the models deliver results that seems to be too dependent on the frequency of the data series. For instance, on annual data for Norway, the model in BZ (1996) suggests two equilibria at 2.1% and 5.5% and a shift to the higher equilibrium level in 1988. On quarterly data in BZ (1998), the estimates are 1.83% and 4.72%, respectively. One may say that these are roughly the same as those on annual data. However, the model implies two periods in which unemployment is at the higher equilibrium level, in 1982:3-1984:4 and 1988:1-1995:4. The shift to the higher equilibrium in the beginning of 1980s appears to be at odds with the actual data that never exceed 4% before 1988, cf. Figure 5.1. BZ (1998) recognise this as a problem with a number of other series too and deal with it outside the models to obtain interpretable results. Secondly, the results may not be invariant to the inclusion of possibly neglected dynamics in the models. The models appears to be too restrictive since they only contain regime dependent intercepts as regressors. As will be shown in Subsection 5.3.1, the estimates of mean rates and the estimated number of switches between equilibria change with the inclusion of autoregressive terms and with their numbers. Similar results have been reported by Clements and Krolzig (1998) in their study of US GNP. Finally, one may argue that the employed Markov regime switching framework is inflexible since it imposes abrupt transitions between possible equilibria at the outset. Although firms may face dichotomous decisions with regard to employment adjustment, it is unlikely that they opt for the same decision simultaneously, unless they are exposed to huge shocks. Hence at the aggregate level and in absence of huge shocks, one is more likely to observe smooth rather than abrupt transitions between possible equilibria.

Skalin and Teräsvirta (1999) employ smooth transition autoregressive (STAR) models that allow for both abrupt and smooth transitions between equilibria, see Granger and Teräsvirta (1993). They use logistic STAR (LSTAR) models for the first differences of the quarterly OECD unemployment series, but with lagged level terms and seasonal dummies. First they test the significance of the apparent asymmetric dynamics in these series, which is supported for most of the OECD unemployment series, except for Norway and a few other countries. For the Norwegian unemployment rate, and the other series that are found to not exhibit asymmetry, they go on and test for multiple equilibria. For Norway they find an abrupt switch to the high unemployment equilibrium in 1988, as in BZ (1996). However, their model does not seem to be suitable for deriving reasonable estimates of unemployment equilibria. Taken at face value, the model implies two equilibria at $0.0012/0.22 \approx 0.006\%$ and $(0.0012 + 0.64)/0.22 \approx 2.91\%$, respectively. Moreover, Skalin and Teräsvirta (1999) use the *time* trend as a transition variable to account for the multiple equilibria. Use of a *time* trend implies permanent transition to the one or the other equilibrium level. A characteristic feature of theory models of multiple equilibria is that they allow for back and forth movements between the equilibria depending on the size and sign of shocks. This feature is lost when a time trend is used as a transition variable, though it may capture slows changes in the equilibrium rates due to gradual changes in the structural and institutional features of the labour markets, as in the case of e.g. Austria and Canada, see Skalin and Teräsvirta (1999). In their analysis of Norway, however, the transition occurs abruptly in 1988.

This essay aims to derive a data consistent univariate model to test for the possibility of multiple equilibria in Norwegian unemployment and asymmetric adjustment. As

Skalin and Teräsvirta (1999) it will allow for both abrupt and smooth transitions between equilibria by adapting the STAR framework, but to model the level of unemployment. To preserve the possibility of back and forth movements between possible equilibria, the transition variable will be a lagged value of the unemployment rate.

The univariate framework adapted here and in general, does not shed light on the sources of shocks, of persistence and on mechanisms which may provide scope for multiple equilibria. However, it provides a convenient way to test for multiple equilibria and to draw out the characteristic features of the unemployment series. By this, it can provide stylised facts to be explained by theoretical and multivariate empirical models.

3. Multiple unemployment equilibria

Models that display multiple equilibria are often based on the existence of reciprocal externalities in various guises. These can arise from trading and exchange opportunities as in Diamond (1982) and Cooper and John (1985), due to spillovers of demand across markets as in e.g. Weitzman (1982) and Murphy et al. (1989) or due to costs associated with layoffs and hirings as in Saint-Paul (1995) and Moene et al. (1997b), respectively. In the presence of reciprocal externalities, the level of activity of one agent depends positively on the level of activity of another agent and vice versa. The positive feedback may potentially lead to multiple equilibria, and "coordination failure" among the agents may cause the economy to get stuck in an equilibrium with an inefficiently low level of activity and employment.

Manning (1990) shows that multiple equilibria can arise from the presence of increasing returns to labour, or in labour and capital together, in a quite standard model of an imperfect economy, for instance as in Layard et al. (1991).

Figure (a) can help us to illustrate the essence in models of multiple equilibria. Here a nonlinear price curve and a linear wage curve are sketched against the (un)employment rate, $1\text{-}u$. The price curve and the wage curve can also be interpreted as a labour demand curve and a labour supply curve, respectively. The curves intersect at two points that correspond to two stable (un)employment equilibria. Upon deviation from a given equilibrium, due to e.g. a small nominal shock, (un)employment reverts to the initial equilibrium. When exposed to a sufficiently large shock, however, the (un)employment converges towards the other equilibrium. The solid and dashed curves in Figure (c) sketch the employment response to a small and a large shock, respectively, when it is at the lower equilibrium level. The shocks may stem from changes in fiscal and monetary policy or from foreign economies through the exposed sectors.

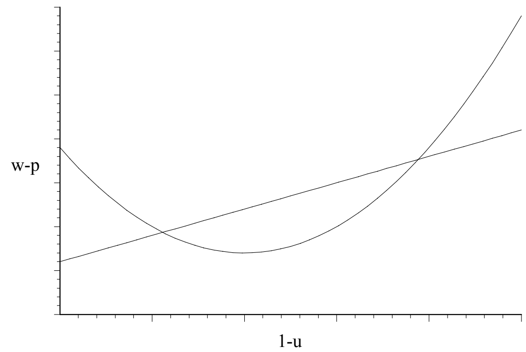


Figure (a)

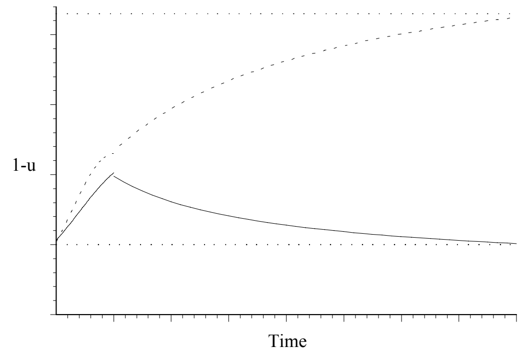


Figure (c)

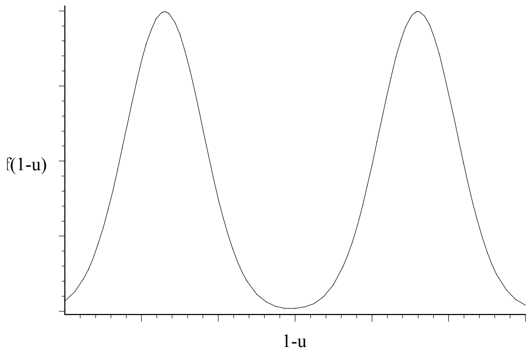


Figure (b)

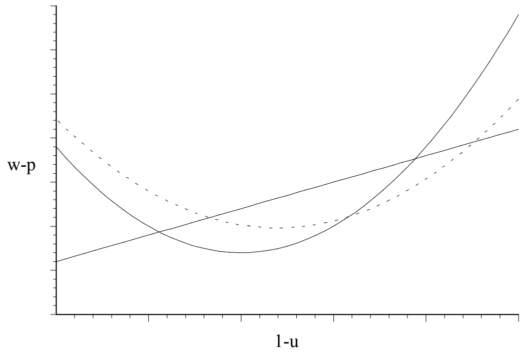


Figure (d)

The solid line in Figure (c) also indicates that employment adjusts at a lower pace towards its equilibrium than it rises. Such asymmetries are commonly ascribed to e.g. asymmetric adjustment costs. For instance, a number of studies suggest that costs of hiring usually exceed costs of separations, see Hamermesh and Pfann (1996) and the references therein. Hence employment will fall faster when exposed to a negative shock than it will rise upon a positive shock. The speeds of adjustment towards the different equilibria may also differ from each other and are dependent on the relative stability of the equilibria. The relative stability can be related to the sources of the sluggishness in the economy, e.g. to adjustment costs of different forms, whether there are nominal and real rigidities in the price-setting and or in the wage-setting and the governments policy towards unemployment, see Manning (1990). The latter study implies that the lower equilibrium is likely to be more stable if the rigidities are mainly nominal in nature and the government pursues a counter-cyclical policy.

The underlying data generating process in the labour market is unobservable but (un)employment data can provide some indications of the underlying process. If the underlying process in the labour market resembles the one sketched in Figure (a), the (un)employment rates are likely to be concentrated around two distinct (un)employment rates. In this case one may observe a bimodal density for (un)employment rates as in Figure (b). However, if the price and wage curves only intersect once, one will observe an unimodal density.

Models of multiple equilibria do not restrict the equilibria to be constant over time. They may vary following structural and institutional changes over time. Figure (d) illustrates the case where such changes shift the position of the price curve. The dotted curve indicates a movement in both of the equilibria. Tests for slow or abrupt changes in the parameters of a multiple equilibria model can provide indications of slow or abrupt movements in the equilibria.

4. Formalizing multiple equilibria

This section formalises the multiple equilibria approach in the Markov regime switching model of Hamilton (1989) and in the STAR model of Teräsvirta (1994). The linear models which imply unique equilibrium or hysteresis in the unit root sense, can be derived by imposing appropriate restrictions on these general models. The sections also touches upon ways to test these nonlinear models against the linear model.

4.1. Markov regime switching model

Equation (4.1) formalises the multiple equilibria approach in the Markov regime switching autoregressive model, hereafter MS-AR(q) model

$$U_t = \mu_{s_t} + \sum_{i=1}^{q} \phi_i (U_{t-i} - \mu_{s_{t-i}}) + \sigma_{s_t} \varepsilon_t \qquad \begin{array}{l} ; \varepsilon_t \sim N(0,1) \\ ; |\varrho = \sum_{i=1}^{q} \phi_i| < 1 \end{array} \tag{4.1}$$

The level of unemployment is assumed to evolve around a variable equilibrium level, μ_{s_t}, which takes on a finite number of values depending on the state, s, of the economy.[1] The regime or state variable s_t is assumed to be an unobserved discrete variable taking on a value in {1, 2, 3,...S}. This is governed by a first-order Markov chain with transition probabilities $\{p_{ij}\}_{i,j=1,2,...S}$. In a first-order Markov chain, the probability of state j in period t depends only on the past through the most recent value of s_t, i.e.

$$P\{s_t = j \mid s_{t-1} = i,\ s_{t-2} = k, ...\} = P\{s_t = j \mid s_{t-1} = i\} \equiv p_{ij}, \tag{4.2}$$

with $p_{i1} + p_{i2} ... + p_{iS} = 1$ for all i. The transition probability p_{ij} denotes the probability of state j conditional on the economy being in state i in the previous period. It is assumed that U_t depends only on the current and q most recent realization of s_t, cf. (4.1).

To fix ideas, consider the case of $q = 0$. When $s_t = 1$, U_t is presumed to have been drawn from a normal distribution with mean μ_1 and standard deviation σ_1, $N(\mu_1, \sigma_1)$, and from $N(\mu_2, \sigma_2)$ when $s_t = 2$, and so on. The transitions between these S regimes or distributions are assumed to be governed by a probability law specified as the Markov chain (4.2).

Shocks which do not change the state of the economy may be interpreted as affecting the level of unemployment through the epsilon, ε_t, in the expression for the residual, $\sigma_{s_t}\varepsilon_t$. In this general formulation, which allows for a state dependent standard error, the effect of such shocks on the level of unemployment may be state dependent. The restriction on the sum of autoregressive coefficients, ϱ, ensures that these shocks only cause a temporary deviation from the state dependent unemployment equilibrium.

Since we do not know in which regime the process was at every date in the sample, neither do we know the distribution responsible for delivering the observed value of U_t at every

[1] In a multivariate framework, one could interpret the model (4.1) as a reduced form equation for unemployment, derived from wage and price formation, and the term μ_{s_t} as a vector of conditional macroeconomic variables representing the state of the economy (cf. Layard *et al.*, 1991). In general, and in contrast to our purpose, the vector of conditional variables may take on an infinite number of values and imply an infinite number of unemployment equilibria. Therefor the multivariate framework would not be suitable to test for a finite number of equilibria.

date. Therefore, probabilistic inference is made about the value of s_t conditional on the history of unemployment and the estimated value of the parameter vector Θ, where: $\Theta = (\mu_1, \mu_2, ...\mu_S,\ \sigma_1,\ \sigma_2, ...\sigma_S,\ \phi_1, \phi_2, ...\phi_q, p_{11}, p_{12}, ...p_{1S},\ p_{21}, p_{22}, ...p_{2S}, ..., p_{S1}, ...p_{SS})'$, i.e.

$$P(s_t = j \mid U_t, U_{t-1}, U_{t-2}, ...U_1; \widehat{\Theta}) \qquad ; j = 1, 2, ...S \tag{4.3}$$

The probability of being in regime j at time t given the observed data and the estimated value of Θ, $\widehat{\Theta}$, is called *filtered probability.* In contrast to filtered probabilities, *smoothed probabilities* are calculated by using the whole sample. Both filtered and smoothed probabilities are calculated for every date in the sample and are useful in dating the transition(s) between the regimes in the series. However, the number of possible regimes S characterising the series has to be determined beforehand. This issue is addressed in Section 5.1.

Note that a change in any moment of the distribution can be interpreted as a change in the regime or state. For instance, two regimes may only be different from each other with respect to the variances. However, in the present context, a change in the regime or distribution will only correspond to a change in the equilibrium level of unemployment if there is a change in the mean value.

The linear autoregressive (AR) model (4.4) is a special case of (4.1) for $S = 1$, i.e. when there is only one possible regime or equilibrium.

$$U_t = \mu + \sum_{i=1}^{q} \phi_i (U_{t-i} - \mu) + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma^2) \tag{4.4}$$

In the more familiar transformation, where $\alpha = \mu(1 - \varrho)$ and $\varrho = \sum_{i=1}^{q} \phi_i$,

$$U_t = \alpha + \sum_{i=1}^{q} \phi_i U_{t-i} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma^2). \tag{4.5}$$

Note that for values of $|\varrho| < 1$, the model is dynamically stable and unemployment is a stationary variable, assuming a single unit root.[2] Thus a shock to the level of unemployment, represented by the residual, ε_t, will only cause a transitory deviation from a given unemployment equilibrium.

The special case of unemployment hysteresis in the unit root sense can be defined by $\varrho = 1$. In which case, the unemployment rate follows a random walk process, augmented by some stationary terms.[3] Thus every shock permanently changes the level of unemployment. The unemployment process does not revert to its initial equilibrium but remains at the post-shock level until disturbed by a new shock. Moreover, if an equilibrium is a state where there is no inherent tendency to change, unemployment is at equilibrium at every point in time, implying an infinite number of equilibria.

Thus one can test the hypothesis of a unique equilibrium against hysteresis or multiple equilibria by testing the null hypothesis of $\mu_s = \mu$ for all s, combined with restrictions on the sum of autoregressive coefficients.[4] The standard asymptotic distribution theory is, however, not valid under the null hypothesis of a constant parameter (linear) AR model. This is because the parameters $p_{11}, ...p_{SS}$ are unidentified, i.e. they may take on any value without affecting the likelihood value, under the null hypothesis of $\mu_1 = \mu_s$ and $\sigma_1 = \sigma_s$, see Hansen (1992) and (1996). However, Hansen (1992) and (1996) suggests a likelihood ratio test that can be used to evaluate the MS-AR(q) against the AR(q) model.

4.2. Smooth transition autoregressive model

The Markov regime switching framework imposes abrupt transitions between equilibria. This assumption can be relaxed by basing the analysis on the smooth transition autoregressive (STAR) model, see e.g. Granger and Teräsvirta (1993).

2

$$U_t = \alpha + \sum_{i=1}^{q} \phi_i U_{t-i} + \varepsilon_t = \alpha + \sum_{i=1}^{q} \phi_i U_{t-1} - \sum_{j=1}^{q-1} \sum_{i=j+1}^{q} \phi_i \Delta U_{t-j} + \varepsilon_t$$

which may be transformed to

$$\Delta U_t = \alpha - (1 - \varrho) U_{t-1} - \sum_{j=1}^{q-1} \sum_{i=j+1}^{q} \phi_i \Delta U_{t-j} + \varepsilon_t.$$

See e.g. Hamilton (1994, pp. 517) for details.

3 Remember that $\alpha = 0$ when $\varrho = 1$.

4 Note that we may have a unique equilibrium even when $\sigma_s \neq \sigma$.

A STAR model of order q can be formulated as follows:

$$U_t = \alpha + \sum_{i=1}^{q} \phi_i U_{t-i} + (\widetilde{\alpha} + \sum_{i=1}^{q} \widetilde{\phi}_i U_{t-i}) F(U_{t-d}) + \varepsilon_t, \qquad \varepsilon_t \sim N(0,\ \sigma^2) \tag{4.6}$$

where $F(U_{t-d})$ is a transition function that monotonically increases with the level of unemployment, lagged d periods, and takes on values in the range [0, 1]. The transition function, interpreted as representing the phase or state of the economy, can be specified as a logistic function

$$F(U_{t-d}) = (1 + \exp[-\gamma\{U_{t-d} - c\}])^{-1}, \qquad \gamma > 0 \tag{4.7}$$

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 constant term and the autoregressive coefficients to change with the value of $F(U_{t-d})$. Thus unemployment is allowed to evolve around distinct equilibria with different dynamics in expansions $U_{t-d} < c$ and contractions $U_{t-d} > c$. The change in parameters occurs with some delay which depends on the value of parameter d. The value of d can be determined together with the tests of a linear AR(q) model against a STAR model, see e.g. Teräsvirta (1994).

The two extreme unemployment equilibria correspond to values of $F(U_{t-d}) = 0$ and $F(U_{t-d}) = 1$ and can be defined as $\mu_1 = \alpha/(1 - \sum_{i=1}^{q} \phi_i)$ and $\mu_2 = (\alpha + \widetilde{\alpha})/[1 - \sum_{i=1}^{q}(\phi_i + \widetilde{\phi}_i)]$, respectively. The sums $\varrho_1 \equiv \sum_{i=1}^{q} \phi_i$ and $\varrho_2 \equiv \sum_{i=1}^{q}(\phi_i + \widetilde{\phi}_i)$ measure the degree of persistence in each of the two states.

Note that both a two regime autoregressive model with abrupt transitions and the AR(q) model are nested in this LSTAR model. When $\gamma \longrightarrow \infty$, the transition function $F(U_{t-d}) = 0$ for $U_{t-d} \leq c$ and $F(U_{t-d}) = 1$ for $U_{t-d} > c$. In this case the LSTAR model reduces to a self-exciting threshold autoregressive (ESTAR) model with threshold value c. The LSTAR model is reduced to an AR(q) model for $\gamma \longrightarrow 0$. In this case $F(U_{t-d}) \longrightarrow 1/2$

for all values of U_{t-d}.

The generality of the LSTAR model makes it suitable to test hypotheses concerning the number of unemployment equilibria, the degree of persistence within each equilibrium and the speed of transition between the equilibria. However, as in the case of the Markov regime switching model, the parameters defining the LSTAR model are not identified under the null hypothesis. It follows from (4.6) and (4.7) that under the null hypothesis of an AR(q) model, i.e. under $\mathrm{H}_0 : \gamma = 0$, the $\widetilde{\alpha}$, $\widetilde{\phi}_i$s' and c may take on any value without changing the likelihood value. These parameters are only identified under the alternative hypothesis of $\gamma \neq 0$. In this case, Teräsvirta (1994) suggests a sequence of tests to evaluate the null hypothesis of an AR(q) model against the alternative of an LSTAR model. The tests are based on estimating the following auxiliary regression for a chosen value of the delay parameter d

$$U_t = \beta_0 + \sum_{i=1}^{q} \beta_{1i} U_{t-i} + \sum_{i=1}^{q} \beta_{2i} U_{t-i} U_{t-d} + \sum_{i=1}^{q} \beta_{3i} U_{t-i} U_{t-d}^2 + \sum_{i=1}^{q} \beta_{4i} U_{t-i} U_{t-d}^3 + \nu_t. \quad (4.8)$$

The test of an AR(q) model against a STAR model (both LSTAR and ESTAR) is equivalent to conducting a joint test of[5]

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

The value of d can be determined by conducting the 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, i.e. 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

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

purpose, the following sequence of tests within the auxiliary regression is suggested

$$\begin{aligned} H_{04} &: \beta_{4i} = 0, \qquad i = 1, 2, ...q \\ H_{03} &: \beta_{3i} = 0 \mid \beta_{4i} = 0, \qquad i = 1, 2, ...q \\ H_{02} &: \beta_{2i} = 0 \mid \beta_{3i} = \beta_{4i} = 0, \qquad i = 1, 2, ...q \end{aligned}$$

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, see Teräsvirta (1994).

5. Data, AR(q) and MS-AR(q) models

This section offers some preliminary evidence against the unique equilibrium approach and mixed evidence in favour of two unemployment equilibria. It is organised as follows. Subsection 5.1 describes the unemployment data and plots its density to check whether its shape is consistent with multiple equilibria or not, see Section 3. Moreover, whether unemployment tends to rise faster than it declines. Subsection 5.2 derives an AR(5) model and explores its properties. In particular, whether it can provide a reasonable estimate of a unique equilibrium or not. 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(5) model will also serve as a reference model for the nonlinear models to be estimated later. Section 5.3 estimates a Markov regime switching model of order 5, MS-AR(5), model and examines its merits. Moreover, it replicates the results in BZ (1996, 1998) for Norway and demonstrate the sensitivity of results to model specification.

5.1. Data

The empirical analysis employs seasonally non-adjusted quarterly data for the open unemployment rate in Norway, see Figure 5.1.[6] The data set covers the period 1972:1-1997:1.

[6] The unemployment figures are based on the Norwegian labour force survey (AKU). The source is the database OECD_MEI and the precise variable name is MEI_SQ0220_424000A0.

During this period, there are three noticeable upswings in unemployment relative to its low level in the early 1970s, in 1974/75, 1982/83 and 1988. It moves down relatively slowly to its initial level after the upswings except for the third upswing, which is also the most striking one. Unemployment continues to rise after the strong upswing, but at a slower pace until 1992. The movement is downward after that. The figure indicates asymmetry in the pace of unemployment between the low and high level. The movements from lower to higher levels seem to be more abrupt than vice versa.

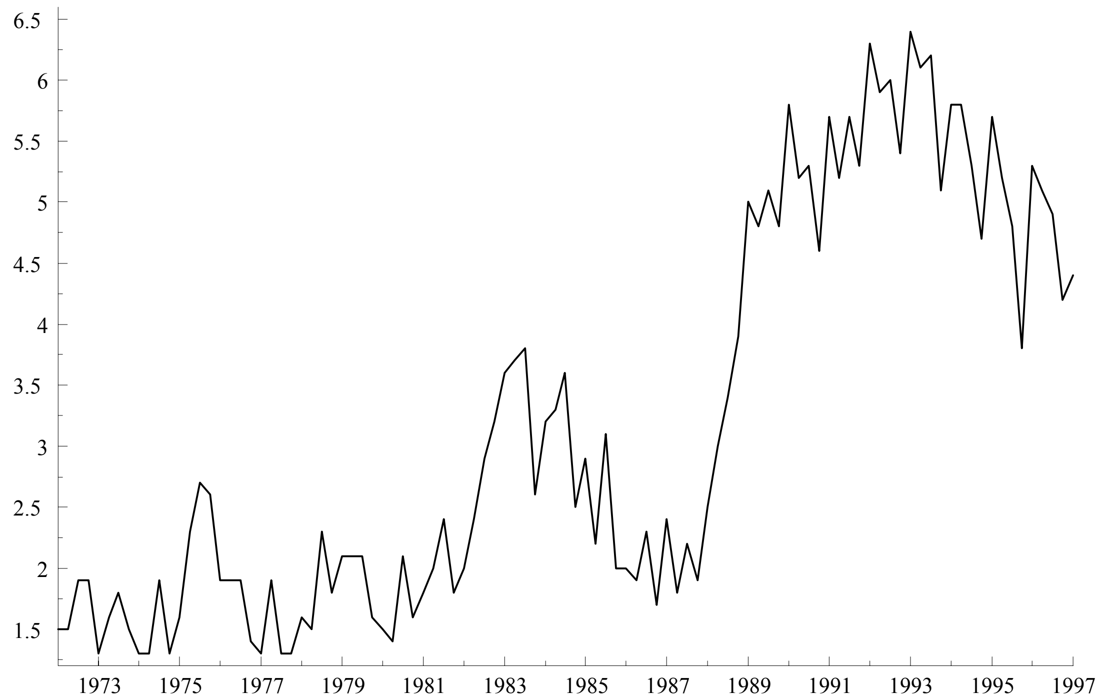


Figure 5.1: *Seasonally non-adjusted quarterly series of open unemployment in Norway, 1972:1-1997:1.*

The observed values lend themselves to interpretation in the light of the multiple equilibria or regime switching model. Figure 5.1 indicates two or possibly three different regimes in the sample. Most values of U_t in the period up to 1982 appear to have come from a "low-mean-distribution", while the values for the period after 1988 are more likely to have been drawn from a "high-mean-distribution". We are, however, more uncertain with regard to the values of U_t in the period from 1982 to 1985. Both regimes are likely to have delivered these observations. One may alternatively assume that these values are from a third distribution with a medium level of mean. Allowing for changes in the mean over the sample, there do not seem to be substantial differences in the magnitude of fluctuations between the regimes.

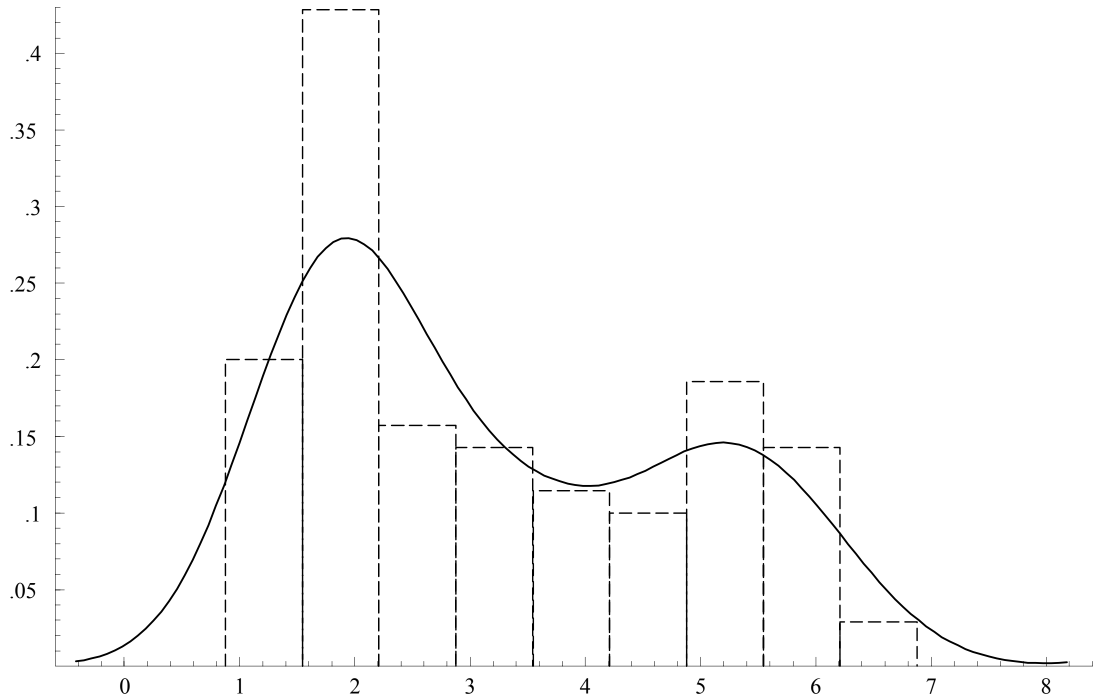


Figure 5.2: *Non-parametric density estimation of open unemployment in Norway (with histogram), 1972:1-1997:1.*

As noted in Section 3, the unemployment rates should be concentrated around two distinct rates, if they are generated by a labour market that has switched between two equilibria. Figure 5.2 plots the frequency distribution of unemployment, which appears to be bimodal with the modes centered at around 2% and 5%, respectively. The spread around the higher mode seems to be slightly larger than the spread around the lower mode. Although Figure 5.2 indicates the presence of two unemployment equilibria, one cannot preclude the existence of a third mode with a value close to either 2 or 5 but with a relatively larger or smaller spread.[7] However, in order to estimate a parsimonious model Section 5.3 assumes two possible unemployment equilibria, i.e. $S = 2$. This is in accordance with BZ (1996, 1998) who use a bootstrap method to identify the number of

[7]See section 22.3 in Hamilton (1994).

states.

5.2. An AR(5) model

The estimated AR model with five lags is presented in Table 5.1.[8] Starting with eight lags, the autoregressive order of five was determined by excluding the statistically insignificant lags of a higher order. A lag order of five was also found to be optimal according to Akaike's information criterion (AIC). Except for signs of autocorrelation in the residuals, there do not seem to be significant violations of the other standard assumptions about residual properties. There are signs of autocorrelation despite the inclusion of the second and the third lag that are insignificant at a 5% level, see the test summary in Table 5.1. The autocorrelation remained a feature of the residuals even when a lag length of 8 was used. Note that a significant value for the autocorrelation test may also result when a linear functional form is imposed on a data-generating mechanism that can be more properly characterised by a nonlinear functional form or two or more linear segments. However, the regression specification test (RESET) does not indicate significant functional form mis-specification, at least not at standard levels of significance, see Ramsey (1969). However, this test is constructed to have power against general forms of functional mis-specification and may thus have low power against specific nonlinear forms.

One-step ahead Chow tests in Figure 5.3.a indicate noticeable changes in the model parameters in 1982/83 and in 1988. The recursive OLS estimates of the constant and the first autoregressive coefficient indicate non-constancies at these dates, see Figures 5.3.b and 5.3.c. The recursively estimated standard deviations of the 1-step residuals in Figure 5.3.d, indicate a slight increase in the variance around 1982 and then a further increase in 1988. The graphical presentations above do not seem to strongly contradict the assumption of two regimes characterizing the unemployment data.

Note that the estimated value of ϱ is not significantly different from 1. As noted earlier, a value of ϱ close to 1 is a common finding in studies of European unemployment series that use the linear AR framework, thus not contradicting the null hypothesis of hysteresis in

[8] The model was also estimated with the method of maximum likelihood in order to derive the log likelihood value.

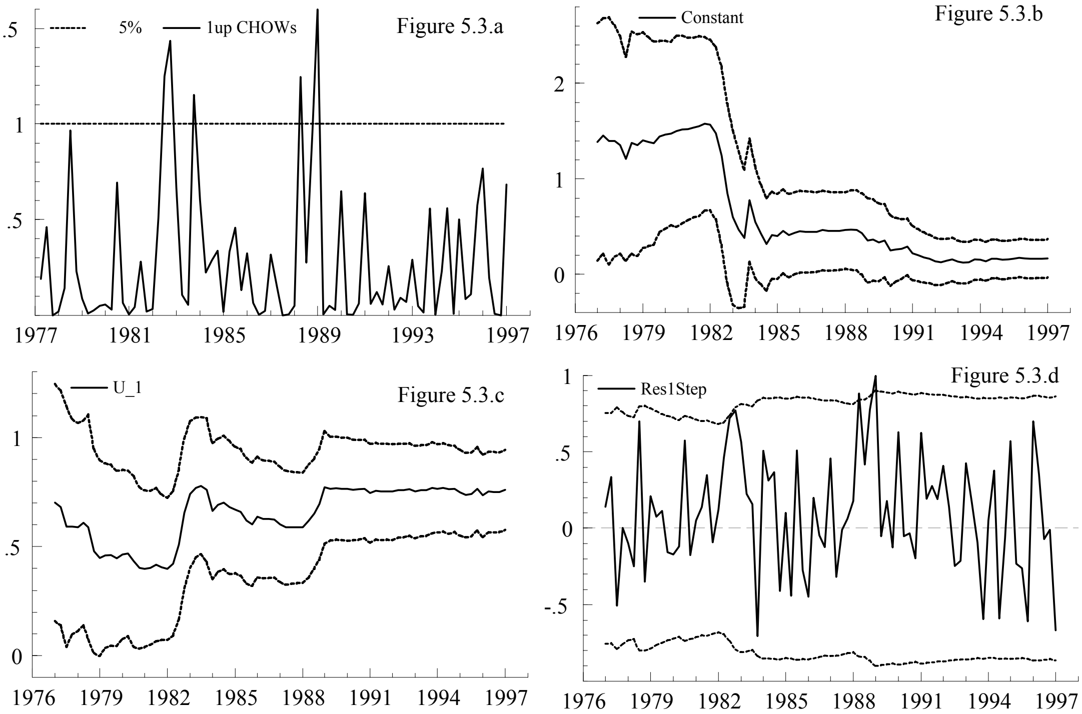


Figure 5.3: *One-step ahead Chow tests with critical values at 5% (5.3.a), recursive OLS estimates of the constant term, the first autoregressive coefficient and the one-step ahead residuals, respectively, each with ±2 SE, see (5.3.b), (5.3.c) and (5.3.d). Initial period 1973:2-1977:1.*

unemployment. The AR(5) model implies an equilibrium level of unemployment, μ, equal to 3.74 per cent, which can be questioned in the light of Figure 5.1. Accordingly, the actual unemployment rate was systematically below the equilibrium level before 1988 and systematically above the equilibrium level afterwards.

5.3. A MS-AR(5) model

This section estimates the MS-AR(5) model, see Section 4.1, and evaluates it against the AR(5) model. It is found that the results from the relatively "well specified" MS-AR(5) model are largely inconclusive and are apparently at odds with the data. Subsection 5.3.1, however, shows that apparently interpretable results can be obtained if the autoregressive terms are omitted from the Markov regime switching model.

The maximum likelihood estimates (MLE) of Equation (4.1) are presented in the second panel of Table 5.1. The estimates are derived under the assumption of two regimes

Table 5.1: Estimated models of Norwegian unemployment

1. AR(5) model: OLS/ML estimates

Log likelihood = 35.8 $\quad \widehat{\alpha} = \underset{(0.099)}{0.17} \quad \widehat{\mu} = 3.74 \quad \widehat{\sigma}^2 = 0.19$

$\widehat{\varrho} = 0.96 \quad \widehat{\phi}_1 = \underset{(0.092)}{0.76} \quad \widehat{\phi}_2 = \underset{(0.104)}{0.19} \quad \widehat{\phi}_3 = \underset{(0.105)}{-0.12} \quad \widehat{\phi}_4 = \underset{(0.103)}{0.61} \quad \widehat{\phi}_5 = \underset{(0.093)}{-0.48}$

Test summary

AR $_{1-5}$ *F(5, 82)* = 2.85[0.02*], *ARCH 4 F(4, 79)* = 0.66[0.62],
Heterosced.: Xi ^{2} : *F(10, 79)* = 0.90[0.54], *XiXj*: *F(20, 69)* = 0.86[0.64]
Normality $\chi^2(2) = 0.57[0.75]$, *RESET: F(1, 89)* = 1.73[0.19]

2. MS-AR(5) model: MLE[1]

Log likelihood = 40.5 $\quad \widehat{\mu}_1 = \underset{(0.844)}{3.24} \quad \widehat{\mu}_2 = \underset{(0.832)}{3.86} \quad \widehat{\sigma}_1^2 = \underset{(0.026)}{0.07} \quad \widehat{\sigma}_2^2 = \underset{(0.025)}{0.09}$

$\widehat{\phi}_1 = \underset{(0.084)}{0.86} \quad \widehat{\phi}_2 = \underset{(0.087)}{0.07} \quad \widehat{\phi}_3 = \underset{(0.087)}{-0.09} \quad \widehat{\phi}_4 = \underset{(0.088)}{0.82} \quad \widehat{\phi}_5 = \underset{(0.086)}{-0.70}$

$\widehat{p}_{11} = \underset{(0.104)}{0.79} \quad \widehat{p}_{22} = \underset{(0.075)}{0.86} \quad \widehat{p}_{12} = 0.21 \quad \widehat{p}_{21} = 0.14 \quad \widehat{\pi}_1 = 0.40 \quad \widehat{\pi}_2 = 0.60$

3. MS-AR(0) model with const. variance: MLE[2]

Log likelihood = $-14.0 \quad \widehat{\mu}_1 = \underset{(0.082)}{2.13} \quad \widehat{\mu}_2 = \underset{(0.116)}{5.26} \quad \widehat{\sigma}^2 = \underset{(0.062)}{0.44}$

$\widehat{p}_{11} = \underset{(0.010)}{0.99} \quad \widehat{p}_{22} = \underset{(0.016)}{0.99} \quad \widehat{p}_{12} = 0.01 \quad \widehat{p}_{21} = 0.01 \quad \widehat{\pi}_1 = 0.59 \quad \widehat{\pi}_2 = 0.41$

4. MS-AR(0): MLE

Log likelihood = $-16.8 \quad \widehat{\mu}_1 = \underset{(0.063)}{1.92} \quad \widehat{\mu}_2 = \underset{(0.163)}{4.81} \quad \widehat{\sigma}_1^2 = \underset{(0.042)}{0.20} \quad \widehat{\sigma}_2^2 = \underset{(0.234)}{1.02}$

$\widehat{p}_{11} = \underset{(0.021)}{0.97} \quad \widehat{p}_{22} = \underset{(0.026)}{0.97} \quad \widehat{p}_{12} = 0.03 \quad \widehat{p}_{21} = 0.03 \quad \widehat{\pi}_1 = 0.55 \quad \widehat{\pi}_2 = 0.45$

The effective sample when 5 lags covers the period 1973:2-1997:1. The std. errors are in parentheses below the estimates. *Denotes significance at the 5% level. [1]Assuming one regime, the OLS estimates have been used as the starting values in the maximum likelihood (ML) estimation. [2]For the sake of convergence the initial values for the means were set equal to 2 and 4 here. The estimates have been obtained using PcGive 9.10, see Doornik and Hendry (1997) and Gauss 3.2. The Gauss software has been kindly provided by J. D. Hamilton.

or equilibria. As in the AR(5) model, increasing the number of lags to five, $q = 5$, lead to a large increase in the log likelihood value, see Table 5.2. The parameter estimates of the AR(5) model in the first panel were used as starting values. However, the results were found to be quite robust to changes in the starting values.

The MS-AR(5) model appears to provide a better fit to the historical unemployment record compared with the AR(5) model. The variances of residuals in the regime switching model, which are almost equal across the two regimes, are about half the value of those in the AR(5) model. The log likelihood value of the MS-AR(5) model is also higher compared with the log likelihood value of the AR(5) model. They are 35.8 and 40.5, respectively. The estimated equilibria are, however, not significantly different from each other at conventional levels of significance, implying weak evidence of two unemployment equilibria. Indeed, we are not able to reject the AR(5) model against the MS-AR(5) model when using the appropriate likelihood ratio (LR) test suggested by Hansen (1992) and (1996. The standardised LR value is 2.19 with a p-value of 0.223.[9] Note that the estimated level of $\widehat{\mu}_1$ at 3.24% is higher and the estimate of $\widehat{\mu}_2$ at 3.86% is lower than what one would expect from a mere glance at Figures 5.1 and 5.2 that presents the actual data.

The transition probabilities indicate that the probability of a switch from the high to the low unemployment equilibrium is smaller than vice versa, $\widehat{p}_{21} = 0.14$ versus $\widehat{p}_{12} = 0.21$. The unconditional probabilities of unemployment evolving around the low and high equilibrium levels are $\widehat{\pi}_1 = 0.4$ and $\widehat{\pi}_2 = 0.6$, respectively. These are long run functions of the transition probabilities, see Hamilton (1994, pp. 681-84). This asymmetry in the transition from one level to another seems to comply with the observed movements of unemployment from one level to another, cf. Figure 5.1. However, the transition probabilities also indicate that the equilibria are quite short dated, which appears to be in conflict with the impression from Figure 5.1. The probability of remaining in the state of high or low unemployment equilibrium in the next period, given that one is in the high or low unemployment equilibrium is 0.86 and 0.79, respectively. Thus the expected duration of the high unemployment equilibrium is only $1/(1 - \widehat{p}_{22}) \approx 7$ quarters, while it is less than five quarters $(1/(1 - \widehat{p}_{11}) \approx 4.8)$ for the lower equilibrium.

Figure 5.4 displays the filtered and smoothed probabilities of being in a state with high unemployment equilibrium over the sample. The graph indicates large discrepancies between the filtered and smoothed probabilities. The smoothed probabilities, based on the full sample, seem to be more decisive regarding the state of unemployment than the filtered probabilities that are derived from observations up to period t, in which the observation

[9] The Gauss software was downloaded from Bruce C. Hansen's home page.

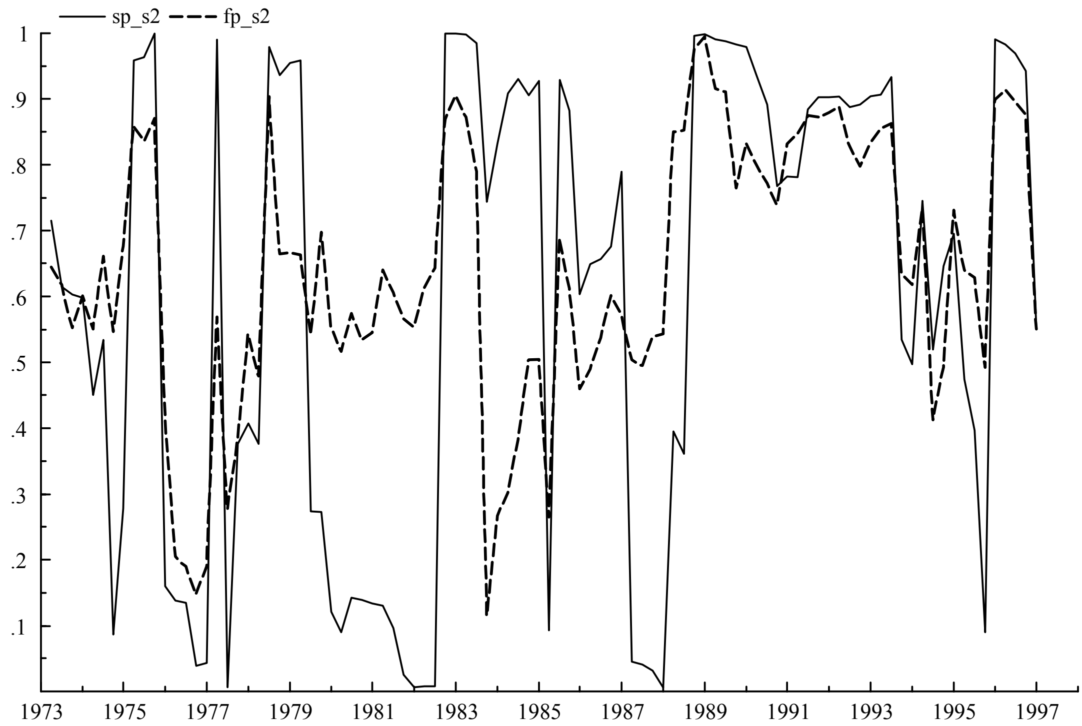


Figure 5.4: *Smoothed and filtered probabilities of being in regime 2, sp_s2 and fp_s2, respectively, 1973:2-1997:1.*

is made. The smoothed probabilities are mostly close to 0 or 1, and in contrast to filtered probabilities, less often close to 0.5.

The smoothed probabilities indicate four brief periods of the high equilibrium state during the 1970s, one major period in the 1980s and another major period extending from the end of the 1980s to the end of the sample in 1997:1. The four brief periods are 1973:2-74:2, 1975:2-75:4, 1977:2 and 1978:3-79:2. The two major periods are 1982:4-87:1 and 1988:4-1997:1.[10] The major periods are, however, interrupted for one quarter and two quarters, in 1985:2 and 1995:3-95:4, respectively. The timing of switches in the mean, μ, during the 1980s is roughly consistent with our earlier observations, see Figures 5.1 and 5.3.a-d. However, the implied changes in parameters during the 1970s, 1985:2, and 1995:3-95:4 are not apparent in these figures, see e.g. one-step ahead Chow tests in Figure 5.3.a.

[10] Observations at time t which are equally likely to have been generated by both regimes are assumed to belong to the regime at time $t-1$.

The MS-AR(5) model seems to attribute quite brief periods, e.g. even single isolated observations, to structural changes in the sense of a change in the parameters. These brief periods are perhaps better characterised as outliers within a given regime rather than representing structural changes. This problem of "spurious switches", low transition probabilities and small deviations between the estimated equilibria may be assigned to the non-rejection of the linearity hypothesis. The problem of "spurious switches" between regimes has also been remarked on by Clements and Krolzig (1998) in their study of US GNP and dealt with by BZ (1998).

Comparison with existing studies

As noted in Section 2, the Markov regime switching model has previously been employed by BZ (1996, 1998) to study unemployment rates in OECD countries. They have reported evidence of shifts between multiple equilibria for a number of countries, including Norway. However, BZ (1996, 1998) estimate the equilibrium rates without autoregressive terms, i.e. under the (a priori) restrictions of $\phi_i = 0 \; \forall \; i$, for all countries. In BZ (1996), the analysis of Norwegian unemployment rate is based on annual data for the period 1960-1993. Assuming constant residual variance across two possible regimes, they estimate the unemployment equilibria to be 2.1% and 5.5%, and find a one-time transition to the high unemployment equilibrium in 1988. On quarterly data for the period 1970:1-1995:4, BZ (1998) allow the residual variance to change across the two regimes. In this case they estimate the equilibrium levels to be 1.83% and 4.72% while the residual variance is estimated to be 0.12% and 1.16% in the low and high unemployment regime, respectively. In this case, their model predicts two periods of high unemployment equilibrium, 1982:3-1984:2 and 1988:1-1995:4.

We are almost able to replicate the results in BZ (1996, 1998) on our quarterly data set by following their course. Panel 3 of Table 5.1 reports estimates quite close to BZ (1996) when the autoregressive terms are excluded from the model (4.1) and a constant variance across regimes is assumed. These restrictions also lead to a one-time transition to the state of high unemployment equilibrium in 1988:4, i.e. in 1988, as in BZ (1996). Panel 4 in Table 5.1 reports the results when the assumption of constant residual variance is relaxed.

Table 5.2: Sensitivity of the results from the Markov switching model to number of autoregressive terms, q.

	q					
	0	1	2	3	4	5
$\widehat{\mu}_1$	1.92	3.66	4.83	4.04	2.82	3.24
$\widehat{\mu}_2$	4.81	4.29	5.40	4.66	3.74	3.86
$\widehat{\phi}_1$		0.97	0.43	0.62	0.50	0.86
$\widehat{\phi}_2$			0.56	0.71	0.26	0.07
$\widehat{\phi}_3$				-0.35	-0.20	-0.09
$\widehat{\phi}_4$					0.41	0.82
$\widehat{\phi}_5$						-0.70
$\widehat{p}_{11}$	0.99	0.34	0.38	0.79	0.60	0.79
$\widehat{p}_{22}$	0.99	0.59	0.84	0.38	0.99	0.86
$\widehat{\sigma}_1^2$	0.20	0.07	0.03	0.16	2.67	0.07
$\widehat{\sigma}_2^2$	1.02	0.15	0.21	0.28	0.19	0.09
likl.	-16.8	19.13	21.68	24.78	23.25	40.50

likl. = Log likelihood value, see Table 5.1 for details.

These are close to the results in BZ (1998) for both the equilibria and the variances. Now the model also predicts two periods of high unemployment equilibrium, 1982:4-1984:3 and 1988:2-1997:1, with one period of low unemployment equilibrium in-between, i.e. from 1984:4 to 1988:1. These dates only differ from those of BZ (1998) by one quarter. The relatively fewer switches, compared with the more general model in the second panel of Table 5.1, are also reflected in the transition probabilities. They suggest that movements between the equilibria are quite unlikely.

The models in panel 3 and 4 of Table 5.1, as BZ (1996, 1998), provide estimates of the equilibria that are close to the values of modes observed in Figure 5.2 and in accordance with the visual impression from Figure 5.1. In addition, the number of switches and their datings appear to be consistent with the unemployment behaviour over time, cf. Figures 5.1 and 5.3.*a-d*. However, these models have lower explanatory powers, i.e. lower log likelihood values, than the AR(5) model and the regime switching models for any number of autoregressive terms, from 1 to 5, cf. panel 1 and 2 in Table 5.1 and Table 5.2. Selecting a model that accounts for the dynamics by autoregressive terms, however, provide results that appear to be at odds with some of the observed features of the unemployment series, see Table 5.2. This table prefers the MS-AR(5) model whose properties are discussed above. In addition, it demonstrates the sensitive of the estimates for the equilibria to the inclusion of autoregressive terms and their number. Similar results are reported in Clements and Krolzig (1998, Table 6.).

One might conclude that the Markov regime switching models provide mixed evidence in favour of multiple equilibria. Models that are better at accounting for the dynamics and have higher explanatory power provide results that are difficult to interpret, both within the unique equilibrium and the multiple equilibria approach. The contrary appears to be true for models that neglect the dynamics. In the latter case, the estimated levels of the equilibria and shifts between them seems to be reconcilable with the data and interpretable within the multiple equilibria approach.
The next section adopts an alternative framework to see if the evidence can be tilted in one or the other direction.

6. Smooth transition autoregressive model

This section analyses Norwegian unemployment in the STAR framework described in Section 4.2. Section 6.1 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 section. Section 6.2 conducts a number of tests to assess the data consistency of the obtained model and Section 6.3 explores its dynamic properties. In particular, it demonstrates that the long run effects of a shock depends on its size, as suggested by (theory) models of multiple equilibria, see Section 3.

6.1. An LSTAR model

While the AR(5) model may not be rejected in favour of the MS-AR(5) model, the following 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. The results in Table 6.1 rejects the AR(5) model at about 5% level of significance when $d = 5$. And for $d = 5$, the lower panel of the table shows rejection of H_{02} at 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 can be rejected in favour of a STAR model for all values of d at 10% level of significance, but still most strongly for $d = 5$ with a p-value of 0.3%. Moreover, H_{02} can also be rejected at the same p-value of 0.3% for $d = 5$. The F-tests may have more power in these parsimonious auxiliary regressions than in their general versions.

Table 6.1: Testing the linear AR(5) against a nonlinear STAR model, and LSTAR against an ESTAR model

d	Testing linearity	p-values
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 nonlinearity	p-values
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]**
Method: OLS, Initial sample: 1973:2-1997:1. * and ** denotes significance at 5% and 1%, respectively.		

Table 6.2 presents an estimated LSTAR model with five autoregressive terms when $d = 5$. The model was estimated by nonlinear least squares (NLS). The results indicate that most of the lagged terms in the nonlinear part as wells of the linear part of the model are insignificant, see the upper panel of Table 6.2. One reason for this might be the high degree of correlation between the regressors, the lagged terms of U_t. Indeed, when the general model is sequentially reduced to a parsimonious model, several of the lagged terms in the nonlinear part of the model become significant, see the parsimonious model in panel 2 of Table 6.2.

The estimate of the transition parameter γ is associated with a high degree of uncertainty, see panel 1 or 2. This finding may be 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

Table 6.2: LSTAR model and its properties

General LSTAR model

$$\widehat{U}_t = \underset{(0.23)}{0.45} + \underset{(0.12)}{0.79} U_{t-1} + \underset{(0.15)}{0.32} U_{t-2} - \underset{(0.16)}{0.14} U_{t-3} + \underset{(0.19)}{0.20} U_{t-4} - \underset{(0.16)}{0.36} U_{t-5}$$

$$+(\underset{(1.24)}{0.03} - \underset{(0.25)}{0.11} U_{t-1} - \underset{(0.25)}{0.32} U_{t-2} + \underset{(0.24)}{0.08} U_{t-3} + \underset{(0.24)}{0.75} U_{t-4} - \underset{(0.32)}{0.31} U_{t-5})\text{x}$$

$$[1 + \exp\{-\underset{(2.54)}{3.70}(U_{t-5} - \underset{(0.59)}{3.63})\}]^{-1}$$

Sample 1973:2-1997:1, Log likelihood value = 43.81 and $\widehat{\sigma} = 0.42$

Parsimonious LSTAR model

$$\widehat{U}_t = \underset{(0.23)}{0.50} + \underset{(0.09)}{0.73} U_{t-1} + \underset{(0.12)}{0.37} U_{t-2} - \underset{(0.13)}{0.31} U_{t-5} +$$

$$(-\underset{(0.16)}{0.41} U_{t-2} + \underset{(0.15)}{0.95} U_{t-4} - \underset{(0.19)}{0.43} U_{t-5})\text{x}[1 + \exp\{-\underset{(2.28)}{3.48}(U_{t-5} - \underset{(0.26)}{3.57})\}]^{-1}$$

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

Dynamic properties

$\widehat{F}(U_{t-5}) = 0 : \widehat{\varrho}_1 = \sum_{i=1}^{5} \widehat{\phi}_i = 0.78,$ $\quad \widehat{\mu}_1 \equiv \frac{0.5}{1-\widehat{\varrho}_1} \approx 2.3$

$\widehat{F}(U_{t-5}) = 1 : \widehat{\varrho}_2 = \sum_{i=1}^{5} (\widehat{\phi}_i + \widehat{\widetilde{\phi}}_i) = 0.90,$ $\quad \widehat{\mu}_2 \equiv \frac{0.5}{1-\widehat{\varrho}_2} \approx 5.1$

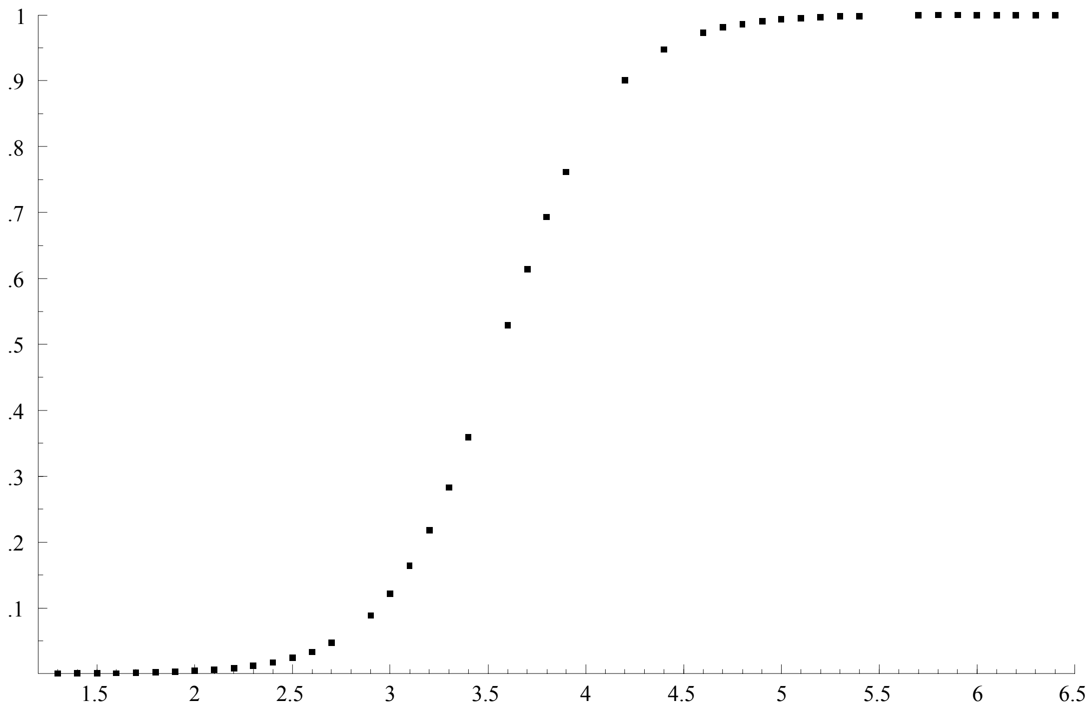


Figure 6.1: *A cross plot of the transition function $\widehat{F}$ (vertical axis) against the transition variable (horizontal axis). One dot represents at least one observation.*

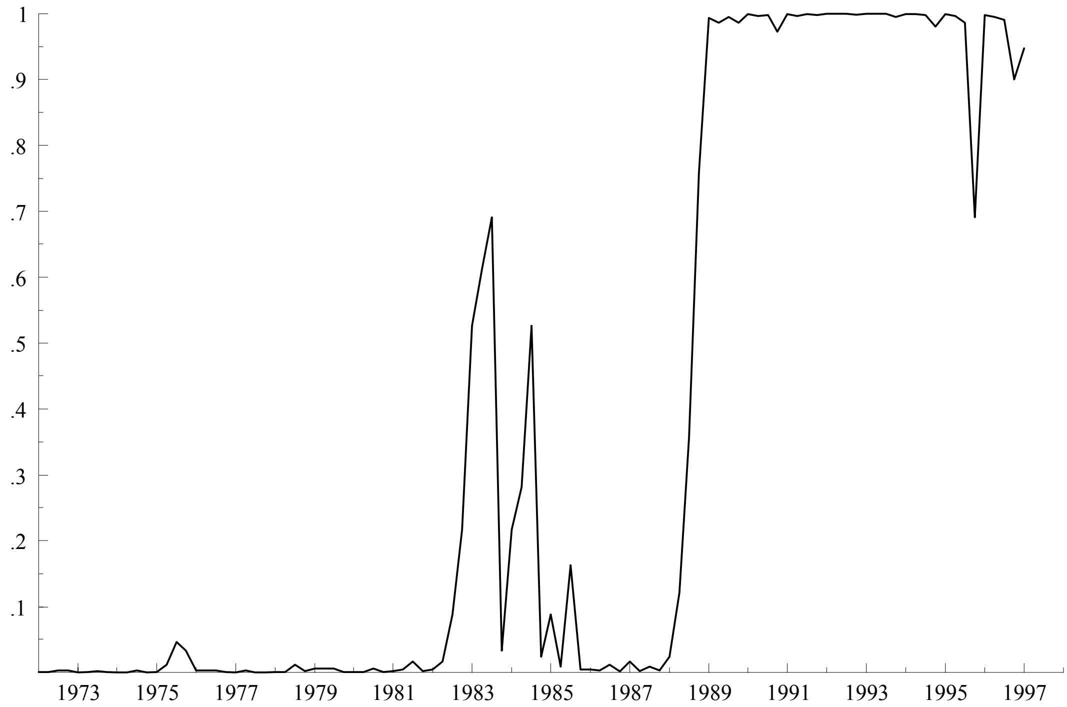


Figure 6.2: *The transition function ($\widehat{F}$) from the parsimonious LSTAR model over the period 1972:1-1997:1.*

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 $\widehat{\gamma}$ is assumed to reflect this feature. In such cases, many observations in the neighbourhood of c are required to obtain precise estimates of γ, cf. Teräsvirta (1994). In the present data set, however, most values of U_{t-d} are clustered around levels of 2% and 5% while c is (quite precisely) estimated at a value of about 3.6%, cf. Figure 5.2.

The large value of $\widehat{\gamma}$ implies that even a 1 percentage point deviation of unemployment from $\widehat{c} \approx 3.6$ is sufficient to bring the estimated transition function $\widehat{F}(U)$ close to 0 or 1, see Figure 6.1.[11] Consequently, the LSTAR model resembles a two regime threshold model because the unemployment rate has mainly been close to either 2% or 5% over the sample period.

Figure 6.2 shows that $\widehat{F}(U)$ moves quite fast from 0 to 1 during 1988 and stays close to

[11] To date the transition of unemployment process from regime to another, the values of the estimated transition function $\widehat{F}(U)$ are obtained by using U_t rather than U_{t-d}.

1 in the subsequent periods of the sample. In the period 1982/83, however, $\widehat{F}(U)$ displays a transitory increase from 0 but falls short of reaching 1. The unemployment process in this period may be interpreted as being in between the two regimes. This is in contrast to the results from the MS-AR(5) and MS-AR(0) models that implies full transition to the high unemployment regime in this period. However, the transition to the high unemployment regime during 1988 is consistent with these models and with Skalin and Teräsvirta (1999) who use a time trend as transition variable. The indicated parameter changes in the early 1980s and in 1988 are consistent with the results from the parameter constancy tests in Figures 5.3.a-d.

Before presenting the dynamic properties of the LSTAR model, it would be appropriate to examine its explanatory power and investigate whether it offers an adequate description of the data or not.

6.2. Model evaluation

The (parsimonious) LSTAR model has slightly higher explanatory power than the AR(5) model and the MS-AR(5). The standard deviation of its residuals is smaller than that of the AR(5) model. The ratio between the standard deviations of residuals is about $0.93 = 0.41/0.44$. The log likelihood value of this model is 42.37 compared with 35.8 and 40.5 for the AR(5) model and the MS-AR(5), respectively. We are, however, not confident about the significance of the differences between these log likelihood values since the appropriate distributions are unknown.

The results in Table 6.3 indicate that the LSTAR model is data consistent. None of the tests regarding the assumptions about residuals is rejected at 5% or 10% levels of significance. The null hypotheses that are tested are: absence of autocorrelation up to 5 lags, no heteroscedasticity, including ARCH type up to order 5, and that the residuals have a normal distribution. Note that the residuals of the linear AR model were autocorrelated even if 8 lags were included, see Table 5.1. The tests for no remaining nonlinearity of STAR type do not indicate any remaining nonlinearity in the model. The test of parameter constancy suggests that the initial non-constancies in the parameters have been modelled satisfactorily, though not completely since the p-value is 10%. The non-rejection of this test implies that there have not been significant changes in the two equilibria over time, see Section 3. Although the alternative hypothesis is that of smooth changes in the parameters, the parameter constancy test has power even if the alternative hypothesis is that of abrupt changes in the parameters, see Eitrheim and Teräsvirta (1996).

Table 6.3: Testing the adequacy of the parsimonious 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]
Parameter constancy: $F(21,66) = 1.54$ [0.10]					
Heteroscedasticity (X_i^2): $F(16, 70) = 0.87$ [0.60]					
Heteroscedasticity (X_iX_j): $F(41, 45) = 0.77$ [0.80]					
Normality, $\chi^2(2)$ =1.15 [0.56]					

The tests for no autocorrelation, no remaining nonlinearity of STAR type and parameter constancy are those suggested by Eitrheim and Teräsvirta (1996). p-values of test statistics in large brackets [.], and std. deviations in small brackets (.). 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 nonlinearity of STAR type were done 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 Doornik and Hendry (1997)).

6.3. Dynamic properties of the model

The LSTAR model suggests that the unemployment is stationary in both regimes, though quite persistent. The equilibria associated with the regimes $\widehat{F}(U) = 0$ and $\widehat{F}(U) = 1$ are at about 2.3% and 5.1%, respectively. The two regimes have different dynamics. The sums of the autoregressive coefficients during the expansion, $\widehat{F}(U) = 0$, and contraction, $\widehat{F}(U) = 1$, are 0.78 and 0.9, respectively. Thus, unemployment becomes more persistent when it rises to higher levels and its response towards a shock become more sluggish. It tends to revert more quickly towards the lower equilibrium upon a (small) deviation from it compared with when it deviates from the higher equilibrium. In other words, the lower equilibrium is relatively more stable than the upper one. This might be a reflection of the actual policy of the Norwegian government that has been aimed at low unemployment during the sample period, cf. Manning (1990). In addition, the wage bargaining institutions are relatively centralised making it difficult to leave the low unemployment regime, cf. Calmfors and Driffill (1988). The relative stability of the lower equilibrium implies that the asymmetries in the adjustment process are likely to increase with the level of unemployment and to be more apparent at the higher equilibrium than at the lower equilibrium.

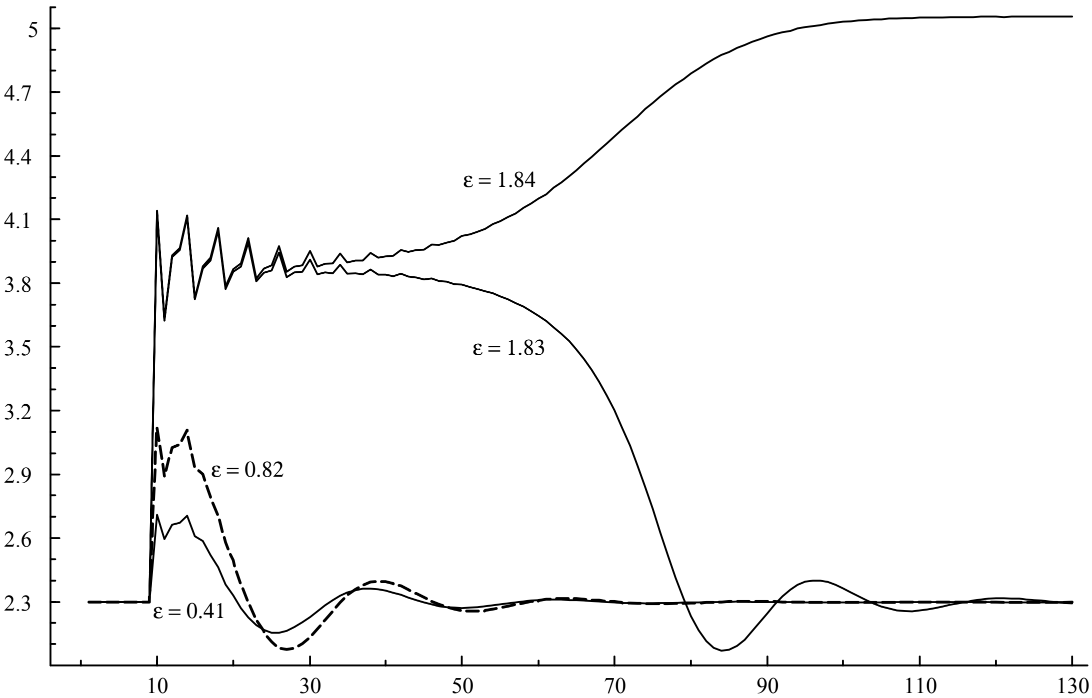


Figure 6.3: *Shocks (ε) of different sizes when in the low unemployment equilibrium of 2.3 %.*

Both equilibria are stable, implying that if unemployment is at one of the equilibria, it will remain there unless disturbed by a shock that causes a transition to the other equilibrium. Figures 6.3 and 6.4 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. Figure 6.3 shows that unemployment returns to the lower equilibrium of 2.3% when exposed to positive single shocks ($\varepsilon > 0$) of sizes up to 1.84 i.e. about 4.5 times the standard error of residuals ($\widehat{\sigma}$). A larger shock, however, makes it converge towards the high unemployment equilibrium at 5.1%. From there, it requires a shock of size -1.28, i.e. of about -3$\widehat{\sigma}$ to revert to the lower equilibrium rate, see

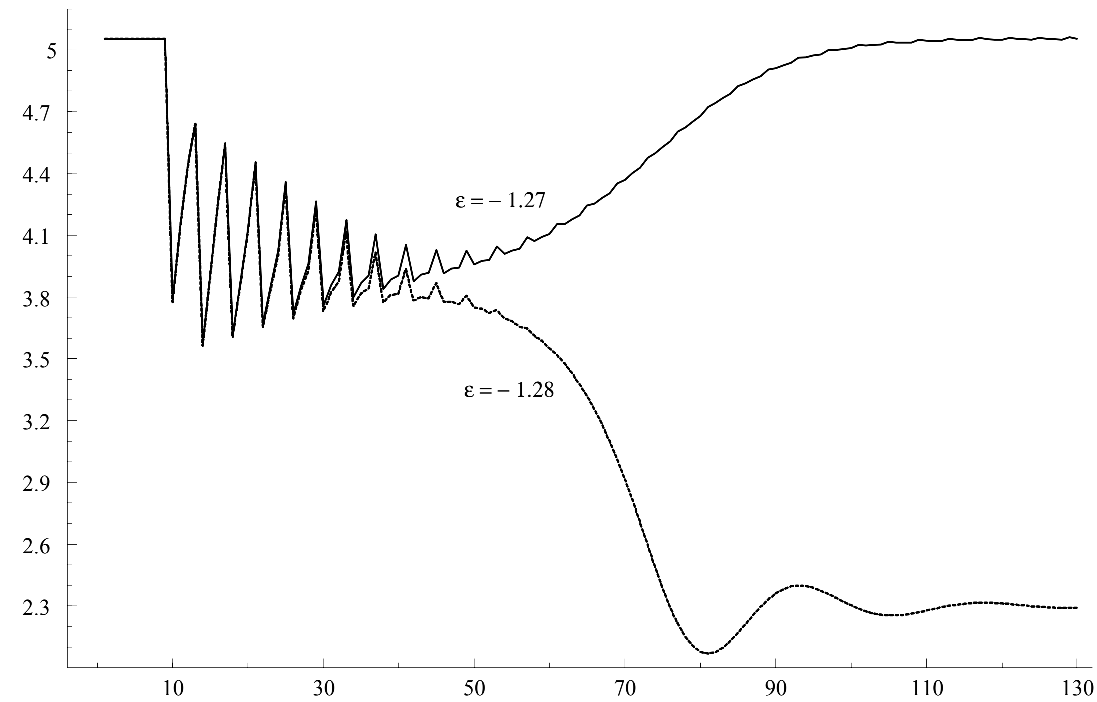


Figure 6.4: *Shocks (ε) of different sizes when in the high unemployment equilibrium of 5.1%.*

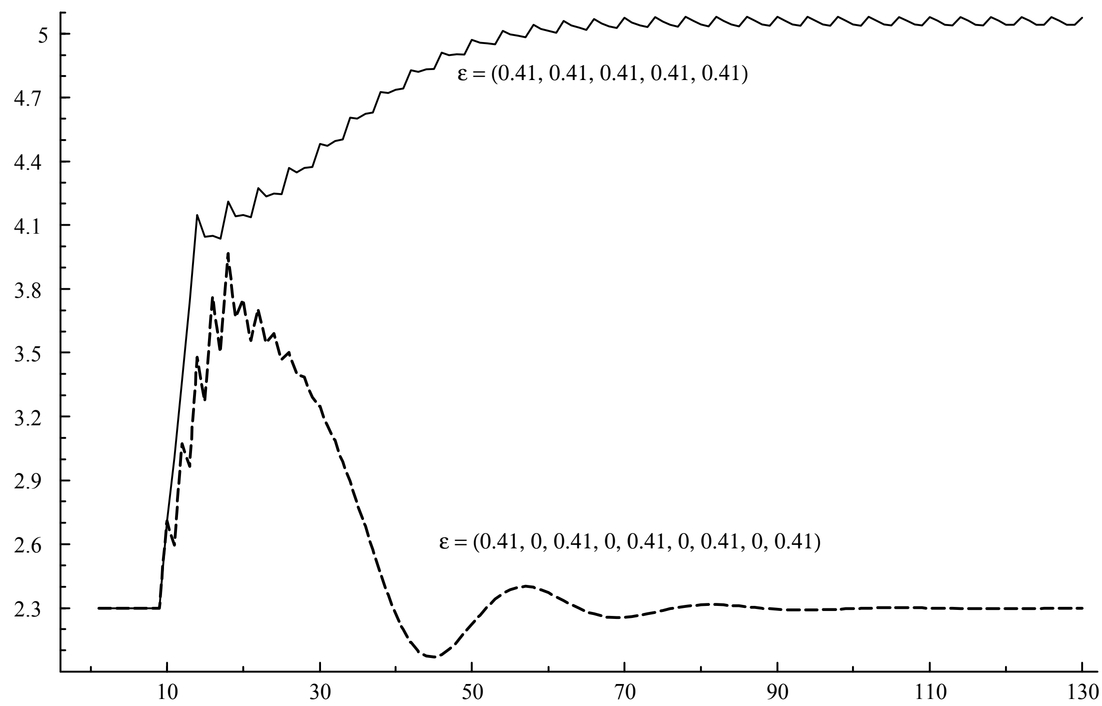


Figure 6.5: *A continuous sequence of five shocks, each equal to $\widehat{\sigma} = 0.41$, and a discontinuous sequence of the same shocks when in the low unemployment equilibrium.*

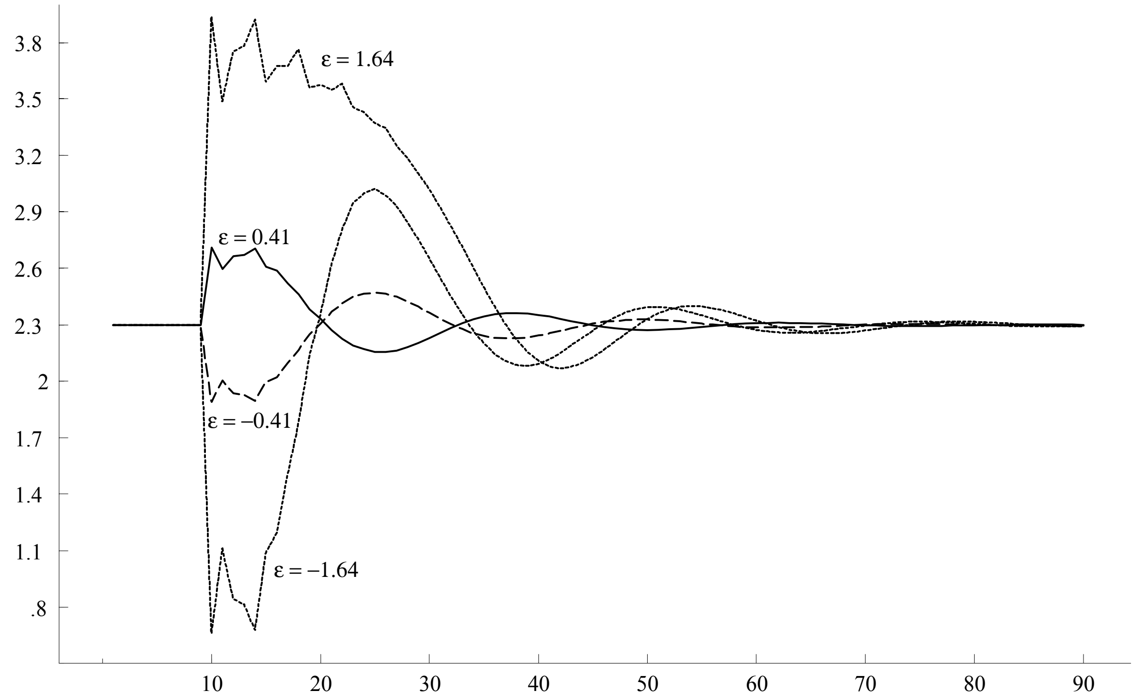


Figure 6.6: *Two positive shocks of sizes 0.41 and 1.64 when in the lower equilibrium and two negative shocks of the same magnitudes from the same initial level.*

Figure 6.4.[12] Otherwise, it will get stuck at the high unemployment equilibrium. Because of the weaker gravitation of the higher equilibrium compared with the lower equilibrium, unemployment requires a weaker impetus to leave the higher equilibrium.

Note that the model does not require a large single shock to cause a 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 ($U_{t-5} - 3.57$) to cause a transition from one regime to another. However, they have to be larger in sum than a large single shock, because of slight reversions towards the initial equilibrium during the intervals between the shocks. Figure 6.5 displays the effects of a continuous sequence of five small shocks of size $\widehat{\sigma} = 0.41$ when unemployment is initially at the lower equilibrium. Unemployment converges to the higher equilibrium in this case too. The same figure lays out the response to a discontinuous sequence of five small shocks of the same size. In

[12] For small shocks unemployment seems to display a limit cycle around the high unemployment equilibrium, see figure 6.5. This is to say that a set of values repeat themselves when approaching the high unemployment equilibrium.

this case unemployment reverts to the initial equilibrium. This failure to leave the initial equilibrium in face of discontinuous small shocks is also related to the local mean reversion property of the process. When the second small shock arrives after the pause of one period, unemployment has already moved slightly back towards the initial equilibrium. Therefore, it requires larger shocks than 0.41 in the periods afterwards to converge towards the higher unemployment regime.

The model implies asymmetric unemployment response to sufficiently large positive and negative shocks. In the following, assume that the shocks fall short of causing a switch between the equilibria. Now, suppose that unemployment is at the lower equilibrium level. Then, for sufficiently large positive shocks the measure of persistence, $\widehat{\varrho}_1$, rises from 0.78 towards 0.91, but remains equal to 0.78 if the shock is negative, irrespective of its size in absolute terms. The opposite happens when unemployment is at the higher equilibrium. In that case the measure of persistence is $\widehat{\varrho}_2$, which remains at 0.91 if unemployment is exposed to a positive shock of any size, but falls from 0.91 towards 0.78 if the shock is large and negative. In both of these cases the recovery of unemployment is faster from the below of a given equilibrium than from the above. In other words, it rises faster than it declines. Figure 6.6 illustrates the case when unemployment is initially at the lower equilibrium and is disturbed by a positive and a negative shock of size 1.64 and -1.64, respectively. Its overshooting from the below stands in contrast with its sluggish recovery from the positive shock.

The importance of the sign of shocks is even more pronounced if they are sufficiently large to cause a switch between the equilibria. In that case, large positive shocks can cause a switch from the lower to the higher equilibrium, while negative shocks only cause a transitory deviation from it. The opposite happens when the unemployment is in the vicinity of the higher equilibrium. However, the effects of small positive or negative shocks are symmetric around a given equilibrium. Figure 6.6 shows this for the case of a positive and a negative shock of size 0.41 and -0.41, respectively.

The figures above demonstrate that the size and the sign of a shock matters for the unemployment dynamics and determines whether it leaves its initial equilibrium or not. Moreover, they elucidate the importance of the initial equilibrium for the response towards a given shock.

7. Conclusions

This essay aimed to derive a data consistent univariate model to test for the possibility of multiple equilibria in Norwegian unemployment and shed light on the issue of whether or not it adjusts asymmetrically when exposed to positive and negative shocks. To this end an LSTAR model has been derived. It has been shown that this model outperforms a linear AR(5) model in explanatory power and appears to be data congruent. The latter property has been established by testing the residual properties and the constancy of the model parameters.

The LSTAR model provides evidence for two stable unemployment equilibria at 2.3% and 5.1%, respectively. The degree of persistence in the vicinity of these equilibria is 0.78 and 0.90, respectively, which implies that the lower equilibrium is relatively more stable than the higher one. The unemployment process tend to exhibit equilibrium reversion around 2.3% before 1988 and around 5.1% afterwards, i.e. until the end of the sample in 1997:1. Although the LSTAR model allows smooth transitions between equilibria, the results indicate a rather abrupt transition from the low to the high unemployment regime in 1988. These results are largely consistent with the conclusions of Bianchi and Zoega (1996) and (1998).

The model implies that large transitory shocks or a sequence of small shocks may cause a transition between equilibria and thus exert permanent effects on the level of unemployment. However, the small shocks must be larger in sum than the single large shock due to the equilibrium reversion property. In addition, due to the relatively lower degree of equilibrium reversion at higher unemployment levels, a smaller shock is required to make a transition from the higher equilibrium to the lower one than vice versa.

The model also implies that unemployment displays asymmetric response to large positive and negative shocks, while the response is symmetric to small positive and negative shocks. In other words, unemployment recovers faster from a fall than a rise, only when the disturbances are large. This result is intuitively appealing since it seems reasonable that small increases or decreases in the labour stock can be made with about the same costs. The findings regarding the asymmetries in the adjustment process are in contrast to Skalin and Teräsvirta (1999) for Norway, but consistent with their results for a number of other OECD countries.

This essay has also examined a linear AR(5), as noted above, and different versions of Markov regime switching models. The AR(5) model was found to be inconsistent with the data and offered an unrealistic estimate of the unique equilibrium at 3.74%. However, the high degree of implied persistence could also be interpreted as evidence of hysteresis in the unit root sense. The results from the Markov regime switching models were found to be highly sensitive to whether one attempted to capture the unemployment dynamics or not. Neglecting the dynamics lead to results that seemed to be consistent with the raw unemployment observations. These favoured two equilibria at around 2% and 5% and two periods of high unemployment equilibrium, as in Bianchi and Zoega (1998). However, attempts to control for dynamics lead to estimates of the equilibria that were close to each other and a quite a large number of switches between equilibria. Although such models offered a better fit, the implied results were difficult to reconcile with the main characteristics of the unemployment data.

3. Employment adjustment in slack and tight labour markets

(*with Ragnar Nymoen*)

Abstract

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

[0] This essay is a revised version of the one presented in the submitted thesis.

1. Introduction

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

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

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

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

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

2. Friction, persistence and multiple equilibria

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

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

$$\ln(N_t) = \Gamma_1 Z_t - f(U^e_{t+1}) + v_t, \quad f' \geq 0, \tag{2.1}$$

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

In order to establish the aggregate consequences of a relationship like (2.1), we express the unemployment rate as

$$U = \ln(N^S) - \omega_1 \ln(N_t) - \omega_2 \ln(N^{rest}) + \varepsilon_t, \;\; \omega_1 + \omega_2 = 1, \tag{2.2}$$

where N^S denotes labour supply and N^{rest} is labour demand in the rest of the economy. ω_1 and ω_2 (and the residual term ε_t) are due to the log linearisation. Assume that **a)** $\ln(N^{rest}_t)$ depends on a set of variables Z^{rest}_t, **b)** U_{t-1} has predictive power for U_{t+1} and that firms use this information, at least. In addition, that **c)** N^S depends linearly on past unemployment due to e.g., "discouraged worker effect", see Pencavel (1986) inter alia, and on a set of explanatory variables Z^S. Then, (2.1) and (2.2) imply:

$$U_t = \delta + \rho U_{t-1} + \omega_1 f(U_{t-1}) + \theta' Z_t + \epsilon_t, \tag{2.3}$$

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

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

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

which implies

$$U_t = \delta + \kappa U_{t-1} + \theta' Z_t + \epsilon_t. \tag{2.4}$$

Since

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

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

$$\mathsf{E}[U_t \mid U_0,\ Z] = \frac{(1-\kappa^t)}{(1-\kappa)}[\theta' Z \delta + \delta] + \kappa^t U_0\ , \tag{2.5}$$

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

$$\mathsf{E}[U_t \mid Z] \approx \frac{\theta' Z}{(1-\kappa)} + \frac{\delta}{(1-\kappa)}, \tag{2.6}$$

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

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

$$f(U_{t-1}) = \frac{1}{1 + e^{-\xi(U_{t-1} - c)}}, \tag{2.7}$$

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

$$\mathsf{E}[U_t \mid Z] \approx \begin{cases} \dfrac{\theta' Z}{(1-\rho)} + \dfrac{\delta}{(1-\rho)} \equiv \mu_1 \\ \dfrac{\theta' Z}{(1-\rho)} + \dfrac{\delta + \omega_1}{(1-\rho)} \equiv \mu_2 \end{cases}, \tag{2.8}$$

where $\mu_1 < \mu_2$. Note that a nonlinear $f(U_{t-1})$ also implies multiple equilibria in sectoral employment. For example, (2.8) and (1) implie

$$\mathsf{E}[\ln(N_t) \mid Z] \approx \begin{cases} \Gamma_1 Z - \mu_1 \\ \Gamma_1 Z - \mu_2 \end{cases}. \tag{2.9}$$

In order to test whether non-linear friction effects can explain the existing evidence of multiple equilibria in the Norwegian labour market, it is necessary to employ multivariate models, see Skalin and Teräsvirta (1999), Bianchi and Zoega (1998) and Akram (1999) for the evidence. A shortcoming of univariate studies that contain evidence of multiple equilibria is their inability to identify the underlying mechanisms at work, e.g., non-linear adjustment costs or labour hoarding, increasing returns to scale, effects on labour supply or perhaps quite simply a mean shift in one or more of the forcing variables in Z.

The next sections explain industry employment and aggregate unemployment in Norway. Specifically, we estimate generalisations of (2.1) together with an equation for the rate of unemployment. The average number of working hours per employed wage earner in industry is also included in the empirical model, since changes in working hours (not only persons) affect total labour input.[1]

3. The econometric framework

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

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} - \alpha (Y - Y^*)_{t-1} + \omega \Delta Z_t + \Omega \varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I). \tag{3.1}$$

Y^* represents the equilibrium level of Y which depends on the level of the Z variables. In our analysis, the Y vector contains the (natural) logs of employment in Norwegian industry (n), of the average working hours of industrial workers (h) and of the economy-wide unemployment rate (u); The Z variables include logs of wage costs, indicators of product demand and capital stock. In Section 5.1, we use cointegration analysis within the context of the corresponding VAR model to estimate the relationships that define Y^*, see Johansen (1988) and (1995b). A deviation between Y and Y^* in a given period is partially adjusted in the subsequent period: $0 < \alpha < 1$. $\sum_{i=1}^{k} \Gamma_i$ also conveys information about the dynamic behaviour of Y. ΔZ_t represents short run effects of the Z variables.

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

The disturbance term is a vector $\Omega\varepsilon_t$ with zero mean and covariance matrix $\Omega'\Omega$, as ε_t is by assumption an identically, independently distributed vector with standard normal distribution.

The constant parameter VEqCM encompasses the theoretical model in Section 2 for the case of linear cyclical adjustment costs, i.e. linear $f(u_{t-1})$. For example, if the long run employment equation contains the rate of unemployment u, persistence in the unemployment rate can be (partly) ascribed to linear adjustment costs in employment.

A generalisation of (3.1) that allows for shifts in e.g., the dynamics of Y and the short run effects of forcing variables is given by

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha(s_t)(Y - Y^*)_{t-1} + \omega(s_t)\Delta Z_t + \Omega(s_t)\varepsilon_t, \qquad \varepsilon_t \sim IIDN(\mathbf{0},\ I), \tag{3.2}$$

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.[2] We assume that s_t is an unobservable state variable that takes on discrete values in the space $\{1,\ 2,...,\ S\}$ governed by a first-order Markov chain, see e.g., Hamilton (1989) and Krolzig (1997). Since s is unobservable, probabilistic inference about the value of s_t is based on the information available at time τ and the estimated values of all parameters in the system for all states, say $\widehat{\Theta}$. 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:

$$P(s_t = j \mid Y_t,\ Z_t;\ \widehat{\Theta}), \qquad j = 1,\ 2, ...,\ S \text{ and } t = 1, 2, 3..., T. \tag{3.3}$$

A potential shortcoming of model (3.2) is that it imposes symmetric effects on Y of positive and negative changes in its determinants, in a given state. It is not unlikely

[2]The case of constant parameters, model (3.1), corresponds to $s_t = 1,\ \forall\ t$.

that employment responds more slowly to positive impulses than to negative ones, if e.g., cycle independent hiring costs are larger than firing costs. The empirical relevance of this shortcoming can be assessed by considering a slightly generalised version of the model with state dependent effects. For example, one may use the following model, which allows for different responses to overmanning $(Y - Y^*)^+$ and undermanning $(Y - Y^*)^-$ and to positive and negative changes in the exogenous variables, ΔZ^+ and ΔZ^-, respectively, in state s. Here, superscript "+" denotes that a variable $X^+ = X$ iff $X \geq 0$ while $X^+ = 0$ iff $X < 0$; similarly, $X^- = X$ iff $X \leq 0$ while $X^- = 0$ iff $X > 0$.

$$\begin{aligned} \Delta Y_t &= \sum_{i=1}^{p} \Gamma_i(s_t)\Delta Y_{t-i} - \alpha^+(s_t)(Y - Y^*)^+_{t-1} - \alpha^-(s_t)(Y - Y^*)^-_{t-1} + \\ &\quad \omega^+(s_t)\Delta Z^+_t + \omega^-(s_t)\Delta Z^-_t + \Omega(s_t)\varepsilon_t. \end{aligned} \tag{3.4}$$

Given the large number of parameters to estimate, (3.4) requires a relatively large number of observations to provide precise coefficient estimates and conclusive results.

4. Data

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

The elements of the Y vector, in levels, are displayed in Figure 4.1. The number of persons employed in the manufacturing and construction sector displays 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, compared with 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 1988/89, and it peaks

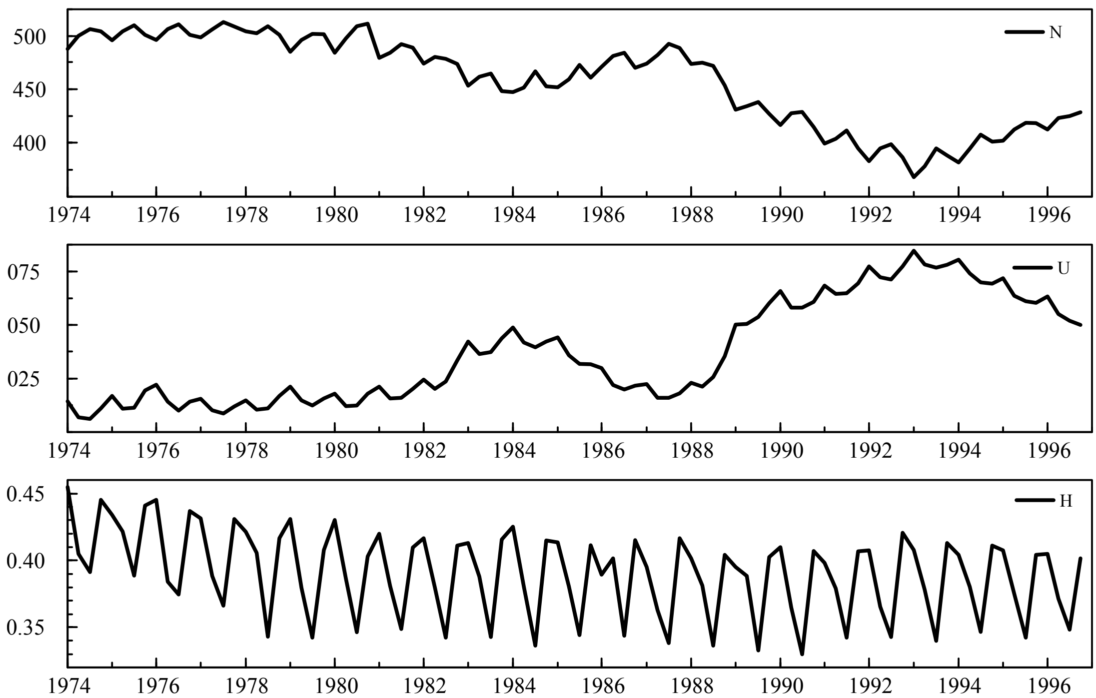


Figure 4.1: *Time series of the Y variables (in levels) over the sample period: 1974(1)–1996(4). Persons employed in manufacturing and construction in thousands (N), the aggregate unemployment rate (U) and average working hours in manufacturing and construction (H) in thousands.*

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 1988/89 that led to a shift in its long run mean, see e.g., Bianchi and Zoega (1998) and Skalin and Teräsvirta (1999). Similarly, the downward shift in industry employment in the late 1980s can be interpreted as a shift in the long run mean of employment.

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

Augmented Dickey Fuller (ADF) tests presented in the appendix suggest that logs of N, U and H (denoted by small letters) may be considered as 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,

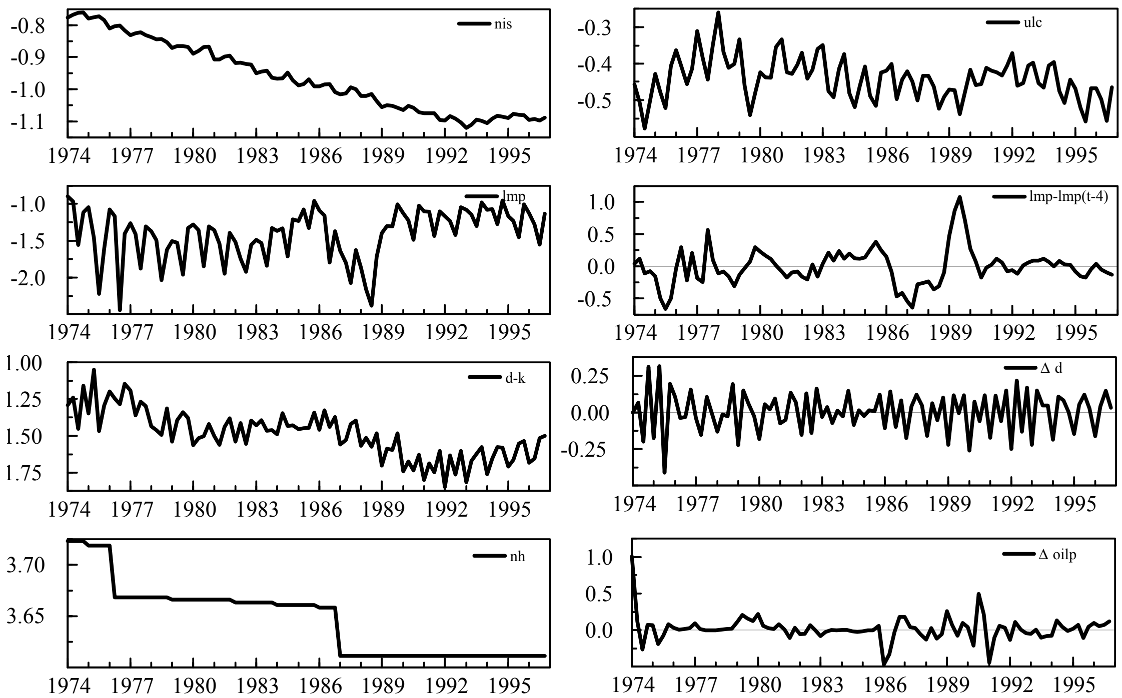


Figure 4.2: *Time series of the Z variables and their transformations over the period 1974(1)-1996(4). From left (in logs): Share of industry employment in total employment* (*nis*), *unit labour costs (ulc), the programme ratio (lmp*) and the annual growth in the *programme ratio* ($\Delta_4 lmp$), *indicator of capacity utilisation* ($d - k$), *quarterly growth in aggregate demand* (Δd), *normal working hours (nh) and finally, quarterly growth in crude oil prices* ($\Delta oilp$).

it contains unit labour costs (ulc), normal (institutional) working hours per week (nh), demand relative to capital stock ($d - k$), the labour market program ratio (lmp), crude oil prices ($oilp$) and finally the share of industry employment in total employment (nis). Figure 4.2 shows a downward trend in nis over most of the sample period. This trend is negatively correlated with e.g., the secular rise in the female labour participation rate and in part-time work; with technological changes; with the tendency towards decentralisation of the wage bargaining process; and with increases in social welfare programs. These structural developments may have contributed to a rise in the unemployment rate over time, see e.g., Dornbusch and Fischer (1994, pp. 511) and Layard et al. (1991).

Most elements of the Z and ΔZ vectors are displayed in Figure 4.2. 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.

Table 5.1: Diagnostics for 5th order conditional VAR for industry employment, working hours and aggregate unemployment rate (in logs); 1974(1)-1996(4); p-values in square brackets.

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

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

5. A linear model

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

Table 5.1 reports the outcome of tests for residual misspecification. The results suggest that the empirical system is adequately specified.

5.1. Cointegration

We next tested for cointegration using the Johansen (1988) procedure, within a system that restricted ulc_{t-1}, nh, nis_{t-1}, $(d-k)_{t-1}$ and a deterministic trend to the cointegration space, while the constant term, $\Delta_4 lmp_{t-1}$, Δd_t, $\Delta_4 d$ and the dummy variables were entered unrestricted, cf. Harbo et al. (1998) and Doornik et al. (1998). The results are reported in Table 5.2. It contains the relevant eigenvalues and the associated trace (*Tr*) statistics employed in testing the hypothesis of $(r-1)$ versus r cointegration vectors. The critical values are from Table 2 in Harbo et al. (1998).

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

Table 5.2: Cointegration rank.

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

Table 5.3: Restricted cointegration analysis, identification of 3 cointegration vectors

$\widehat{\beta}'$	n	u	h	ulc_1	nh	$(d-k)\_1$	nis_1
1	−1	−0.14	−1	−0.13	0	0.20	0
		(0.01)		(0.05)		(0.03)	
2	−1.81	−1	0	0	0	0	−3.97
	(0.78)						(0.64)
3	0	0	−1	0	1	0	0

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

Table 5.3 therefore imposes relevant restrictions on the β and α vectors, which are jointly acceptable with $\chi^2(11) = 16.86$ [0.11].[3] Figure 5.1 shows the recursive estimates of the β-coefficients and their 95% confidence intervals denoted as $\pm 2SE$. The estimates of the unrestricted β-coefficients appear statistically significant and stable over the period 1985(1)–1996(4).

The restricted cointegration vectors are interpretable. A rise in u reduces the equilibrium level of employment which may suggest a reduction in labour hoarding in the face

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

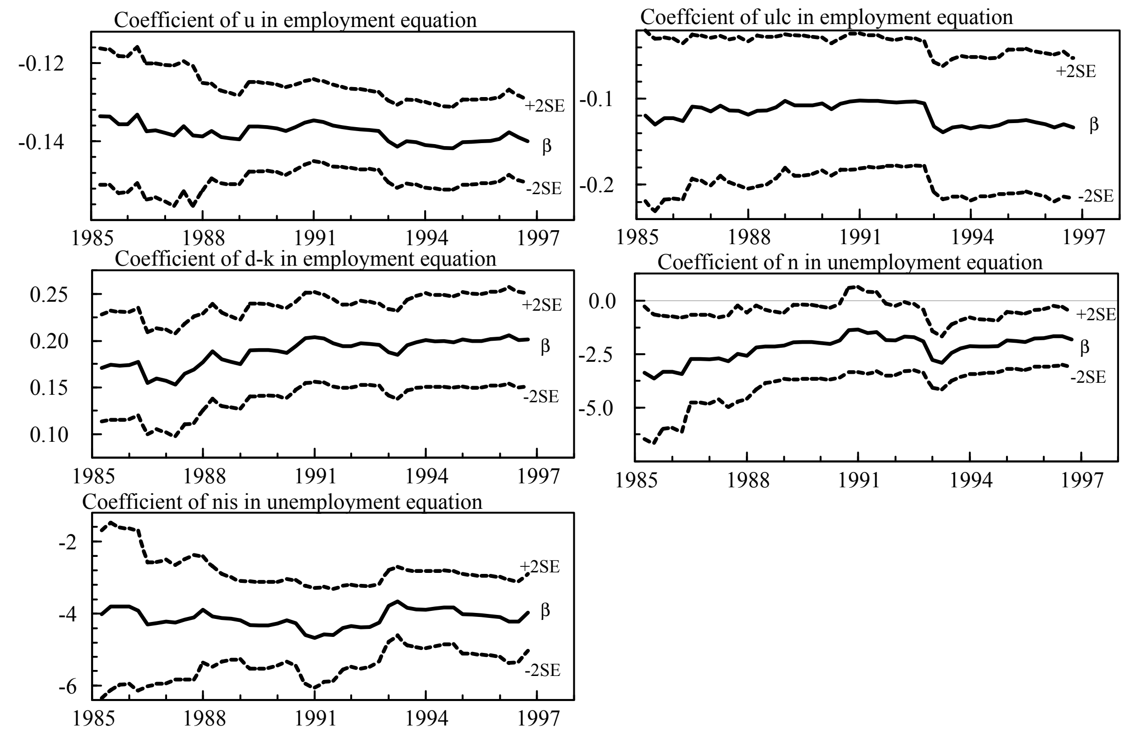


Figure 5.1: *Recursive estimates of the cointegration vectors with +/-2SE. Initial sample: 1974(1)–1984(4).*

of easier access to labour, see Section 2. h and n appear to be perfect substitutes in the long run, which is consistent with the "labour sharing view". A rise in ulc reduces the employment, consistent with a downward sloping demand curve for labour. The positive coefficient estimate of $(d-k)$ suggests that higher capacity utilisation raises employment, or alternatively, a rise in the capital stock (k) substitutes employment. The second vector implies a reduction in u following a rise in sectoral employment n, though the coefficient estimates are imprecise. Furthermore, the proxy for structural changes, nis, is significant in the long run unemployment equation. The third vector suggests that average working hours follow the institutionally determined working hours, nh.

The restricted $\widehat{\alpha}$ matrix in Table 5.3 shows that both n and u respond to deviations between the actual and the equilibrium values of n and u.[4] The test of joint restrictions on the α and β' matrices accepts the weak exogeneity of h for the long run parameters in

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

the employment and unemployment equations. This seems inconsistent with the common finding that working hours act as a buffer against deviations between actual and equilibrium level of employment, cf. Jacobson and Ohlsson (2000) *inter alia*. However, the more restricted simultaneous equation model in the next subsection does not support the weak exogeneity of hours. This apparently contradictory result may be ascribed to low test power in the unrestricted VAR.

5.2. A simultaneous equation model with linear friction effects

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

$$\begin{aligned} n - n^* &= n - \{0.20(d-k) - h - 0.13(u + ulc)\}, \qquad (5.1)\\ u - u^* &= u - \{-1.81n - 3.96nis\}, \qquad (5.2)\\ h - h^* &= h - nh. \qquad (5.3) \end{aligned}$$

Using these equilibrium correction terms, the conditional VAR model was reformulated as a (conditional) VEqCM of order 4 in the differences. Thereafter, parsimony was sought through data consistent coefficient restrictions. Further, the parsimonious version of the model was reformulated as a structural VEqCM with contemporaneous effects between the endogenous variables, cf. Bårdsen and Fisher (1999) and Boswijk (1995). Accordingly, $(n-n^*)_{t-1}$ was restricted to the equation of Δn_t while $(u-u^*)_{t-1}$ was restricted to the equation of Δu_t. Table 6.2 presents the preferred specification of the structural VEqCM which has been estimated by FIML. The diagnostics indicate that the standard assumptions regarding the residuals are not violated at the standard levels of significance. The test for overidentifying restrictions shows that it parsimoniously encompasses the initial VEqCM.

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

Table 5.4: Simultaneous equation model with linear friction effects

Industry employment

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

Aggregate unemployment

$$\widehat{\Delta u_t} = \underset{(0.130)}{0.735} - \underset{(0.386)}{0.523}\,\Delta_3 n_t + \underset{(0.063)}{0.343}\,\Delta u_{t-1} - \underset{(0.069)}{0.150}\,\Delta u_{t-2}$$
$$+ \underset{(0.050)}{0.171}\,\Delta u_{t-3} + \underset{(0.072)}{0.467}\,\Delta u_{t-4} - \underset{(0.034)}{0.192}\,(u-u^*)_{t-1}$$
$$- \underset{(0.044)}{0.156}\,\Delta op_{t-1} - \underset{(0.075)}{0.240}\,\Delta_4 d_t + \underset{(0.539)}{2.04}\,\Delta_4 nh_t$$
$$- \underset{(0.027)}{0.103}\,\Delta_4 lmp_{t-1} - \underset{(0.026)}{0.150}\,CS_{t-1}$$
$$\hat{\sigma}_u = 6.660\%$$

Industry hours

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

Diagnostics

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

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

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

The model in Table 6.2 is considered as an empirical counterpart to the theoretical model in Section 2, with linear specification of friction effects $f(U_{t-1})$. Section 2 shows that a non-linear $f(U_{t-1})$ may imply multiple equilibria; more specifically, shifts in the long run means of $u - u^*$ and $n - n^*$, as u^* and n^* are interpreted as counterparts to $\dfrac{\theta' Z}{(1-\rho)}$ and $\Gamma_1 Z + \dfrac{\theta' Z}{(1-\rho)}$ in Section 2. To investigate this possibility we defined $f(U_{t-1})$ as a logistic function of U_{t-1}, as in equation (2.7). The value of the threshold parameter (c) was set to 0.04 and that of the steepness parameter (ξ) to 100; since estimates of c and ξ were found to be quite imprecise when the method of Maximum Likelihood was applied to the employment equation in Table 6.2, cf. Teräsvirta (1998). Consequently, $f(U_{t-1})$ behaves as a step function with a value close to 1 (high friction) when $U_{t-1} < 0.04$ and close to 0 (low friction) when $U_{t-1} > 0.04$.

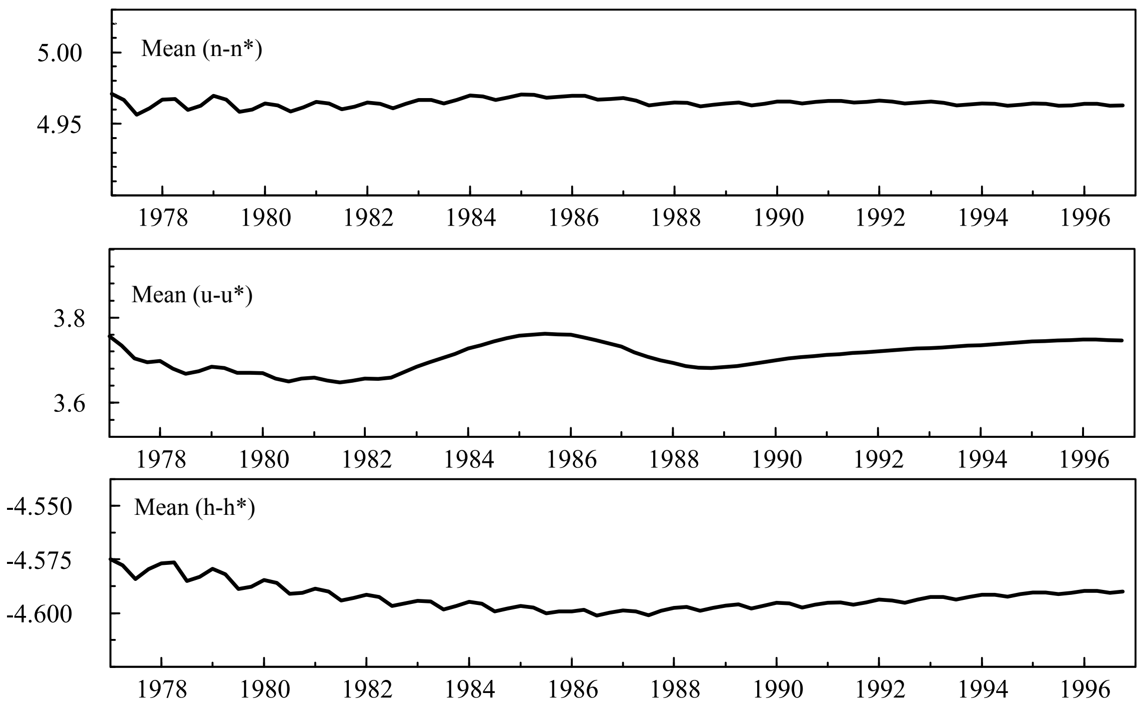


Figure 5.2: *Recursive estimates of the means of* $n - n^*$, $u - u^*$ *and of* $h - h^*$ *over the period 1977(1)–1996(4). The initial estimates are based on observations from the period 1974(1)–1976(4).*

Notably, the joint test of the significance of the logistic $f(U_{t-1})$ when added to the employment and unemployment equations in Table 6.2 yielded $\chi^2(2) = 0.023[0.989]$, 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 5.2 suggests that possible changes in the marginal means of u and n should be attributed to the non-modelled variables and not to labour market friction effects. The figure displays recursive estimates of the means of $n - n^*$, $u - u^*$ and $h - h^*$ over the period 1977(1)–1996(4). The stability of the parameter estimates defining n^*, u^* and h^* is shown above, in Figure 5.1.

Apparently, tests of the overall stability of the structural VEqCM in Figure 5.3 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 exceed the critical value of 1 over the period 1985(1)–1996(4).

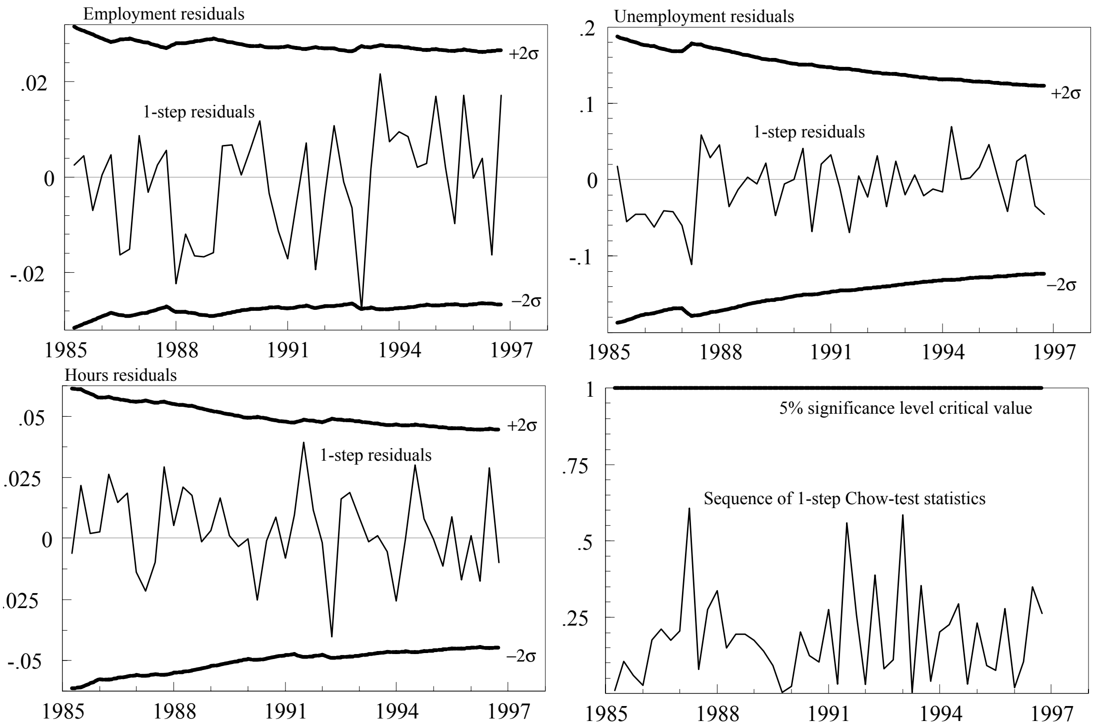


Figure 5.3: *1-step ahead residuals ±2 estimated standard errors based on the equations of employment, unemployment and hours. Also, a sequence of 1-step Chow tests scaled by their critical values at the 5% level of significance.*

However, these tests may understate possible non-constancy in the short run parameters of the VEqCM because the long run parameters appear remarkably constant over the sample in Figures 5.1 and 5.2. Hendry (2000) shows that even large shifts in short run parameters, representing e.g., dynamics, adjustment speeds and intercepts, are difficult to detect if parameters defining the long run equilibrium remain unaltered. Note that the full sample estimates of the long run means of $n - n^*$ and $u - u^*$ in Figure 5.2 are close to the derived long run estimates of the composite constant terms in Table 6.2, $0.543/0.11 \approx 5$ and $0.735/0.192 \approx 4$. This suggests that the composite constant terms in the employment and unemployment equations mainly consist of the evidently stable equilibrium means of $n-n^*$ and $u-u^*$, times the associated equilibrium correction coefficients; Implicitly, other components of the composite constant terms, including the autonomous growth rates in employment and unemployment, seem to be numerically small or to outweigh each other. In the equation for hours, the equilibrium mean of $n - n^*$ seems to be cancelled by the equilibrium mean of $h - h^*$, which may explain the insignificance and hence the exclusion of a constant term in the hours equation, see Figure 5.2.

Section 6 investigates whether the short run parameters of the VEqCM, characterising persistence in employment and unemployment and their response to changes in exogenous variables, depend on the cyclical phase of the economy. In line with common practice, we assume that a model of hours (h) with state dependent parameters is not called for. Commonly, adjustment in working hours is modelled independently of the phase of the economy since costs in adjusting hours are small relative to the costs associated with adjusting persons, see e.g., Hamermesh and Pfann (1996) and Bosworth et al. (1996). The time series of H in Figure 4.1 lends support to this practice.

6. State dependent adjustment

The employment and unemployment equations in Table 6.2 were estimated separately assuming two states, i.e. $S = 2$.[5] The estimation was conducted by Maximum Likelihood (ML) using a version of the Expectation Maximisation (EM) algorithm proposed by

[5]Results based on $S = 3$ turned out to be difficult to interpret. Also, estimation of both equations when all (short run) parameters in both equations were subjected to common shifts (i.e. imposing a common cycle) did not seem feasible; In particular, estimation of the reduced form of these equations subject to common shift led to failure of convergence for both $S = 3$ and $S = 2$.

Hamilton (1990), see Krolzig (1997). The parameter estimates and the series of filtered and smoothed probabilities are obtained jointly by iterations between (preliminary) estimates of the parameters and those of the probabilities. The ML estimators are consistent and asymptotically normal under quite general regularity conditions, see e.g., Hamilton (1993) and (1996) Krolzig (1997).

The outcomes for the employment and the unemployment equations are presented in Table 6.1 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 4.1 and the filtered and smoothed probabilities of $s_t = 2$ for the employment and unemployment in Figure 6.1.

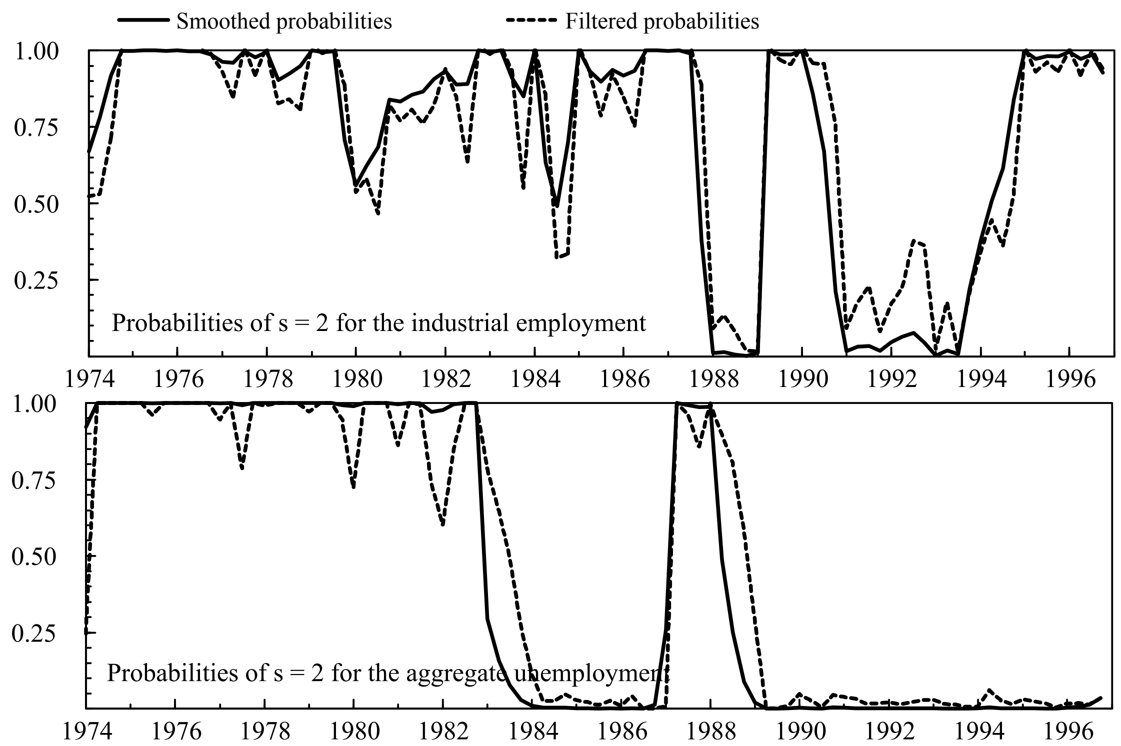


Figure 6.1: *The filtered and smoothed probabilities of industrial employment and aggregate unemployment being in state 2: the expansion phase.*

Figure 6.1 suggests some differences in the cycles of the industry employment and the aggregate unemployment rate. Notably, the dates of switches between the contraction and expansion phases are different from about 1984. In particular, the probabilities related to the unemployment series suggest a recession even after 1993, in contrast to the probabilities

Table 6.1: Models with state dependent parameters.

Industry employment

In recession:

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

In expansion:

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

Aggregate unemployment

In recession:

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

In expansion:

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

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

related to employment. This is not surprising given that the unemployment rate was still more than twice its size in the 1970s and the early 1980s. Also, the filtered and smoothed probabilities based on the unemployment behaviour offer a clearer classification into the two regimes than the corresponding probabilities for the employment behaviour.

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

Table 6.1 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 characterised by labour shortage. In recession, the autoregressive coefficient is insignificantly different from zero and the absolute value of the estimated equilibrium correction coefficient is more than three times its size than in the expansion phase of the economy, 0.285 versus 0.083. Furthermore, the coefficient estimates of all the other regressors (except the impulse dummies) tend to double, at least, when there is a switch from expansion to recession. The corresponding coefficient estimates in Table 6.2 are largely between the state dependent coefficient estimates. This implies that a linear (constant parameter) characterisation of the employment behaviour may underestimate the employment response to shocks in recessions and overestimate the response in expansions.

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

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

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

7. Asymmetric response to shocks?

The employment response may depend on the sign of a shock since hiring costs are believed to be higher than firing costs, see e.g., Hamermesh and Pfann (1996). Table 7.1 presents a generalised version of the employment equation in Table 6.1 where the employment is allowed to respond asymmetrically to positive and negative shocks, as in equation (3.4). Specifically, in each of the two states, the response is allowed to vary with positive and negative deviations from the equilibrium employment and to positive and negative changes in the other regressors, except the autoregressive and deterministic terms.

The increased flexibility of this model has led to a large reduction in the standard errors of the residuals in both states. However, the coefficient estimates are less precise than in the previous models. Also, the coefficient estimate of Δn_{t-4} has become larger in recession than in expansion, relative to the estimates in Table 6.1; The opposite has happened in the case of CS_{t-1}. These changes possibly call for a more adequate representation of seasonal

Table 7.1: Model with sign and state dependent parameters.

Industry employment

In recession:

$$\widehat{\Delta n_t} = \underset{(0.139)}{0.874} + \underset{(0.080)}{0.416}\,\Delta n_{t-4} - \underset{(0.010)}{0.067}\,\Delta u_t^+ + \underset{(0.009)}{0.002}\,\Delta u_t^-$$
$$+ \underset{(0.010)}{0.026}\,\Delta u_{t-2}^+ - \underset{(0.016)}{0.108}\,\Delta u_{t-2}^- - \underset{(0.027)}{0.196}\,\Delta h_t^+ - \underset{(0.035)}{0.088}\,\Delta h_t^-$$
$$- \underset{(0.028)}{0.176}\,(n-n^*)_{t-1}^+ - \underset{(0.028)}{0.179}\,(n-n^*)_{t-1}^- + \underset{(0.019)}{0.055}\,\Delta d_t^+ + \underset{(0.020)}{0.056}\,\Delta d_t^-$$
$$- \underset{(0.005)}{0.0050}\,i81q1_t + \underset{(0.007)}{0.022}\,i86q1_t + \underset{(0.005)}{0.003}\,CS_{t-1}$$
$$\hat{\sigma}_{n,\,1} = 0.380\%$$

In expansion:

$$\widehat{\Delta n_t} = \underset{(0.236)}{0.331} + \underset{(0.075)}{0.193}\,\Delta n_{t-4} - \underset{(0.013)}{0.010}\,\Delta u_t^+ + \underset{(0.015)}{0.018}\,\Delta u_t^-$$
$$+ \underset{(0.012)}{0.004}\,\Delta u_{t-2}^+ - \underset{(0.012)}{0.002}\,\Delta u_{t-2}^- - \underset{(0.043)}{0.199}\,\Delta h_t^+ - \underset{(0.034)}{0.066}\,\Delta h_t^-$$
$$- \underset{(0.047)}{0.065}\,(n-n^*)_{t-1}^+ - \underset{(0.048)}{0.062}\,(n-n^*)_{t-1}^- + \underset{(0.022)}{0.000}\,\Delta d_t^+ + \underset{(0.018)}{0.062}\,\Delta d_t^-$$
$$\underset{(0.012)}{0.042}\,i81q1_t + \underset{(0.009)}{0.025}\,i86q1_t + \underset{(0.006)}{0.020}\,CS_{t-1}$$
$$\hat{\sigma}_{n,\,2} = 0.687\%$$

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

effects in the model.

Table 7.1 offers mixed evidence of an asymmetric response to positive and negative changes in the explanatory variables. In particular, the response to over- and undermanning, $(n-n^*)_{t-1}^+$ and $(n-n^*)_{t-1}^-$, is symmetric across the two states.[6] The exceptions are the response to changes in working hours (Δh) and in aggregate demand (Δd) which appear asymmetric. The coefficient estimates of Δh_t^+ are more than twice the size of the coefficient estimates of Δh_t^- in both states, suggesting that a reduction in employment can be achieved faster than an expansion. As regards the demand shocks, the coefficient estimate of Δd_t^+ is zero while that of Δd_t^- is 0.062 in an expansion. This finding also suggests that a reduction is easier than an expansion. However, in a recession, this asymmetry seems to disappear as the coefficient estimate of Δd_t^+ and of Δd_t^- are almost identical.

However, Table 7.1 substantiates the evidence in favour of state dependent employ-

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

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

The table also supports the relevance of the dynamic friction effects, at least if we look at the case of a recession. (In the state of expansion, the estimates of the unemployment terms become small relative to those in the state of recession and statistically insignificant at the 5% level). Also, in this highly non-linear model, the implied equilibrium solution of employment is the same in both states and equal to that implied by the models in Table 6.2 and 6.1.

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

8. Conclusions

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

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

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

Appendix: Data definitions and properties

The data set has been extracted from the database of RIMINI: the quarterly macroeconometric model used in Norges Bank (The Central Bank of Norway). Square brackets include the variable name in the RIMINI data base. The data set is available on request.

- CS: Centred seasonal for the first quarter in a year.
- D: Indicator of aggregate demand. [DEMIBA.2].
- H: Average working hours per employed wage earner in manufacturing and construction. Thousand hours. [FHIBA].
- i19yy:q1: Impulse dummy, 1 in 19yy:1 and zero elsewhere.
- K: Stock of physical capital in manufacturing and construction. Mill. 1993 NOK. [KIBA].
- LMP: Number of unemployed on labour market programs divided by total unemployment. [AMUN].
- NH: Normal weekly working hours in Norway. Hours. [NH]
- N: Employment in manufacturing and construction. 1000 persons. [NWIBA].
- NIS: Total employment in manufacturing and construction relative to total employment in mainland Norway. Rate. [NWIBA/NWF].
- OILP: Spot price of Brent Blend crude oil in US $, indexed. [OLJEPIND].
- U: Total unemployment rate as a fraction of total labour force. [UTOT2].
- ULC: Unit labour costs (inclusive pay roll tax) in manufacturing and construction deflated by the producer price index. 1993 NOK. [WCIBA/PYIBA.ZYIBA].

Table 8.1: ADF tests of unit roots; 1974(1)-1996(4)

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

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

4. When does the oil price affect the Norwegian exchange rate?

Abstract

Major changes in the Norwegian exchange rate have often coincided with large fluctuations in the price of crude oil. Previous empirical studies have however suggested a weak and ambiguous relation between the oil price and the exchange rate. In contrast to these studies, this essay explores the possibility of a non-linear relation between oil prices and the exchange rate. An examination of daily observations reveals a negative relation between the oil price and the nominal value of the currency. The strength of this relation depends on whether the oil price is below, inside or above the range of 14–20 US dollars a barrel. Moreover, it depends on whether the oil price is displaying a falling or rising trend. The relation is relatively strong when oil prices are below 14 dollars and are falling. These non-linear effects are tested and quantified within equilibrium correcting models of the exchange rate, derived on monthly and quarterly data to control for the influence of other macroeconomic variables. The models with non-linear oil price effects outperform similar models with linear oil price effects. The latter models grossly underestimate the exchange rate response to oil price changes in a state of low oil prices. The essay undertakes an extensive evaluation of the derived models to demonstrate the robustness of the results.

1. Introduction

The price of crude oil is commonly believed to have a significant influence on the Norwegian exchange rate. The Norwegian currency crises in the 1990s, i.e. the appreciation pressure in 1996/97 and the depreciation pressure in 1998/1999, have been attributed to the rise and fall of oil prices, see e.g. Alexander et al. (1997), Haldane (1997) and Norges Bank (1998) for details. Likewise, the large devaluation of the krone in 1986 is often explained with reference to low oil prices in 1985/86, see e.g. Norges Bank (1987, pp. 17).[1] The assumed link between the oil price and the value of the krone is based on the size of the petroleum sector relative to GDP, 10-20 % since the mid 1970s, and its relatively large share in Norway's total export of goods and services, see Aslaksen and Bjerkholt (1986) and Statistics Norway (1998). For example, in the period 1991-1997, Norway's oil production has been about 1.7 to 3 million barrels a day and oil and gas exports have made up more than 1/3 of its total export of goods and services.

A number of arguments can be put forward to explain why the nominal exchange rate of an oil producing country may appreciate when the oil price rises and depreciate when it falls. Firstly, higher oil prices increase demand of the currency of an oil exporting country and thereby raise its price relative to other currencies. Secondly, if the long run real exchange rate depends on oil prices, higher oil prices may create a wedge between the long run (equilibrium) real exchange rate and the actual real exchange rate, cf. Alexander et al. (1997).[2] Consequently, the nominal exchange rate may appreciate, even overshoot its equilibrium value if prices are sticky, to bring the actual real exchange rate in line with its equilibrium value, cf. Dornbusch (1976) and Mark (1990). Thirdly, if the real exchange rate is constant in the long run, as implied by the purchasing power parity (PPP) theory, higher oil prices may still bring about a short run appreciation of the real and nominal exchange rates through mechanisms that are well known from the Dutch disease literature, see e.g. Corden (1984). Accordingly, higher oil prices lead to a revaluation of petroleum

[1] In May 1986, the krone was devalued by 12 per cent relative to a trade weighted currency basket, mainly composed of (western) European currencies, see Norges Bank (1987, pp. 35-38) for details about the composition of the basket. The appreciation in 1996/97 and depreciation in 1998 were of around 10 per cent to the ECU.

[2] The real exchange rate may depend on the oil price indirectly through other variables that are affected by changes in oil prices, e.g. the stock of net foreign assets and current account.

wealth and increase revenues from the oil exports, see Golub (1983). This wealth and income effect can increase aggregate consumption and raise the demand of (internationally) traded and non-traded goods. As a result of higher demand of the latter goods, domestic prices may rise and place appreciation pressure on the real exchange rate, and thereby induce a transfer of resources from the sector of tradables to the sector of non-tradables.[3] Due to sticky prices, however, the nominal exchange rate may appreciate in the short run and speed up the real exchange rate appreciation to levels consistent with the temporary transfer of resources between the sectors, see Bruno and Sachs (1982). In the long run, however, the nominal exchange rate is not directly related to other variables than domestic and foreign prices. When oil prices fall, the arguments above can be reversed to explain depreciation pressure.

Empirical studies have, however, provided mixed support for the assumed covariance between the oil price and the Norwegian exchange rate, see e.g. Bjørvik et al. (1998) and Akram and Holter (1996).[4] These studies find a statistically insignificant and/or numerically weak relation between the oil price and the value of the krone. Figure 1.1, which shows a cross plot between the Brent Blend oil price in US dollars and the krone/ECU exchange rate (indexed) together with the associated regression line, illustrates the existing empirical results. Contrary to the theory, the cross plot does not indicate any obvious relation between the oil price and the Norwegian exchange rate. The regression line even indicates a small positive covariance and not a negative one as expected. Such empirical findings are puzzling in the light of the theoretical literature and the widely shared belief that the oil price has been an important factor behind the major fluctuations in the value of the krone during the 1990s and the devaluation in 1986.

However, the empirical results can be interpreted in two ways. One interpretation is that the "true" relation between the oil price and the value of the Norwegian currency is weak, at best, and the empirical results are a reflection of this fact. Hence the common belief has no firm ground. Indeed, empirical studies of the oil price and exchange rates

[3] The real exchange rate is defined as $R \equiv E(P^f/P)$. E denotes the nominal exchange rate, i.e. the price of foreign currency in terms of domestic currency, while P^f and P symbolises the foreign and domestic price levels, respectively.

[4] Regretfully, there does not seem to be any study published in English on this issue.

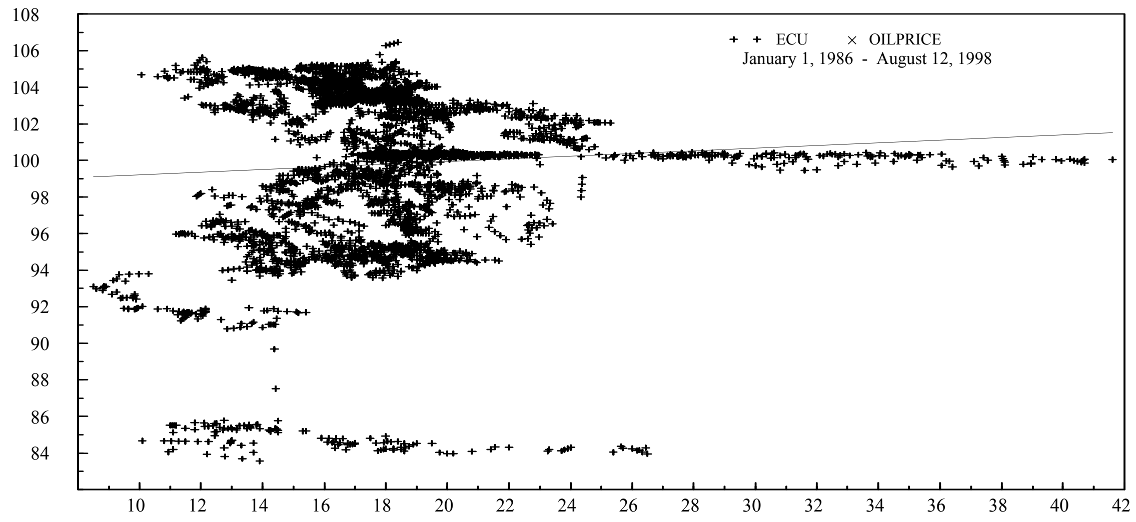


Figure 1.1: *Cross plot of the ECU index, an index for the krone/ECU exchange rate, and the price of crude oil in US dollars (horizontal axis) together with a regression line. The plot is based on 4608 daily observations over the period 1.1.1986-12.8.1998.*

of other countries often report an unstable relation between these variables, characterised by changes in the sign and size of coefficients over different sub-samples, see e.g. Shazly (1989) and De Grauwe (1996, pp. 146-149). Furthermore, at least for the appreciation pressure in 1996/97, an alternative explanation is offered by e.g. Kvilekval and Vårdal (1997). It is argued that the appreciation pressure arose as a result of higher interest rates in Norway relative to those in the EU countries throughout 1996, cf. Figure 4.1 in the appendix of this essay.

The second interpretation is that the common belief is not baseless, but the puzzle arises from the empirical approach towards estimating the relation between oil prices and the exchange rate. The present essay tests for this second interpretation.

A common feature of most studies that measure the link between the oil price and the exchange rate, including those conducted on Norwegian data, is that they (implicitly) assume symmetric effects on the exchange rate from an increase and a decrease in the oil price. Furthermore, the oil price effects are assumed to be independent of the level of oil prices. Accordingly, (log) linear models are employed to estimate their effects on a given exchange rate. This study questions whether linear models, imposing symmetric oil price effects, tend to underestimate oil price effects on the Norwegian exchange rate and hence fail to explain major changes in the exchange rate in the face of large fluctuations in oil prices.

A non-linear relation between oil prices and the Norwegian exchange rate seems to be reasonable in the light of the Norwegian monetary policy and the role of the central bank in its conduct. Since 1972, the Norwegian monetary policy has been aimed at exchange rate stabilisation against (western) European currencies, see Alexander et al. (1997), Norges Bank (1987) and (1995) for details and overview. In this monetary policy framework, the nominal exchange rate will display (excessive) fluctuations, due to appreciation or depreciation pressure arising from changes in e.g. oil prices, only if the central bank is unable or unwilling to ensure stability in the exchange rate. It follows that one is more likely to observe a negative relation between oil prices and the value of the krone when the authorities abandon the practice of currency stabilisation. Studies of currency crises suggest that a central bank is often more willing to and capable of resisting pressure for currency appreciation than depreciation pressure, cf. Flood and Marion (1998) and the references therein. This asymmetry is explained by pointing to the higher costs of resisting depreciation pressure than appreciation pressure. The costs are usually measured in terms of sacrifices of objectives other than exchange rate stabilisation pursued by a central bank. These may be concerns for unemployment, competitiveness, economic growth, inflation and/or the viability of financial institutions due to its role as a lender of last resort, cf. Obstfeld (1990) and Calvo (1998).[5]

The form of a possibly non-linear relation between oil prices and the Norwegian exchange rate is not known and has to be assumed. This is however a general problem when non-linear relations between variables are considered and not specific to this case. Moreover, tests of a linear relation against a non-linear relation are often designed to have power against specific non-linear forms, see e.g. Teräsvirta et al. (1995). To avoid making *a priori* assumptions about the form of a possibly non-linear relation between the oil price

[5] Generally a trade off will exist between realisation of these additional objectives where it may appear less costly for a central bank to e.g. lower interest rates in the face of appreciation pressure than raise interest rates in the face of depreciation pressure. Especially, if it is more concerned with the "side effects" on activity level than on inflation.

and the exchange rate, this essay starts out with an examination of the observed values of these variables using graphs and basic descriptive measures. The findings from this analysis are thereafter formalised and tested within the framework of multivariate models of the Norwegian exchange rate.

The essay proceeds as follows: The next section (2) examines daily observation of the krone/ECU exchange rate (hereafter referred to as the ECU index) and the oil price over the period January 1986-August 1998, in search for empirically stable patterns.[6] A regular pattern is likely to emerge more clearly in daily observations due to their large number than in observations collected at lower frequencies. The choice of the ECU index reflects the Norwegian policy of exchange rate stabilisation against the ECU during the 1990s. The examination turns out to reveal a non-linear, or state dependent, relation between the oil price and the ECU index. This bivariate analysis is, however, unable to control for the influence of other exchange rate determinants that might explain the apparent non-linearity.

This limitation is overcome in Section 3, which tests the findings from the bivariate analysis and estimates the non-linear oil price effects using equilibrium correction models (EqCMs) of the exchange rate, see Hendry (1995). To cross-check the findings, this section models the ECU index using monthly data over the period 1990:11 to 1998:11 and the nominal effective exchange rate ($\overline{E}$) using quarterly data over the period 1972:2 to 1997:4. The quarterly data set covers almost all oil price shocks in the OPEC era and exchange rate fluctuations since the end of the Bretton Woods system. Thus, the model of $\overline{E}$ enables a sound assessment of the results implied by the bivariate analysis and the model of the ECU index.

In addition, both the ECU index and the $\overline{E}$ are modelled using linear and non-linear specifications of oil price effects. The models with linear oil price effects serve as our reference models and help us to judge whether a change in the representation of oil price effects leads to better model properties and different estimates of the oil price effects.

Furthermore, we undertake an extensive evaluation of the models with non-linear oil

[6] All empirical results and graphs are obtained using PcGive 9.10 and GiveWin 1.24, see Hendry and Doornik (1996) and Doornik and Hendry (1996).

price effects to examine the robustness of the obtained results. In particular, we investigate whether our preferred model and the implied oil price effects remain invariant when exposed to additional information in the form of extra variables and observations not used in the derivation of the model. Also, we compare its merits against an alternative model with linear oil price effects, but with deterministic variables to account for the apparent non-linearity; focusing on their in-sample and out-of-sample explanatory power, in particular.

Section 4 reiterates the main findings while the appendix contains precise definitions of the variables, their sources, graphs and tables with their time series properties.

2. Empirical regularities

The analysis in this section is based on daily observations of the ECU index and the Brent Blend spot oil price in US dollars per barrel. The ECU index represents the krone/ECU exchange rate with $100 = 7.9440$, which refers to the central value of the krone/ECU rate when the krone was pegged to the ECU on October 22, 1990. The sample consists of 4608 observations covering the period from January 1, 1986 to August 12, 1998.

Before embarking on the descriptive analysis, it should be kept in mind that this sample contains observations from two different exchange rate and capital mobility regimes, which are likely to affect the observed relation between the oil price and the ECU index over the sample period. On the one hand, the krone was stabilised against the trade weighted basket of currencies ($\overline{E}$) before the peg to the ECU in October 1990, which allows a possible covariation between the ECU index and the oil price to emerge more clearly than during the 1990s (when the krone was more closely linked to the ECU). But on the other hand, the Norwegian foreign exchange rate regulations were not dismantled before July 1, 1990. These limited the capital mobility between Norway and other countries and thereby the fluctuations in the exchange rate, see Olsen (1990). Thus it is not obvious whether possible covariation between the krone/ECU exchange rate and the oil price is allowed to emerge more clearly during the 1990s or during the 1980s.

Subsection 2.1 characterises the ECU index and the oil price over the sample and examines their time series properties. Subsection 2.2 reports some patterns in the bivariate

Table 2.1: Testing the presence of a unit root in the ECU index and OILPRICE

Variable	ADF(d)	$\widehat{\varrho}$	*t-value*
ECU	5	-0.0019	-2.702
ECU^{2ID}	5	-0.0020	-2.945*
OILP	5	-0.0047	-3.338*

Note: The ADF-tests employ daily observations over 01.01.1986-12.08.1998, see Table 4.1 in the appendix for details.. * indicates significance at the 5% level. The critical values at 5% and 1% are -2.863 and -3.435, respectively. 2ID indicates that 2 impulse dummies have been used to adjust for the break in the series on the 10. and 11. December 1992.

relation between the oil price and the ECU index which clearly suggest a non-linear relation between the oil price and the exchange rate.

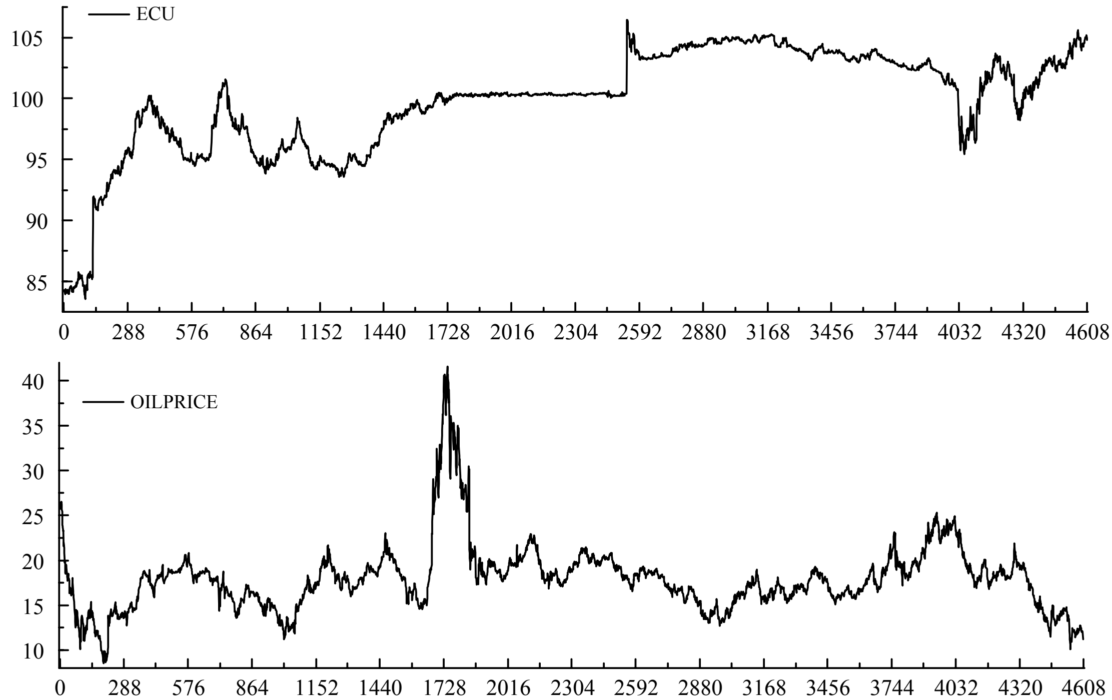
2.1. The exchange rate and the oil price

The daily observations of the ECU index and the oil price are displayed in Figure 2.1. There are relatively large swings in the index in the beginning and in the last part of the sample, especially from the end of 1996 (about observation 4000) and onwards. The early part of the sample covers the devaluation of the krone in May 1986 while the latter part of the sample covers the appreciation of the krone in 1996/97 and the depreciation in 1998. The period of the formal peg from October 22, 1990 to December 10, 1992 (from observation 1756 to observation 2537) is distinguished by a high degree of stability. The variability in the index increases after the abandonment of the formal peg. However, it continues to be relatively small compared with the period before 1990, particularly before the autumn of 1994: the interval before observation 3168.

The lower part of Figure 2.1 shows that the oil price has mainly fluctuated in the range of about 14-20 dollars (per barrel), see also the histograms in Figure 2.2. Most of the prices outside this band can be confined to specific periods. Prices below 14 dollars occur mostly in 1986 and 1998. During these periods oil prices even fell below 10 dollars. Prices in excess of 20 dollars are mostly from the Gulf war period in 1990/91 and from 1996/97 when they increased up to 42 dollars and 25 dollars, respectively.

The overall impression is that both the ECU index and the oil price can be characterised

Figure 2.1: *ECU index (above) and the price of crude oil. Daily observation from 01.01.1986 to 12.08.1998, 4608 observations.*



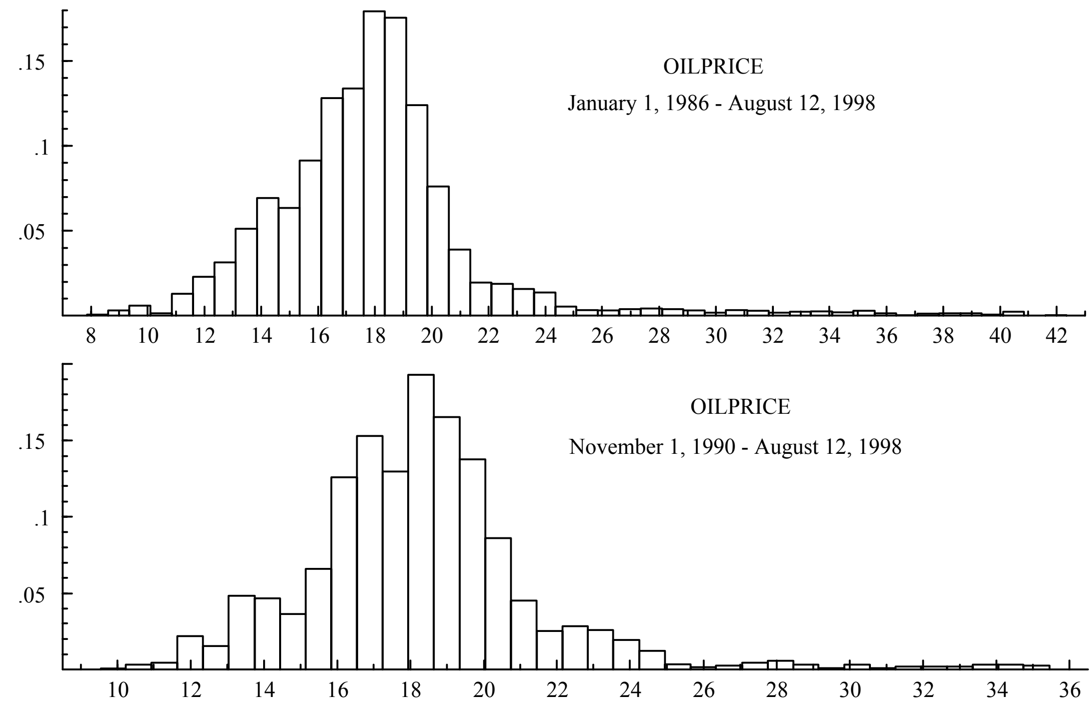


Figure 2.2: *Histograms of the daily observations of the oil price using the whole sample (above) and those from the period of peg to the ECU (bottom).*

as mean reverting processes, especially if one accounts for the break in these series. This impression is supported by the results in Table 2.1, which reports the result of augmented Dickey-Fuller (ADF) tests, see e.g. Banerjee et al. (1993, ch. 4). The null hypotheses are of unstable (or integrated) processes for the ECU index and the oil price. The null hypothesis for the ECU index is not rejected at the strictly 5% level of significance, but is rejected when the breaks in December 1992 are accounted for.[7] The latter result is as expected in the light of the Norwegian policy of exchange rate stabilisation. It is well known that an ADF test tends to underreject the null hypothesis when there are breaks in a series, see Perron (1989).

The null hypothesis of an integrated oil price process is rejected at the 5% level of significance, even when the exceptionally high oil prices during the Gulf War are not controlled for. The result is consistent with Horsnell and Mabro (1993, pp. 186) who use about three years of daily observations, and with studies based on longer samples of data, see e.g. Perron (1989) and Green et al. (1996). Given the support for a mean reversion property in the oil price, the range 14-20 dollars can be interpreted as the *normal range* of the oil price in the sample period.

2.2. Covariance between the exchange rate and the oil price

This subsection takes a closer look at the bivariate relation between the oil price and the ECU index. It points out that these variables generally display a negative covariance but the strength of this covariance depends on the level of the oil price and on whether or not the oil price is falling.

Figure 2.3 suggests that, in general, the covariance between the oil price and the ECU index is negative and relatively strong when the oil price moves outside the range of about 14-20 dollars, hereafter referred to as the normal range, but becomes negligible with a positive or negative sign when it fluctuates inside this range. As noted above, there seem to have been four (main) periods with prices outside the normal range, 1986 and 1998 with prices below 14 dollars and 1990/91 and 1996/97 with prices above 20 dollars. In three of

[7] One can argue for the use of more conservative critical values since the test is based on a model with more deterministic variables than in the standard case, cf. Banerjee et al. (1993, ch. 4).

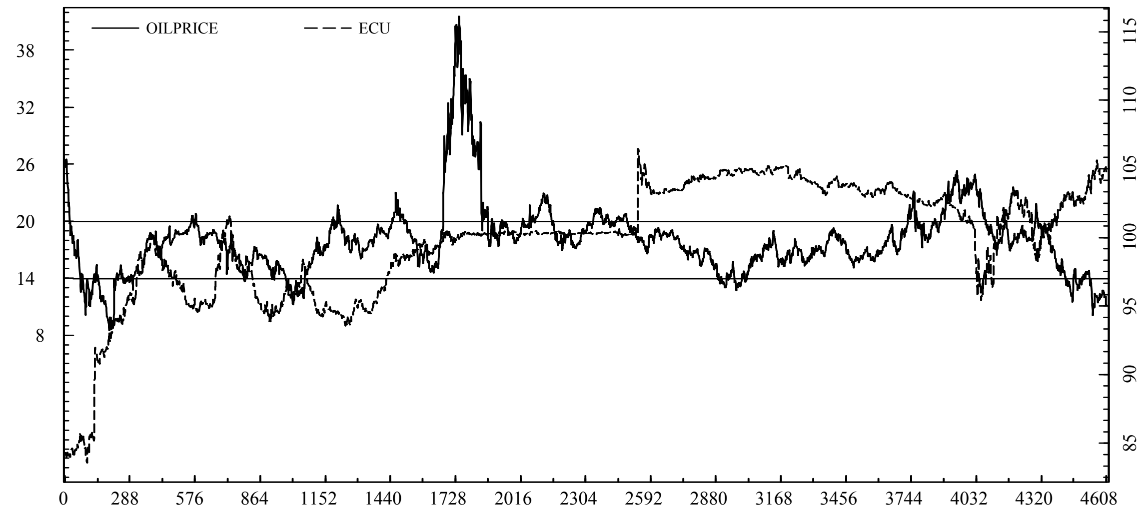


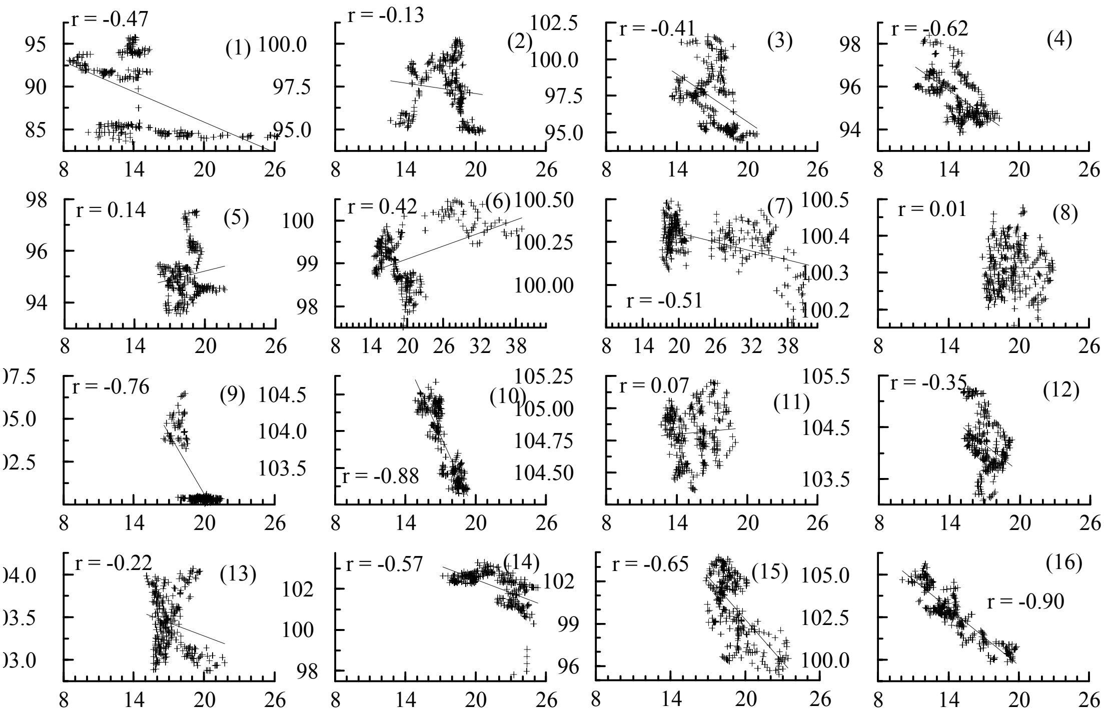
Figure 2.3: *ECU index (dashed) and the oil price from 01.01.1986 to 12.08.1998. The ECU index is mean and variance adjusted to the oil price.*

these periods, movements in the oil price coincide with large fluctuations in the exchange rate. More specifically, both the devaluation in 1986 and the depreciation in 1998 coincide with prices below 14 dollars while the appreciation in 1996/97 coincides with prices above 20 dollars. Note also that after the appreciation pressure, the ECU index seems to crawls back to the pre-appreciation level and this appears to coincide with a return of the oil price to the normal range. However, the unprecedented high oil prices during the Gulf war in 1990/91 do not lead to any noticeable appreciation of the krone measured by the ECU index. The krone was formally pegged to the ECU in this period, but one may still wonder at the absence of any considerable appreciation pressure during this period.[8] The positive covariance or zero/negligible covariance can be clearly observed in the observations up to 1700, which corresponds to the period before January 1990.

Figure 2.4 focuses on the covariation between the ECU index and the oil price in different sub periods. It displays the covariation between these variables in 16 equally sized samples consisting of 288 (= 4608/16) observations. Each sample covers a non-

[8]Norges Bank (1990, pp. 145) records a net purchase of foreign currency equivalent to about 6 billion NOK during August and the first half of september 1990. This in an effort to avoid the strengthening of the krone because of "the higher oil prices". During the appreciation pressure in 1996/97, however, the banks net purchase of foreign currency was equivalent to about 75 billion NOK, see Norges Bank (1997).

Figure 2.4: *Cross plots of the ECU index on the oil price using non-overlapping samples of equal size. These samples are derived by splitting the 4608 daily observations from January 1, 1986 to August 12, 1998 into 16 subsamples. Each of them consists of 288 observations and covers a period of about 9 and 1/2 months.*



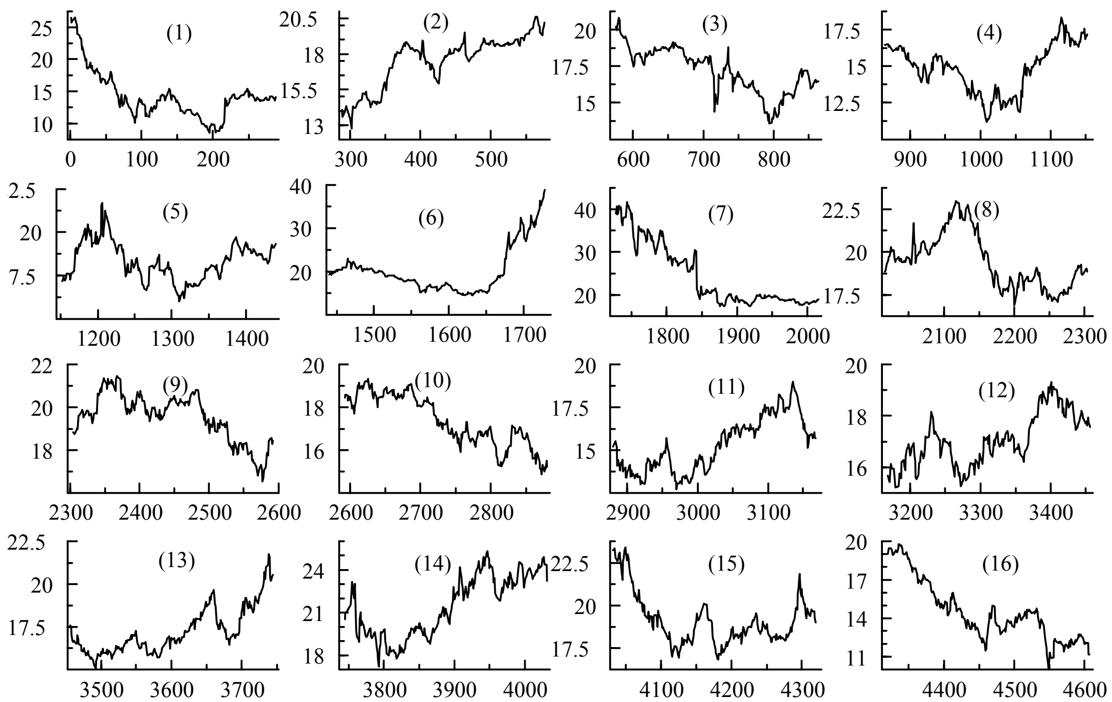


Figure 2.5: *The oil price in the 16 non-overlapping periods from January 1, 1986 to August 12, 1998, cf. figure 2.4.*

overlapping period of about 9 1/2 months. For example, sample 1 consists of the first 288 observations from January 1, 1986 to October 15, 1986, while sample 2 consists of the next 288 observations from October 16, 1986 to July 30, 1987, and so on. The degree of correlation in each sample is also reported. Figure 2.5 displays the level of the oil price in the corresponding samples. For instance, sample 1 in Figure 2.5 plots the first 288 observations of the oil price over the period January 1, 1986 to October 15, 1986, and so on.

Figure 2.4 confirms the impression from Figure 2.3 but also adds some new insight. It shows that:

- There is negative covariance in most of the samples but positive or negligible covariance in sample 5, 6, 8 and 11. In these samples the oil price is mostly inside the normal range, except in sample 6 where the positive covariance can be ascribed to the high oil prices during the Gulf War.

- The strength of the negative covariance seems to depend on whether the oil price is inside or outside the normal range. It is quite weak in sample 2, 12 and 13 but stronger in sample 1, 9, 10, 14, 15 and 16. In the former samples, the oil price is mostly inside the normal range, while the latter samples, with the exception of sample 10 and perhaps 15, contain a relatively large number of oil price observations outside the normal range. In sample 10, the observations are mostly inside the normal range.

- The negative covariance seems to be stronger when the oil price is falling compared with when it displays a rising trend. For example, samples 10 and 9 may be compared with 12 and 13, respectively. In the first pair of samples (10 and 12) and in the second pair (9 and 13), oil prices fluctuate in approximately the same price ranges, see Figure 2.4. However, as is evident from Figure 2.5, the oil price displays a falling trend in the periods covered by samples 10 and 9 and a rising trend in the periods covered by samples 12 and 13. Figure 2.4 shows that the negative correlation is stronger in samples 10 and 9 compared with the correlation in samples 12 and 13.

- The negative covariance seems to decrease with the level of the oil price. Figure 2.4 shows that the spread of observations around the regression lines is wider at higher oil prices than at lower oil prices. This is especially apparent in samples 1, 5, 9, 13 and 14, when the oil price is around 20 dollars. This pattern is more pronounced in larger samples of the data, as shown in Figure 2.6.

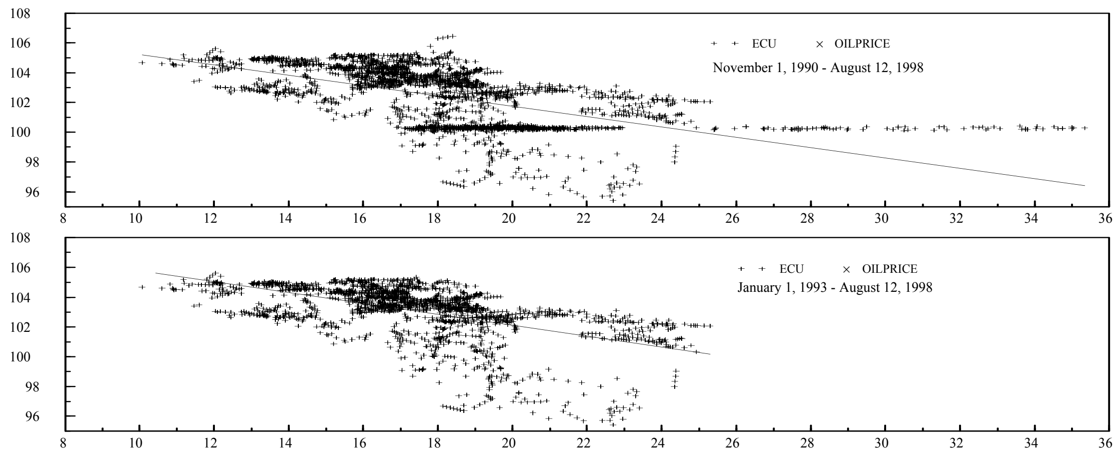


Figure 2.6: *Cross plots of the ECU index on the oil price using daily observations. The straight lines are the corresponding regression lines.*

To summarise, the graphical analysis suggests both *level* and *trend* dependent oil price effects on the exchange rate. In general, there is a negative covariance between the oil price and the exchange rate. The degree of covariance, however, is stronger when the oil price is outside the normal range of about 14-20 dollars than when it is inside this range, which appears as the normal range of oil prices in this sample. The covariance also shows a tendency to decrease with the level of the oil price, which also implies that the oil price effect is stronger when oil prices are below the normal range compared with when they are above this range. In addition to the level effects, the covariance seems to become stronger when the oil price is on a downward trend rather than on an upward trend. Thus, the covariance appears to be negligible when the oil price is inside the normal range, unless it displays a falling trend.

The bivariate analysis of this subsection, however, does not control for the possible influence of other exchange rate determinants. Hence it offers a potentially biased impression of the relations between the oil price and the exchange rate. The next section makes attempts to correct for this.

3. Multivariate exchange rate models

It is well documented that fluctuations in oil prices lead to considerable changes in macroeconomic variables, see e.g. Hamilton (1983) and Mork et al. (1994). Some of these variables, such as the current account, interest rates, level of economic activity and inflation are also regarded as important determinants of nominal exchange rates, see e.g. Frankel and Rose (1995). Because of the correlation between macroeconomic variables and oil prices, partial effects of the oil price on the Norwegian nominal exchange rate may be quite different from those indicated by the bivariate analysis. Moreover, the non-linear effects of the oil price might have emerged due to our failure to control for the influence of these variables in the preceding analysis. It cannot be precluded that once they are accounted for, the effects of oil prices, if any, are linear and independent of the level and the trend in the oil price.

The purpose of this section is to control for the influence of potentially relevant variables and factors when testing: (a) whether oil prices have non-linear effects on the Norwegian nominal exchange rate, (b) whether the effects are *level* and *trend* dependent, as suggested by the bivariate analysis, and (c) whether a linear representation of oil price effects leads to underestimation of oil price effects on the exchange rate.

To this end, we derive single equation multivariate models of the exchange rate; specifically, equilibrium correcting models (EqCMs) of the Norwegian nominal exchange rate. These models are not derived from a particular exchange rate theory but from quite general models containing variables that are interpretable within different exchange rate theories. The general models are thereafter simplified by following a "general to specific" modelling strategy in which parsimony is sought through data based coefficient restrictions, see Hendry (1995). The exchange rate literature is quite pessimistic with regard to the ability of macroeconomic variables to explain exchange rate movements, see e.g. Frankel and Rose (1995). However, most empirical studies confine their attention to variables and parameter values implied by a preferred exchange rate theory. Relatively general models without a priori coefficient restrictions seem to be better equipped to explain exchange rate movements and to serve their purpose in the present context.

Secondly, to convince ourselves that the obtained results are not an artefact of a given data sample, a model or are regime specific, this section models relative changes in both the ECU and the trade weighted nominal exchange rate ($\overline{E}$). The ECU index is modelled on a monthly data set that covers the period 1990:11 to 1998:11, while the $\overline{E}$ (also indexed) is modelled on a quarterly data set covering the period 1972:2-1997:4. The monthly data set is from the period with peg to the ECU while the quarterly data set covers almost the whole history of the Norwegian exchange rate since the end of Bretton Woods system, and all the major oil price shocks in the OPEC era. The data sets also differs with regard to capital mobility regimes. In contrast to the quarterly data set, the monthly data set only covers the period with unregulated capital mobility. The Norwegian foreign exchange regulations were gradually removed in the second half of 1980s and were fully dismantled by July 1990, see Olsen (1990).

Thirdly, we attempt to isolate the contribution of non-linearisation on the estimates of oil price effects and models' properties by deriving models with linear and non-linear oil price effects for both the exchange rate indices. The two models of a given exchange rate index are similar to each other but for the representation of oil price effects. The differences in the estimates of oil price effects and in the properties of models across a given pair of models may therefore be ascribed to differences in the representation of oil price effects; despite potential shortcomings with a given pair of models such as possible bias owing to our use of single equation models and due to neglect of potentially relevant explanatory variables. Note that the focus on single equation models implicitly assumes that the conditioning variables are weakly exogenous for the parameters of interest, in this case the coefficients of oil prices, see Engle et al. (1983). A violation of this assumption can lead to inconsistent and inefficient estimates of the parameters of interest. A test of this assumption however requires models of all the conditioning variables, which are beyond the scope of the present essay. Though we cannot claim that our estimators of the oil price effects provide consistent and efficient estimates of the *partial effects* of oil prices on the exchange rate, or of any other variable, possible differences in the oil price effects across the pairs of models still enable us to address the issues that are the focus of this study: (a), (b) and (c) noted above.

Finally, to further substantiate the results, we undertake an extensive evaluation of the models with non-linear oil price effects. Specifically, they are exposed to a battery of tests aimed at testing whether they are well specified and especially whether the oil price effects are adequately characterised. Also, their explanatory power is measured against the models with linear oil price effects. In addition, we consider whether the quarterly model (with non-linear oil price effects) remains intact when we include variables that are neglected during its derivation, and whether it remains stable when reestimated on an extended data set. The latter data set contains new observations for 7 quarters over the period 1998:1-1999:3, a period in which oil prices and the Norwegian exchange rate have displayed excessive fluctuations. Furthermore, the properties of the quarterly model are compared against a model with linear oil price effects but with deterministic variables to account for the apparent non-linearity. Properties that are focused upon are: the in-sample and out-of-sample explanatory power.

The remainder of this section is organised as follows: The general model is formulated in Subsection (3.1), which also motivates the choice of variables. Subsection (3.2) considers the models with linear oil price effects and formally tests the appropriateness of a linear representation of oil price effects, in particular. Subsection (3.3) formulates and derives the models with non-linear oil price effects while Subsection (3.4) undertakes the evaluation.

3.1. A general EqCM of the exchange rates

Equation (3.1) presents a general EqCM of the nominal exchange rate e_t, where $e = ecu$, $\overline{e}$.

$$
\begin{aligned}
\Delta e_t &= \alpha_0 - \phi[e - (cpi - cpi^f)]_{t-1-j} + \beta_1 \left[R - R^f\right]_{t-j} + \\
&\sum_{j=0}^{p}[\alpha_i \Delta e_{t-1-j} + \pi_{1j}\Delta cpi_{t-j} - \pi_{2j}\Delta cpi^f_{t-j} - \beta_{2j}\Delta R_{t-j} \\
&+\beta_{3j}\Delta R^f_{t-j} + \mu_j \Delta FI.Y_{t-j} + \Gamma_j \mathbf{Z}_{t-j}] + \Psi \mathbf{B}_t \\
&+ f_t(\sum_{j=0}^{p}\Omega_{1j} oilp_{t-j},\ \sum_{j=0}^{p}\Omega_{2j}\Delta oilp_{t-j}) + v_t. \qquad (3.1)
\end{aligned}
$$

Variables in small letters indicate that they are natural logs of the original variables, e.g. *ecu* is the natural log of the ECU index. Δ denotes a change over one month or a quarter while p indicates the number of lags. These will be determined during the estimation of the models' parameters represented by the Greek letters and supposed to be positive and constant over time. The appendix provides precise definitions of the variables, their graphs and reports their time series properties, i.e. whether a given variable is stationary or non-stationary in unit root sense. As evident from the tables in the appendix, the general model appears to be balanced, since the left hand side variables and the right hand side variables and terms can be characterised as stationary processes, see Banerjee et al. (1993, Ch. 3). Finally, v_t is the residual assumed to be independently, identically and normally distributed with zero mean and variance σ^2, i.e. IIDN(0, σ^2).

The term $[e - (cpi - cpi^f)]$, possibly in addition to the constant term α_0, represents a deviation from the equilibrium level of the nominal exchange rate under the PPP hypothesis. Accordingly, the nominal exchange rate reflects the ratio between domestic and foreign prices in the long run. In general, the empirical evidence in favour of this assertion is mixed, see e.g. Rogoff (1996) for a survey of the literature. However, it has not been rejected when tested for between Norway and its trading partners, see Edison and Klovland (1987) and Akram (2000a).[9]

[9] Klovland and Edison (1987) use annual data from 1874 to 1971 to model the Norwegian prices. They do not reject the PPP-hypothesis between Norway and UK when using a model that controls for the excessive exchange rate volatility in the period 1914 and 1928.

Akram (2000) tests for PPP between Norway and its trading partners using the same quarterly data set as employed in the present study. It is shown that the proportonality between the exchange rate and

Short run fluctuations in the exchange rate can be attributed to variables that affect the demand of domestic assets relative to foreign assets, cf. portfolio balance models of exchange rates, see e.g. Hallwood and MacDonald (1994, pp. 186-205). The exchange rate tends to appreciate when the demand of domestic assets increases relative to the demand of foreign assets, i.e. assets denominated in domestic and foreign currencies, respectively. This is likely to take place when the opportunity costs of holding foreign assets increase or are expected to increase. The opportunity costs of holding domestic assets increase due to expected depreciation of the domestic currency, a rise in foreign interest rates and/or due to a fall in domestic interest rates.

The model allows for short run effects of domestic and foreign prices, represented by the inflation rates Δcpi and Δcpi^f. If PPP holds, the difference between Δcpi and Δcpi^f can be interpreted as expected depreciation. Expected depreciation may also be proxied by the spread between domestic and foreign interest rates. According to the *uncovered interest rate parity (UIP)* hypothesis, the interest rate differential is equal to expected depreciation.[10]

Changes in domestic and foreign interest rates, ΔR and ΔR^f, are supposed to capture changes in the opportunity cost of holding domestic assets rather than foreign assets. When modelling *ecu* on monthly data from the 1990s, we employ three month Norwegian and European money market rates: $R = R3$ and $R^f = R3^f$. However, when modelling $\overline{e}$ on quarterly data since the early 1970s, we employ the Norwegian government bond rate and the trade weighted government bond rate: $R = RB$ and $R^f = RB^f$. The domestic bond rate was chosen in preference to the money market rates as the latter displays a quite erratic behaviour until the end of the 1970s, probably because of a thin domestic money market and regulations of international capital flows, see Figure 4.2 in the appendix. The use of the government bond rate in trading countries is motivated by the policy of exchange rate stabilisation against the currencies of a majority of these countries. This policy is expected to entail a close link between the domestic and the foreign interest rates. However, the

the relative consumer price indices between Norway and its trading partners is not rejected, either in a univariate or in a system framework.

[10] If UIP does not hold, the spread may reflect risk premia in addition to expected depreciation, see e.g. Gibson (1996) and Hallwood and MacDonald (1994).

spread between these interest rates was relatively large in the 1970s and 1980s, see Figure 4.2. These can be partly ascribed to the Norwegian capital regulations which were not fully dismantled before 1990 and were relatively tight during the 1970s, but may also reflect devaluation expectations and risk premia due to the frequent devaluations in this period, see Alexander et al. (1997).

In the model, $\Delta FI.Y$ is foreign net financial investment in Norway relative to the Norwegian GDP (Y). Financial investment abroad tend to place upward pressure on the value of the domestic currency. Hence a rise in $\Delta FI.Y$ is expected to bring about exchange rate depreciation.

The vector **Z** represent variables such as changes in the activity level and productivity growth at home and abroad; and different measures of government expenditure at home. Higher activity level and productivity growth at home relative to abroad are often believed to raise the value of the domestic currency, while a rise in domestic government expenditures is believed to have the opposite effect, cf. Balassa (1964), Samuelson (1964) and Gibson (1996).

The content of vector **Z** is sample dependent, due to the availability of data. On the monthly data set we are only able to consider the effects of the differences in the activity level and of changes in government expenditure. Here, the differences in the activity level are measured by changes in the registered rate of unemployment at home and abroad, Δu and Δu^f, while changes in government expenditures relative to GDP are taken into account by the government budget surplus, *Fisc.Y.* When employing the quarterly data set, however, we also consider growth in GDP (Δy) to represent changes in the activity level at home, as an alternative to Δu. In addition, we are able to allow for productivity growth at home and abroad. Productivity at home and abroad is measured by q and q^f, respectively, which are defined as natural logs of the inverse of unit labour costs in Norway (ULC) and in its trading partners (ULC^f): $q = ln(1/ULC)$ while $q^f = ln(1/ULC^f)$. The inverse of unit labour costs can be interpreted as value added per unit labour cost. Changes in government expenditures relative to GDP are taken into account by Δg where g is defined as the sum of government consumption (C_G) and gross real investment (J_G), relative to GDP (Y): $g = (C_G + J_G)/Y$. We also allow C_G/Y and J_G/Y

to enter separately in the model, i.e. without the homogeneity restriction.

The variables represented by the vector **Z** have, however, received mixed support in exchange rate models, see e.g. Frankel and Rose (1995). They are therefore not included in the general models of ecu and $\overline{e}$ at the outset, in order to avoid over-parameterisation relative to the number of observations, which is likely to be a problem if several lags of a given variable are included in the general models. However, we test for their significance upon reaching parsimonious versions of each model. This allows us to examine whether a preferred specification of a model is invariant to inclusion of additional variables.

Vector **B** contains a number of dummy variables to account for outliers and other extreme observations that remain unexplained by the variables explicitly included in the model and thereby reduce the potential omitted-variable bias in parameter estimates. They are also intended to ensure the validity of the residual assumptions.

Finally, the model allows the oil price to affect the exchange rate in both the short run and the long run, since both the relative changes in the oil price ($\Delta oilp$) and the log level of the oil price ($oilp$) are present. In the latter case, oil prices are allowed to create a wedge between the nominal exchange rate and the relative price ($cpi-\ cpi^f$) in the long run. Accordingly, the long run real exchange rate depends on the oil price in contrast to the PPP theory. Both the short run and long run oil price effects are included in an unspecified form. The analysis in the following subsections is essentially aimed at finding the appropriate specification of the oil price effects.

3.2. Models with symmetric oil price effects

Models with symmetric oil price effects, hereafter *linear models,* were formulated by inserting the following specification of $f_t(.)$ into model (3.1):

$$f_t(\sum_{j=0}^{p} \Omega_{1j} oilp_{t-j},\ \sum_{j=0}^{p} \Omega_{2j} \Delta oilp_{t-j}) = \sum_{j=0}^{p} [\Omega_{1j} oilp_{t-j} + \Omega_{2j} \Delta oilp_{t-j}]. \qquad (3.2)$$

The (general) models of Δecu_t and $\Delta \overline{e}_t$ were then estimated by OLS for a common lag length of 2 ($= p$). To allow for a linear approximation of possible non-linear oil price effects,

Table 3.1: An EqCM of the ECU rate with linear oil price effects.

$$\begin{aligned}
\widehat{Decu}_t = & \underset{(1.757)}{0.178} - \underset{(-1.336)}{0.033}\ [ecu - (cpi - cpi^f)]_{t-2} + \underset{(4.092)}{0.183}\ [R3 - R3^f]_t \\
& + \underset{(23.891)}{1.384}\ Decu_{t-1} - \underset{(-9.307)}{0.520}\ Decu_{t-2} - \underset{(-4.673)}{0.274}\ \Delta[R3_{t-2} - Dcpi_{t-1} \\
& - \underset{(-2.905)}{0.858}\ Dcpi^f_{t-1} + \underset{(2.158)}{0.008}\ DFI.Y_t \\
& - \underset{(-5.160)}{0.028}\ i93p12 - \underset{(-4.978)}{0.026}\ i97p1 + \underset{(3.892)}{0.021}\ i97p11 + \underset{(6.300)}{0.034}\ i98p1 \\
& - \underset{(-4.348)}{0.024}\ i98p3 - \underset{(-0.115)}{0.001}\ i98p6 + \underset{(4.605)}{0.026}\ i98p9 - \underset{(-7.815)}{0.049}\ i98p11 \\
& - \underset{(-3.942)}{0.017}\ oilp_{t-1} - \underset{(-3.313)}{0.031}\ \Delta oilp_t
\end{aligned}$$

Sample: 1990:11-1998:11, $T = 97$, $k = 18$. *Method: OLS*

Diagnostics

$\hat{\sigma}_L$	=	0.511%
Log lik	=	*473.3*23
$AR\ 1-6\ F(6,73)$	=	1.242[0.295]
$ARCH(6)\ F(6,67)$	=	1.075[0.386]
$Het.\ Xi^2\ F(26,52)$	=	0.640[0.891]
$Normality\ \chi^2(2)$	=	0.632[0.729]
RESET F(1, 78)	=	4.386[0.040]*
$TT_{\mathrm{L}}\ F(6,73)$	=	2.650[0.027]*

The t-values are in brackets (.) below the estimates and p-values are in large brackets [.] beside the test statistics. AR 1-6 F(df1, df2) tests for autocorrelation in the residuals up to 6 lags. df1 and df2 denote degrees of freedom. ARCH(6) F(df1, df2) tests for autoregressive conditional heteroscedasticity (ARCH) up to order 6, see Engle (1982). Het. Xi^2 F(df1, df2) tests for heteroscedasticity by using squares of regressors, see White (1980). The normality test with chi-square distribution is that by Jarque and Bera (1980). RESET F(df1, df2) is a regression specification test. It tests the null hypothesis of correct model specification against the alternative of misspecification, indicated by the significance of y^2, i.e. the square of the fitted value, in the model, see Ramsey (1969). TT_L F(df1, df2) is defined in the main text. Here and elsewhere in this study, a raised star * indicates rejection of the null hypothesis at the 5% level of significance, while two stars ** indicate rejection at the 1% level.

greater numbers of lags for $oilp_t$ and $\Delta oilp_t$ were also considered. Following a "general to specific" modelling strategy, variables that appeared to have numerically small and statistically insignificant coefficients were excluded from the models, in most cases, for the sake of parsimony. But some were also retained in the models to ease comparison with the non-linear models to be derived later. In addition to the exclusion restrictions, parsimony was also sought through symmetry restrictions on coefficients that had almost the same estimates but opposite signs. For example, coefficients of changes in the domestic interest rates and the inflation rate are restricted to make them interpretable as changes in the domestic "real interest rate".

Monthly changes in the ecu_t (Δecu_t) turned out to be difficult to model satisfactorily with a limited number of lags on the stochastic variables and relatively few deterministic regressors. In comparison, annual changes in the ecu_t ($Decu_t \equiv ecu_t - ecu_{t-12}$) could be characterised parsimoniously using e.g. annualised domestic and foreign inflation rates and foreign net financial investment in Norway, $Dcpi$, $Dcpi^f$ and $DFI.Y_t$, respectively. Still a number of deterministic regressors were needed to capture large fluctuations in the exchange rate, especially during the years 1997 and 1998. In contrast, quarterly changes in the trade weighted exchange rate $\overline{e}$ ($\Delta\overline{e}$) were straightforward to model, only requiring a (centered) dummy with 1 in 1997:1 and -1 in 1997:2 to capture the appreciation and the subsequent depreciation during the first half of 1997.

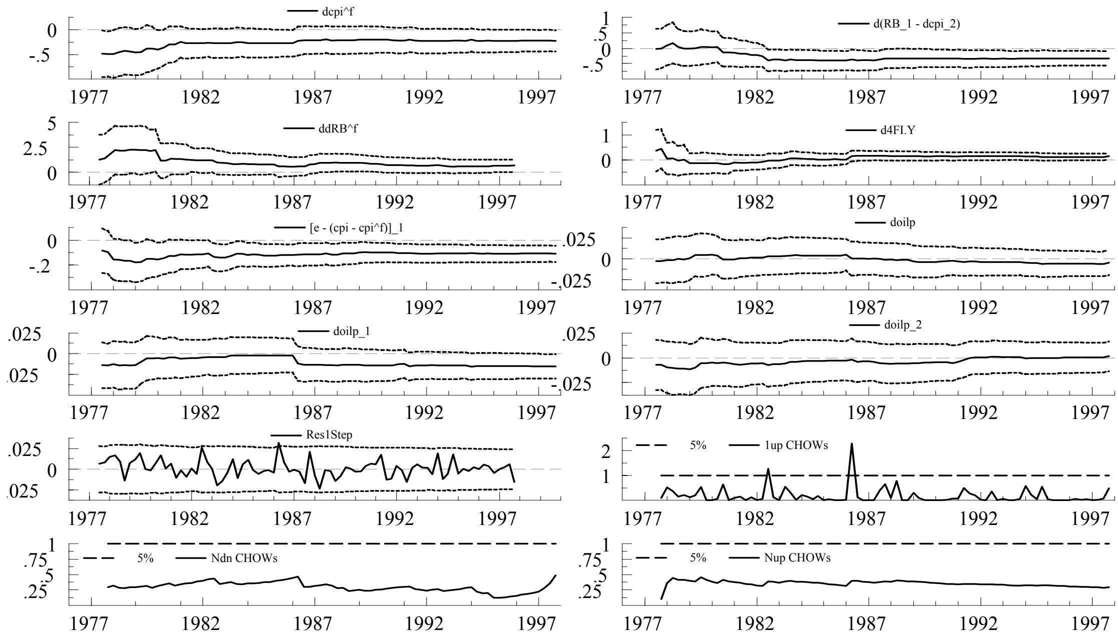


Figure 3.1: *Constancy statistics for the model of* $\Delta\overline{e}$ *with linear oil price effects, see Table 3.2. Initial estimation period is 1972:2-1976:4. Prefix "d" denotes the first difference* Δ*. The graph show the recursive coefficient estimates* $\pm 2SE_t$ *for the indicated regressors, One-step ahead residuals* $\pm 2SE_t$ *and Chow statistics for the model. The latter are scaled by their critical values at the 5% level.*

Table 3.1 and 3.2 present the relatively parsimonious versions of the models of $Decu_t$ and $\Delta\overline{e}_t$, respectively, and the associated model diagnostics. Table 3.2 also reports the

Table 3.2: An EqCM of the effective exchange rate with linear oil price effects.

$$
\begin{aligned}
\Delta\widehat{\overline{e}}_t = & - \underset{(-3.350)}{0.110} [\overline{e} - (cpi - cpi^f)]_{t-1} + \underset{(2.918)}{0.246} \Delta\overline{e}_{t-1} \\
& - \underset{(-2.712)}{0.324} \Delta(RB_{t-1} - \Delta cpi_{t-2}) - \underset{(-2.094)}{0.227} \Delta cpi^f{}_t \\
& + \underset{(2.216)}{0.699} \Delta^2 RB^f{}_{t-1} + \underset{(2.311)}{0.147} \Delta_4 FI.Y_t - \underset{(-3.952)}{0.035} id97q1 \\
& - \underset{(-0.659)}{0.005} \Delta oilp_t - \underset{(-2.073)}{0.015} \Delta oilp_{t-1} + \underset{(0.208)}{0.002} \Delta oilp_{t-2}
\end{aligned}
$$

Sample: 1972:2-1997:4, $T = 103$, $k = 10$. *Method: OLS*

Diagnostics

$\hat{\sigma}_L$	=	1.2%
Log lik	=	408.901
R^2	=	0.46
$AR\ 1-5\ F(5,88)$	=	0.532[0.752]
$ARCH\ (4)\ F(4,85)$	=	0.596[0.666]
$Het.Xi^2\ F(20,72)$	=	0.835[0.665]
$Het.XiX_j\ F(56,36)$	=	0.638[0.936]
$Normality\ \chi^2(2)$	=	1.260[0.533]
$RESET\ F(1,92)$	=	4.407[0.039]*
$TT_L F(9,84)$	=	2.227[0.028]*

Tests for omitted variables

			Δcpi_t	ΔRB_t	ΔRB^f_t	$(RB - RB^f)_{t-1}$	$FI.Y_{t-1}$
Coeff.	*estimate*		0.060	0.049	−0.598	0.002	0.036
Single:	$F(1,\ 88)$		[0.752]	[0.882]	[0.172]	[0.975]	[0.620]
Joint:	$F(5,\ 88)$	[0.806]					

Note: Brackets below the estimates contain t-values while square brackets contain p-values. The value of R^2 is almost the same if a constant term is included. The tests are explained in Table 3.1. Tests for omitted variables impose zero restrictions, individually and jointly, on the coefficients of the indicated variables when included in the presented model.

outcome of variable omission tests and the coefficient estimates of a number of variables when added (jointly) to the quarterly model.

In both models, positive changes in domestic "real interest rates", $\Delta[R3_{t-2} - Dcpi_{t-1}]$ and $\Delta(RB_{t-1} - \Delta cpi_{t-2})$, and foreign inflation place appreciation pressure on the exchange rates. A reduction in foreigners' financial investment in Norway, or equivalently increased Norwegian financial investment abroad, also implies exchange rate appreciation. The equilibrium correction terms have the expected signs in both models, but the equilibrium correction term in the monthly model appears statistically insignificant at the standard levels of significance. In the quarterly model, however, it is statistically significant at the 1% level, even if it is considered as non-stationary (integrated) under the null hypothesis and the critical values are found by the response function of MacKinnon (1991). In which case, the appropriate critical t-value at the 1% level is -2.585.[11]

The spread between the domestic and foreign interest, proxying depreciation expectations, is significant in the monthly model only. In the quarterly model, there was neither a contemporaneous nor a lagged effect from the spread between domestic and foreign bond rates, see e.g. the outcome of variable omission tests in Table 3.2. The effects of changes in foreign interest rate were found to be weak and largely insignificant in the monthly model. In the quarterly model, however, acceleration in foreign bond rates contributes to depreciation pressure in a statistically significant way. One possible explanation for their insignificance in the monthly model can be the high degree of correlation between domestic and foreign interest rates in the sample. The sample covers a period of deregulated capital markets in which the Norwegian government has pursued a stable exchange rate policy. Consequently, the Norwegian interest rates have by and large followed European interest rates for the sake of exchange rate stability. The collinearity between these interest rates may have contributed to the uncertainty in the estimated effects of foreign interest rates and thereby to their insignificance.

The policy of exchange rate stabilisation may also explain the persistence in the rate of depreciation in both models. Especially in the monthly model where the sum of the autoregressive coefficients is above 0.80, which implies long periods of depreciation or appreciation even in the absence of input from other variables. This points to *bandwagon effects* in the foreign exchange market, but may largely reflect the stable exchange rate policy pursued by the Norwegian government in this period.

The effects of oil prices seem to be mixed. In the monthly model, both the contemporaneous change and the lagged level of *oilp* appear statistically significant at the standard levels with negative coefficient estimates. Hence, oil prices have appreciations effects both in the short run and long run. In the quarterly model, however, positive changes in

[11] The 1% critical value for the case when there is no constant term, no trend and one non-stationary variable under the null hypothesis can be calculated as: $-2.5658 - 1.960/103 - 10.04/(103)^2 \approx -2.585$, see MacKinnon (1991).

the oil price place appreciation pressure on the exchange rate, but their effects are not significantly different from zero, perhaps with the exception of $\Delta oilp_{t-1}$. Its full sample estimate is barely significant at the 5% level, but its coefficient estimates on less than the full sample are not, see the recursive estimates in Figure 3.1. The figure also shows that the recursive estimates of the coefficients of $\Delta oilp_t$ and $\Delta oilp_{t-2}$ are quite close to zero in the period 1977:1-1997:4. Actually, there do not seem to be any effects of oil prices at all before 1986. On the basis of these findings one could conclude that oil prices do not have significant effects on the nominal exchange rate, even in the short run. This contrasts with the quite common belief that fluctuations in the oil price tend to have a strong impact on the Norwegian krone, but is nevertheless consistent with earlier empirical findings using linear models, see e.g. Bjørvik et al. (1998) and the references therein.

The diagnostic tests suggest that the residuals from each of the models satisfy the standard assumptions, that is, they appear as IIDN(,) as assumed. However, the regression specification test (RESET) rejects both model formulations at the 5% level. To test whether the rejection is due to possible misrepresentation of oil price effects, we apply the test for non-linearity suggested by e.g. Teräsvirta (1998). This test, denoted as TT_L, is performed on the residuals from each of the models to test the null hypotheses of linear oil price effects against the alternative hypothesis of non-linear oil price effects, with the oil price (*OILP)* as the transition variable. Although the null hypothesis is tested against the alternative hypothesis of neglected non-linearity of smooth transition type, the test also has power against the alternative of abrupt transitions. Table 3.1 and 3.2 show that the null hypotheses of linear oil price effects are rejected at the 5% level.

3.3. Models with asymmetric oil price effects

The non-linear oil price effects suggested by the graphical analysis in Section 2 can be characterised by the following specification of $f_t(.)$ with $c_1 = 14$ and $c_2 = 20$.

$$
\begin{aligned}
f_t(.) &= [\sum_{i=0}^{L}(\phi_i oilp_{t-i} + \widetilde{\phi}_i \Delta oilp_{t-i})] \times F_{t,\ Low} \\
+ \quad & [\sum_{i=0}^{H}(\theta_i oilp_{t-i} + \widetilde{\theta}_i \Delta oilp_{t-i})] \times F_{t,\ High} \\
+ \quad & [\sum_{i=0}^{D}(\psi_i oilp_{t-i} + \widetilde{\psi}_i \Delta oilp_{t-i})] \times F_{t,\ Diff}, \text{ where} \qquad (3.3) \\
F_{t,\ Low} &= [1 + \exp\{\lambda_1(OILP_t - c_1)\}]^{-1}, \qquad \lambda_1 > 0 \qquad (3.4) \\
F_{t,\ High} &= [1 + \exp\{-\lambda_2(OILP_t - c_2)\}]^{-1}, \qquad \lambda_2 > 0 \qquad (3.5) \\
F_{t,\ Diff} &= [1 + \exp\{\delta(OILP_t - OILP_{t-d})\}]^{-1}, \ \delta > 0. \qquad (3.6)
\end{aligned}
$$

The logistic functions $F_{t,\ Low}$, $F_{t,\ High}$ and $F_{t,\ Diff}$ are assumed to reflect the state of the oil price, i.e. whether it is below c_1 US dollar, above c_2 US dollars or below/above the price in period $t-d$, respectively. In addition to the Greek letters in (3.3)-(3.6), c_1 and c_2 are constant parameters while L, H and D denote number of lags in the different states. For (finite) values of λ_1, λ_2 and δ

$$
\begin{aligned}
F_{t,\ Low} &\longrightarrow 1 \text{ when } OILP \ll c_1 \text{ USD} \\
F_{t,\ Low} &\longrightarrow 0 \text{ when } OILP \gg c_1 \text{ USD} \\
F_{t,\ High} &\longrightarrow 1 \text{ when } OILP \gg c_2 \text{ USD} \\
F_{t,\ High} &\longrightarrow 0 \text{ when } OILP \ll c_2 \text{ USD} \\
F_{t,\ Diff} &\longrightarrow 1 \text{ when } \{OILP_t - OILP_{t-d}\} \ll 0 \\
F_{t,\ Diff} &\longrightarrow 0 \text{ when } \{OILP_t - OILP_{t-d}\} \gg 0.
\end{aligned}
$$

Note also that for sufficiently large values of λ_1 and λ_2

$$F_{t,\ Low} \longrightarrow 0 \text{ and } F_{t,\ High} \longrightarrow 0 \text{ for } c_1 < OILP < c_2.$$

This provides a mechanism to represent the state of the oil price when it is fluctuating in the range c_1-c_2 dollars.

Empirical specification of the logistic function requires estimates of the transition parameters λ_1, λ_2 and δ, the threshold values c_1 and c_2 and of the delay parameter d. The estimates of the transition parameters and the threshold values can be obtained by estimating model (3.1) with the non-linear specification of $f_t(.)$, using non-linear least square (NLS) or maximum likelihood (ML), in principle. The estimation requires plausible starting values for these parameters, however. The bivariate analysis in Subsection 2.2 provides plausible starting values for c_1 and c_2 by suggesting that they should be close to 14 and 20 US dollars, respectively. It seems reasonable that the transition parameters λ_1, λ_2 and δ are rather large, though it is difficult to decide upon exact values. Large values of these parameters lead to abrupt transition between 0 and 1 when $OILP_t$ deviates from c_1, c_2 or $OILP_{t-d}$, while small values of these bring about smooth transition between 0 and 1. The intuition for large values of these parameters is based on the observations that the major fluctuations in the Norwegian exchange rate have often coincided with fluctuations in the oil prices and that both oil prices and exchange rates often display abrupt transition from one level to another, in contrast to real economic variables. During estimation, however, we experiment with both low and high values of these parameters. The value of d can be chosen by comparing the explanatory power of the model estimated with different values of d, conditional on values of the transition and threshold parameters. The preferred estimates of these parameters and of the delay parameter can be selected after an iteration process: estimating λ_1, λ_2, δ, c_1 and c_2 conditional on a chosen d, which is thereafter revised conditional on the estimates of λ_1, λ_2, δ, c_1 and c_2, and so on.

Details of how the models with non-linear oil price effects were specified and estimated are provided below.

Estimation: The models with non-linear oil price effects were derived in two steps to ease the estimation of λ_1, λ_2, δ, c_1 and c_2. In the first step, the parsimonious models of $Decu_t$ and $\Delta\bar{e}_t$ in Table 3.1 and 3.2 were reformulated with the function $f_t(.)$ specified in (3.3)–(3.6), replacing the linear oil price effects. A common value of 2 for L, H, and D was

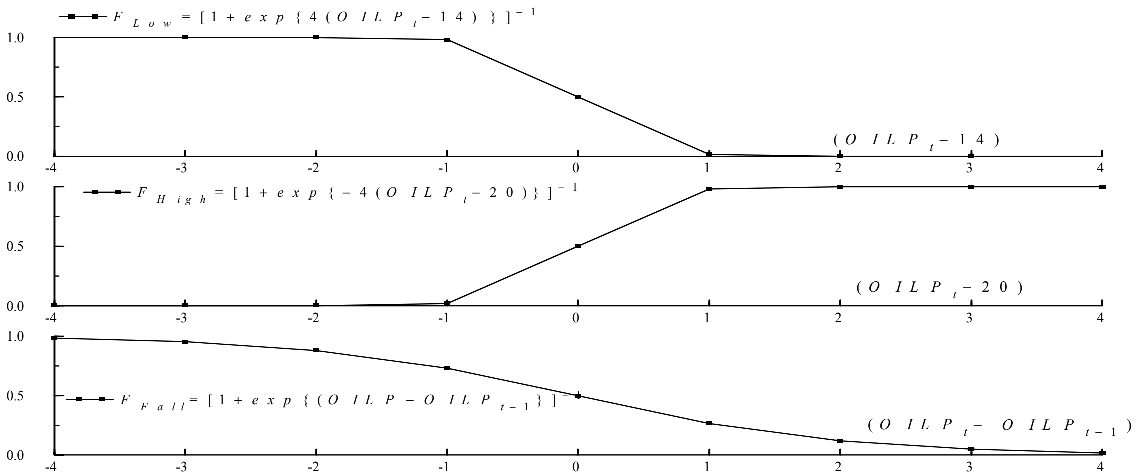


Figure 3.2: *The upper figure describes the value of F_{Low} for deviations of the oil price from 14 dollars. The figure in the middle shows the value of F_{High} for deviations of the oil price from 20 dollars. The figure at the bottom describes the value of F_{Diff} for deviations of the oil price from its value a year ago.*

chosen while the values of d were varied in the interval 1–12 and 1–4 in the case of the model on monthly and quarterly data, respectively. However, the obtained maximum likelihood estimates of λ_1, λ_2, δ, c_1 and c_2 were generally quite uncertain, even when the models were sequentially reduced (and degrees of freedom increased) by excluding oil price terms with numerically insignificant effects. One explanation could be that the uncertainty associated with these parameters reflects the difficulty in distinguishing between the shapes of a transition function, e.g. $F_{t,Low}$, at large values of the transition parameter, cf. Granger and Teräsvirta (1993) and Teräsvirta (1998). To accomplish this, a large number of observations around the threshold value, e.g. $c_1 = 14,$ is required. For example, it is quite difficult to distinguish between the shapes of $F_{t,Low}$ for $\lambda_1 = 4$ and $\lambda_1 = 5$.

The high degree of uncertainty associated with the estimates of the transition and threshold parameters led us to assess their values conditional on each other. Conditional on $c_1 = 14$ and $c_2 = 20$, values of λ_1, λ_2 and δ equal to 4, 4 and 1, respectively, increased the explanatory power of the models compared with a number of other sets of values (for these parameters). For the monthly model, a value of $d = 12$ provided a slightly better fit than any other value in the interval 1-12, while a value of $d = 1$ provided the best fit

Table 3.3: An EqCM of the ECU rate with non-linear oil price effects.

$$
\begin{aligned}
\widehat{Decu}_t = &\underset{(2.851)}{0.268} - \underset{(-2.436)}{0.056}[ecu - (cpi - cpi^f)]_{t-2} + \underset{(3.668)}{0.152}[R3\text{-}R3^f]_t \\
&\underset{(18.627)}{1.227} Decu_{t-1} - \underset{(-5.720)}{0.372} Decu_{t-2} - \underset{(-5.070)}{0.266} \Delta[R3_{t-2} - Dcpi_{t-1}] \\
&- \underset{(-2.385)}{0.649} Dcpi^f_{t-1} + \underset{(2.493)}{0.008} DFI.Y_t \\
&- \underset{(-6.428)}{0.055} i93p12 - \underset{(-5.782)}{0.027} i97p1 + \underset{(3.752)}{0.018} i97p11 + \underset{(6.617)}{0.032} i98p1 \\
&- \underset{(-6.125)}{0.042} i98p3 - \underset{(-3.627)}{0.055} i98p6 + \underset{(5.663)}{0.064} i98p9 - \underset{(-8.190)}{0.080} i98p11 \\
&- \underset{(-3.656)}{0.0175} oilp_{t-1} \\
&- \underset{(-3.575)}{0.2881} \Delta oilp_t \times [1 + \exp\{4(OILP_t - 14)\}]^{-1} \\
&- \underset{(-1.423)}{0.0223} \Delta oilp_t \times [1 + \exp\{-4(OILP_t - 20)\}]^{-1} \\
&- \underset{(-3.152)}{0.0428} \Delta oilp_t \times [1 + \exp\{(OILP_t - OILP_{t-12})\}]^{-1}
\end{aligned}
$$

Sample: 1990:11-1998:11, $T = 97$, $k = 20$. *Method: OLS*

Diagnostics

$\hat{\sigma}_{NL}$	=	0.455%
Log lik	=	*485.703*
LR-Test: $\chi^2(2)$	=	24.761[0.000]**
$AR\ 1-6\ F(6,71)$	=	0.704[0.647]
$ARCH(6)\ F(6,65)$	=	0.718[0.637]
$Het.\ Xi^2\ F(30,46)$	=	0.851[0.676]
Normality $\chi^2(2)$	=	0.281[0.869]
RESET F(1, 76)	=	0.239[0.626]
$TT_{NL}F(21,56)$	=	1.006[0.472]

Note: t-values (ordinary) in brackets (.) below the estimates and p-values in large brackets [.] beside the test statistics. TT_{NL} F(df1, df2) is defined in the main text. See Table 3.1 for details about the other tests.

on the quarterly data. Attempts were also made to estimate the threshold parameters c_1 and c_2 conditional on the proposed values of λ_1, λ_2, δ and d. Still the estimates on the monthly data turned out to be quite uncertain.

On the quarterly data, however, quite precise estimates of c_1 and c_2 at around 14 and 20 were obtained, especially when oil price terms that appeared redundant were excluded from the model. In particular, the level of oil price (*oilp*) had a relatively weak non-linear effect on the exchange rate. Also, the terms representing the oil price effects when *OILP* is above c_2 dollars had relatively weak effects. The estimate of c_1 was found to be

Table 3.4: An EqCM of the effective exchange rate with non-linear oil price effects.

$$\begin{aligned}
\Delta\widehat{\overline{e}}_t = & -\underset{(-3.608)}{0.100}\,[\overline{e}-(cpi-cpi^f)]_{t-1}+\underset{(2.678)}{0.190}\,\Delta\overline{e}_{t-1} \\
& -\underset{(-3.352)}{0.346}\,\Delta(RB_{t-1}-\Delta cpi_{t-2})-\underset{(-2.348)}{0.204}\,\Delta cpi^f{}_t \\
& +\underset{(3.050)}{0.821}\,\Delta^2 RB^f{}_{t-1}+\underset{(1.799)}{0.099}\,\Delta_4 FI.Y_t-\underset{(-4.211)}{0.032}\,id97q1 \\
& [-\underset{(-3.950)}{0.103}\,\Delta oilp_t-\underset{(-2.724)}{0.023}\,\Delta oilp_{t-2}]\times[1+\exp\{4(OILP_t-\underset{(37.711)}{14.197})\}]^{-1} \\
& -\underset{(-3.635)}{0.030}\,\Delta oilp_{t-1}\times[1+\exp\{OILP_t-OILP_{t-1}\})]^{-1}
\end{aligned}$$

Sample: 1972:2-1997:4, $T=103$, $k=10$. *Method: MLE*

Diagnostics

$\hat{\sigma}_{NL}$	=	1.00%
Log lik	=	418.852
R^2	=	0.56
$AR\ 1-5\ F(5,88)$	=	0.410[0.841]
$ARCH\ (4)\ F(4,85)$	=	0.478[0.752]
$Het.Xi^2\ F(20,72)$	=	0.495[0.960]
$Het.Xi\ Xj\ F(56,36)$	=	1.265[0.228]
$Normality\ \chi^2(2)$	=	1.707[0.426]
$RESET\ F(1,92)$	=	0.461[0.499]
$TT_{NL}\ F(22,64)$	=	0.889[0.665]

Note: Brackets below the estimates contain (ordinary) t-values while square brackets contain p-values. R^2 was obtained by OLS with c_1 fixed at 14.197. The value of R^2 is almost the same if a constant term is included in the model. Het. Xi_2 F(df1, df2) tests for heteroscedasticity by using cross products of regressors, see White (1980). TT_{NL} F(df1, df2) is defined in the main text. The other tests are explained in Table 3.1

fairly invariant to starting values when the oil price terms that appeared redundant were excluded. For example, it exhibited fast convergence to about 14.2, with a standard error estimate of around 0.37, when its starting value was varied in the range 12-18. A likely reason for the ability to derive relatively precise estimates of c_1 and c_2 at least, is that the quarterly data set contains more observations of $OILP$ around 14 and 20 USD than the monthly data set, see Figure 3.3.

Given the evidence from the bivariate analysis in Section 2 and the estimates of c_1 and c_2 on quarterly data, we proceeded with the modelling of $Decu_t$ on monthly data with c_1 and c_2 fixed at 14 and 20.

Table 3.3 and 3.4 present parsimonious versions of the two models and the associated diagnostics. In contrast to the model of $\Delta\overline{e}_t$, the model of $Decu_t$ is estimated by OLS

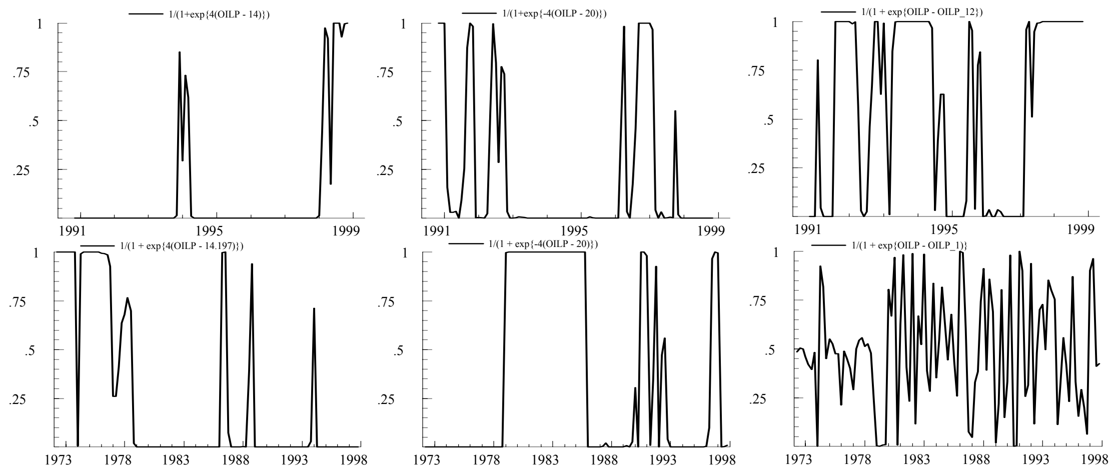


Figure 3.3: *First row: Values of the logistic functions over the period 1990:11-1998:11, monthly data. Second row: Values of the logistic functions over the period 1972:2-1997:4, quarterly data.*

since it is linear in parameters for fixed values of the transition and threshold parameters. A re-estimation by OLS of generalised versions of both models with L, H, D and p set to 2 and $f(,)$ defined by the reported values of λ_1, λ_2, δ, d, c_1 and c_2, also led to the same specific versions as shown in Table 3.3 and 3.4, when statistically insignificant variables and terms were excluded for the sake of parsimony.

Except for the non-linear oil price terms, both models contain the same regressors as the models with linear oil price effects presented in Table 3.1 and 3.2. Note also that the equilibrium correction term $[ecu$ - $(cpi\text{-}cpi^f)]_{t-2}$ in the monthly model has become significant at the 5% level when using the t-distribution, which is not inappropriate given the evidence for stationarity of this term, see Table 4.1 in the appendix. Now both models suggest that the Norwegian nominal exchange rate responds to differences in domestic and foreign prices in the short and long run.

However, the models have different long run solutions. The log level of oil price, $oilp_{t-1}$, appears in a linear way in the model of *Decu* (and even with the same coefficient estimate as in the linear model of *Decu),* see Table 3.1 and 3.3. The implication is that higher oil prices lead to an appreciation of the long run real exchange rate (ecu-(cpi-cpi^f)), in contrast to the PPP theory. On the quarterly data set, however, (log) level of oil prices were not found to have linear or non-linear effects on $\Delta\bar{e}_t$. For example, an F-test accepted zero restrictions on $oilp_t$, $oilp_{t-1}$ and $oilp_{t-2}$ with a p-value of about 90% when they were added to the model in Table 3.4. Thus the long run real exchange rate ($\bar{e}$-(cpi-cpi^f)) equals a constant, zero, in strict accordance with the PPP theory.

Short run oil price effects: Table 3.3 and 3.4 show that oil prices have statistically significant non-linear effects in the short run. The non-linear effects are essentially the same in both models suggesting that changes in oil prices have relatively strong effects when oil prices are below 14 USD and are displaying a falling trend. At levels around or above 20 dollars, oil prices were found to have statistically insignificant effects when judged at the 5% level. As noted above, such terms have been left out from the model of $\Delta\bar{e}$ altogether since their coefficients were also numerically small. The model of *Decu* however contains such a term, though it is only statistically significant at the 10% level. More observations at higher levels of oil prices may provide more firm evidence on its relevance.

Numerically, the effects of oil prices when below 14 dollars are almost 10 times the size of the effects suggested by the corresponding linear models. In a state of falling oil prices, the effects are also relatively stronger at levels of oil prices higher than 14 dollars. It is apparent that a linear representation of oil price effects tends to bring about underestimation of the exchange rate response to fluctuations in oil prices.

The non-linear oil price effects implied by both models are consistent with the level and trend effects suggested by the bivariate analysis in Section 2, apart from the statistical insignificance of oil price effects at oil prices around and above 20 dollars.

Figure 3.4 lays out the level and trend effects of oil prices on the exchange rate implied by the model of *Decu*. As noted above, both models imply essentially the same short run effects of oil prices. For the purpose of illustration, however, we focus on the effects suggested by the model of *Decu* since it also seems to capture the findings based on the bivariate analysis.

The graph at the top in Figure 3.4 illustrates how the strength of short run oil price effects depends on the level of the oil price in dollars, *OILP*. It sketches the total contem-

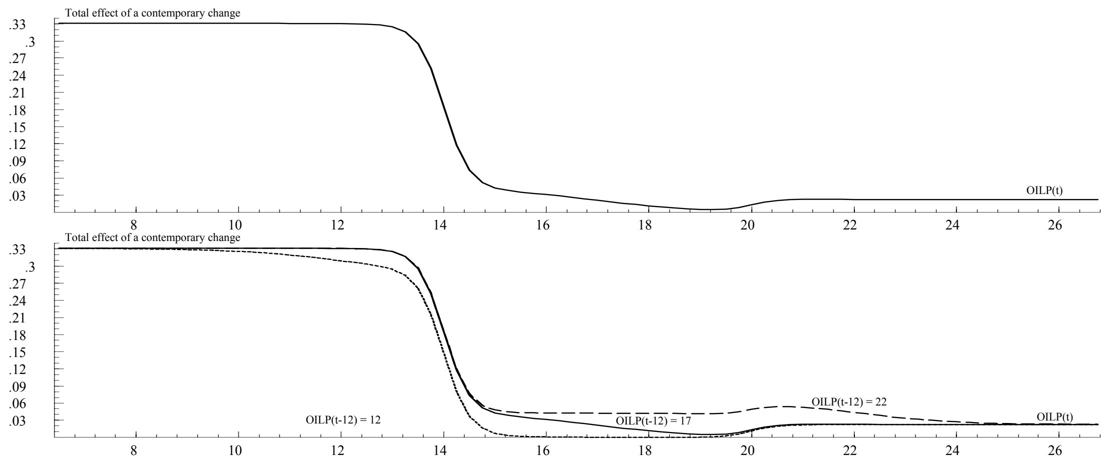


Figure 3.4: *Top: The total contemporary effect (=[*$0.288F_{Low}$*+*$0.022F_{High}$*+*$0.043F_{Diff}$*]* $\times$ $\Delta oilp_t$*), in absolute value, of a change in the oil price (vertical axis) at different levels of the current oil price (OILP(t)).* F_{Diff} *is defined by OILP(t-12) = 17. Bottom: The total contemporary effect of a fall in the oil price if the oil price was 12 (dotted line), 17 (solid line) and 22 (dashed line) dollars a years ago, i.e. at* $t-12$. *Horizontal axis: current level of oil price in dollars (OILP(t)).*

poraneous effect (in absolute value) on the exchange rate from $\Delta oilp_t$ at different levels of $OILP$ under the assumption that $OILP_{t-12}$ is 17 dollars. The relation between the effect on the exchange rate and the oil price in dollars appears convex. The effect is strongest when the oil price is below 13 dollars. This because $F_{t,Low}$ and $F_{t,Diff}$ are fully active since their values are close to 1 at such price levels, see Figure 3.2. As the oil price rises above 14 dollars, the partial effect declines sharply but thereafter at a slower pace as it rises above 15 dollars. The reason is that $F_{t,Low}$ converges to zero as the oil price rises above 15 dollars, see Figure 3.2.

The effect of the trend term, $\Delta oilp \times F_{Diff}$, also declines as the difference between the current price and past price (17 dollars) is reduced, and disappears altogether at 19 dollars, that is, when the oil price has risen by 2 dollars relative to the past, see Figure 3.2. In the range of 15-19 dollars, the contemporaneous effect on the exchange rate is entirely determined by the relative decline or rise in the oil price, cf. Figure 3.2. When the oil price rises above 19 dollars, the effect tends to increase because $F_{t,High}$ starts rising to 1. The effect at high levels of the oil price is however remarkably weak compared with the effect at low levels of the oil price.

The graph at the bottom in Figure 3.4 illustrates the trend effect more clearly. It shows that the strength of short run oil price effects also depends on whether the oil price is displaying a rising or falling trend. For instance, if the oil price was 22 dollars or higher a year ago, the effect of a decrease or increase, $\Delta oilp_t \gtrless 0$, is higher if the current price ($OILP_t$) is below its level of the previous year. The trend effect becomes negligible and vanishes when the current price exceeds past levels.

3.4. Model evaluation

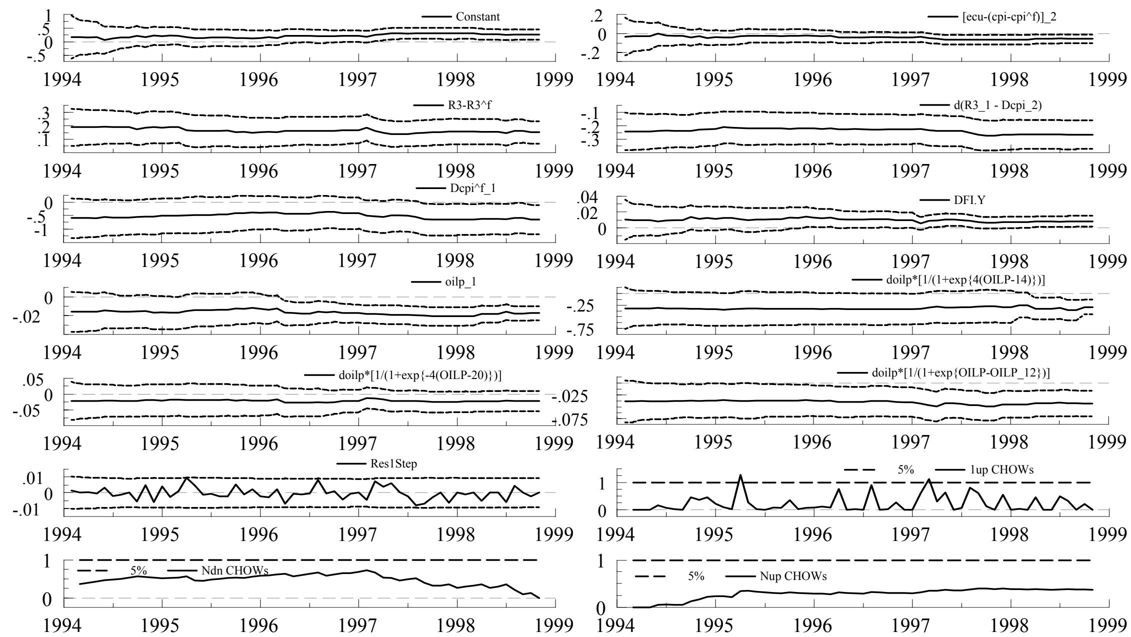


Figure 3.5: Model: *The EqCM of Decu with non-linear oil price effects. Recursive OLS estimates* $\pm 2SE$ *of the coefficients in the non-linear model. Thereafter: 1-Step ahead residuals* $\pm 2SE$*, 1-step ahead Chow tests, break point (Ndn) and forecast Chow tests scaled by their 5% critical values. The initial estimates are based on 40 observations, i.e. on the period 1990:11-1994:2. Prefix "d" denotes the 1. difference, while "D" denotes 12. difference (e.g. Dcpi =* cpi_t*-*cpi_{t-12}*).*

The diagnostics in Table 3.3 and 3.4 show that the regression specification test (RESET) does not reject the functional forms at standard levels of significance any more. The test for no neglected non-linear effects of a specified variable suggested by e.g. Teräsvirta

(1998) is also performed on the residuals from both models to asses whether non-linear effects of oil prices have been adequately characterised. The test is denoted as TT_{NL} and is performed using *OILP* as the transition variable. The tables shows that the null hypotheses of no-neglected non-linear oil price effects are not rejected at the standard levels of significance, in both cases. Also, the standard assumptions about the residuals are not rejected, as in the case of the linear models. The dummy variables are still needed however to ensure adherence to these assumptions in both models.

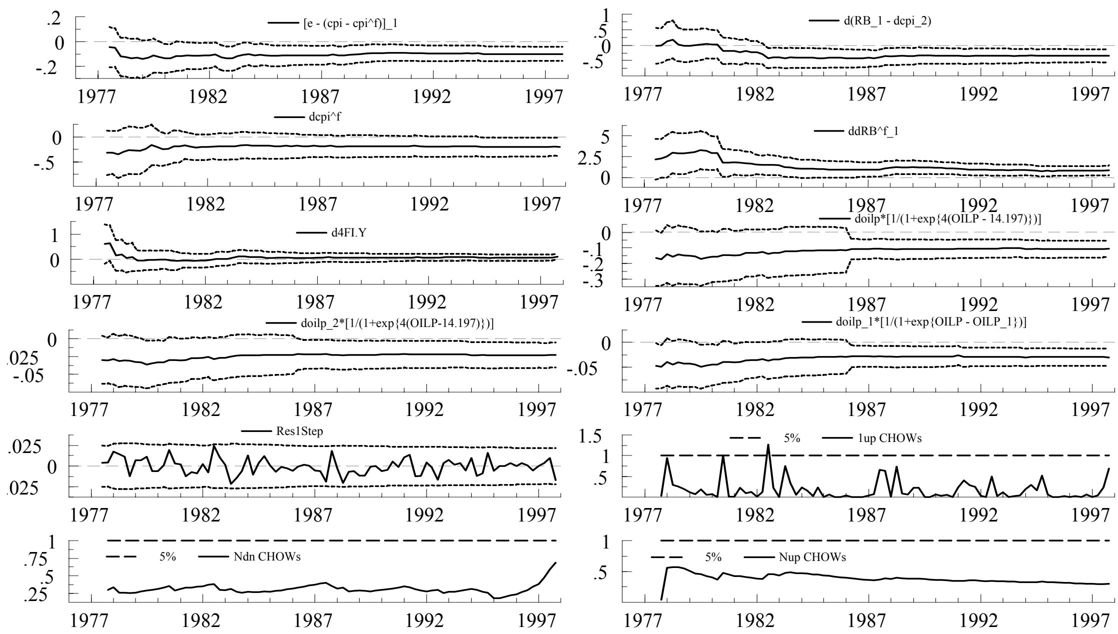


Figure 3.6: *Model: The EqCM of $\Delta\bar{e}$ with non-linear oil price effects. Constancy statistics for the EqCM with non-linear oil price effects. Initial estimation period is 1972:2-1976:4. These have been obtained by fixing c_1 at 14.197 and estimating the model recursively by OLS. Prefix "d" denotes the first difference Δ. The graphs show the recursive coefficient estimates $\pm 2SE_t$ of the indicated regressors, One-step ahead residuals $\pm 2SE_t$ and Chow statistics for the model. The latter are scaled by their critical values at the 5% level of significance.*

The tests for parameter stability over time do not indicate non-constancy in the parameters of both models at the 5% level, see Figure 3.5 and 3.6. In particular, the recursive OLS estimates of the non-linear oil price effects are remarkably stable over time. It is worth noticing that the standard deviations of the oil price effects when oil prices are below 14 dollars decrease relatively fast as more observations below this level come along, but without affecting the coefficient estimates of the oil prices, see e.g. Figure 3.5 in the beginning of 1998 and Figure 3.6 in 1985/1986.

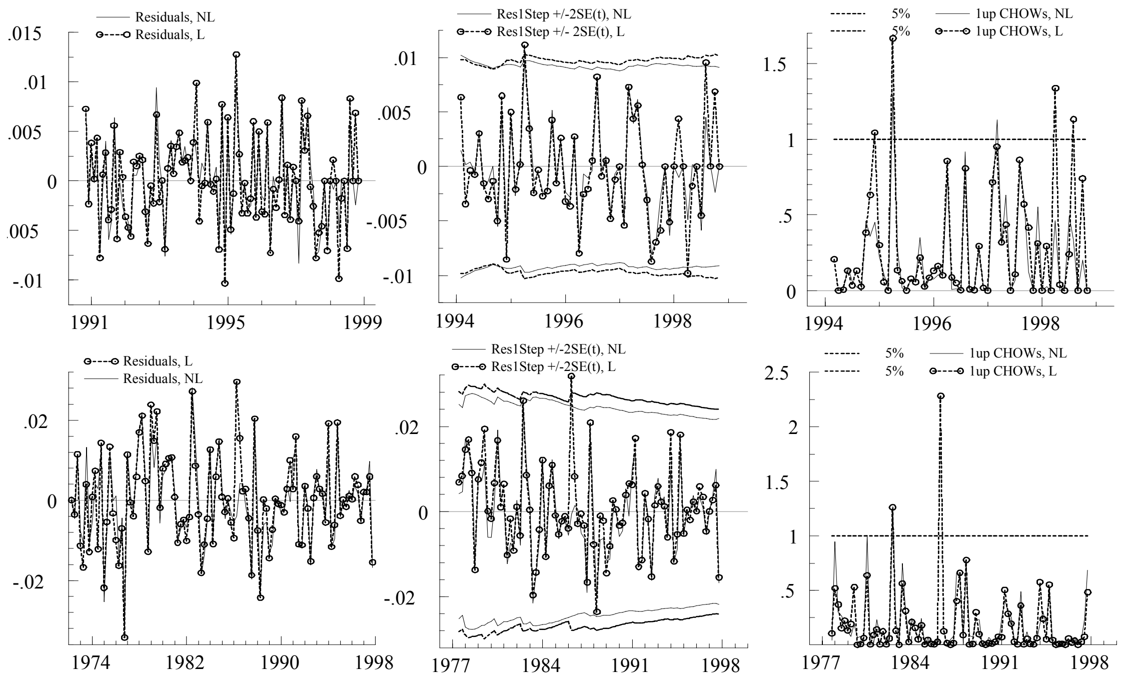


Figure 3.7: *Upper row: Residuals, 1-step ahead residuals* $\pm 2SE_t$, *1-Step ahead Chow statistics and the 5% critical values based on the linear (L) and non-linear (NL) EqCMs of Decu estimated on the monthly data set. Lower row: Residuals, 1-step ahead residuals* $\pm 2SE_t$, *1-Step ahead Chow statistics and the 5% critical values based on the linear (L) and non-linear (NL) EqCMs of* $\Delta\bar{e}$ *estimated on the quarterly data set. Everywhere, solid lines denote values based on the corresponding non-linear models denoted by NL.*

The non-linear models are clearly preferred over the linear models in terms of explanatory power. Table 3.1 and 3.2 show that the linear models have lower explanatory power than the non-linear models. In the case of the models of $Decu_t$, an LR test can be performed. Its outcome, reported in Table 3.3, indicates that the linear model of $Decu_t$ is strongly rejected against its non-linear version.

The higher explanatory power of the non-linear models relative to the linear models is clearly demonstrated in Figure 3.7. For the linear and non-linear models, it displays the residuals, 1-step ahead residuals ± 2 times recursively estimated standard errors ($\widehat{\sigma}$), SE_t, and 1-step ahead Chow test statistics, scaled by the (same) one off critical value at the 5% level. The non-linear oil price effects appear important in explaining the large fluctuations in the exchange rates which lead to lower standard errors of the residuals in the non-linear models than in the linear models. They also lead to more stable parameter estimates particularly in the face of oil price falls, as indicated by the Chow statistics.

For example, Figure 3.7 shows that there are fewer spikes in the residuals from the non-linear model of $Decu_t$ than in the residuals from its linear model, especially around 1994 and in 1998. Figure 3.3 suggests that the non-linear effects of low and declining oil prices are active in these periods and eliminate the spikes. However, Figure 3.7 indicates relatively smaller difference between the explanatory power of the two models in 1996/97, i.e. during the large appreciation pressure. Thus the appreciation pressure seems to have arisen mainly due to other factors than the relatively high oil prices in this period, as suggested by e.g. Kvilekval and Vårdal (1997).

The lower row of Figure 3.7 gives the impression that the non-linear model of $\overline{e}$ mainly owes its superior performance to the coincidence between low oil prices and the devaluation in 1986. There is thus a need to investigate whether low oil prices can explain other incidents of fall in the value of the krone. We turn to this issue later.

Extension of the information set with additional variables: Table 3.5 demonstrates that the model of $\Delta\overline{e}$ with non-linear oil price effects is quite robust to extensions of the information set. The information set has been extended by vector $\mathbf{Z}$; specifically, by productivity growth at home and abroad, Δq and Δq^f, growth in government expenditures (Δg) and growth in domestic GDP (Δy) or changes in unemployment rate Δu. The additional variables were included with up to three lags. Table 3.5 presents the model and the outcome of variable omission tests when Δq, Δq^f and Δg are included with up to three lags.

None of these variables turns out to have statistically significant effects at the 5% level. The fiscal policy variables have relatively large coefficients that indicate appreciation pressure on the nominal exchange rate when government spendings increase. The effects are however insignificant at the 5% level, even if other clearly insignificant terms are omitted from the model. The variable omission tests indicate that none of the added

Table 3.5: The EqCM of the effective exchange rate with additional variables.

$$
\begin{aligned}
\Delta\widehat{\overline{e}}_t = &- \underset{(-3.223)}{0.097}\, [\overline{e} - (cpi - cpi^f)]_{t-1} + \underset{(2.606)}{0.205}\, \Delta\overline{e}_{t-1} \\
&- \underset{(-2.232)}{0.279}\, \Delta(RB_{t-1} - \Delta cpi_{t-2}) - \underset{(-1.966)}{0.257}\, \Delta cpi^f{}_t \\
&+ \underset{(2.538)}{0.765}\, \Delta^2 RB^f{}_{t-1} + \underset{(1.428)}{0.086}\, \Delta_4 FI.Y_t - \underset{(-3.575)}{0.029}\, id97q1 \\
&[- \underset{(-3.564)}{0.103}\, \Delta oilp_t - \underset{(-2.347)}{0.023}\, \Delta oilp_{t-2}] \times [1 + \exp\{4(OILP_t - 14.197)\}]^{-1} \\
&- \underset{(-3.758)}{0.034}\, \Delta oilp_{t-1} \times [1 + \exp\{OILP_t - OILP_{t-1}\})]^{-1} \\
&- \underset{(-1.702)}{0.059}\, \Delta q^f_t - \underset{(-0.765)}{0.026}\, \Delta q^f_{t-1} + \underset{(0.270)}{0.009}\, \Delta q^f_{t-2} + \underset{(0.342)}{0.011}\, \Delta q^f_{t-3} \\
&- \underset{(-0.739)}{0.022}\, \Delta q_t - \underset{(-0.519)}{0.016}\, \Delta q_{t-1} + \underset{(0.409)}{0.013}\, \Delta q_{t-2} - \underset{(-0.495)}{0.015}\, \Delta q_{t-3} \\
&- \underset{(-1.316)}{0.235}\, \Delta g_t - \underset{(-0.544)}{0.111}\, \Delta g_{t-1} + \underset{(0.209)}{0.043}\, \Delta g_{t-2} - \underset{(-1.850)}{0.353}\, \Delta g_{t-3}
\end{aligned}
$$

Sample: 1972:2-1997:4, $T = 103$, $k = 22$, *Method: OLS*

Tests for omission of variables

			Δq^f	Δq	Δg	Δy	Δu
Joint[a]	$F(4,\ 81)$	:	[0.484]	[0.611]	[0.111]		
Separate[b]	$F(4,\ 89)$	:	[0.349]	[0.773]	[0.139]	[0.311]	[0.805]

Note: Brackets below the estimates contain (ordinary) t-values while square brackets contain p-values. Model estimated by OLS with c_1 fixed at 14.197. [a]The F-tests were performed by placing zero restrictions on the contemporary and lagged values of the indicated variable while retaining the additional variables. [b]The F-test were performed by excluding all additional variables except the contemporary and lagged values of the indicated variable.

variables have significant effects and can be omitted from the model altogether. This was also the case when the insignificant variables were sequentially omitted from the model, with one exception. The contemporaneous effect of productivity changes in the trading countries (Δq^f_t) was found to be significant at the strictly 5% level when all the other additional variables were excluded from the model, but with the "wrong" sign, i.e. negative. Government expenditures on consumption and real investment were also entered separately, i.e. as $\Delta(C_G/Y)$ and $\Delta(J_G/Y)$ with up to three lags, but neither of them were found to have significant effects.

Growth in GDP (Δy) was also included in the model with up to three lags, both when the additional variables were present and when they were left out, but no significant effects were found and there no noteworthy change was detected in the coefficient estimates of the initial variables. Changes in GDP, as well in the rate of unemployment (Δu), are likely to reflect the influence of different variables that affect the Norwegian economy. Including Δy, or Δu, in the model may to some extent, mop up the effects of variables which are not explicitly included in the model. The variable omission tests, however, show that neither Δy nor Δu has statistically significant effects on the nominal exchange rate.

Similar results were obtained in the case of the monthly model with non-linear oil price effects. For example, an *F-test* accepted zero restrictions on the associated $\mathbf{Z}$ vector when included with two lags in the model. The $F(9,\ 68)$ statistics was 0.999 with a p-value of 0.45. As noted earlier, vector $\mathbf{Z}$ vector includes $Fisc.Y$, Δu and Δu^f in the case of the monthly model.

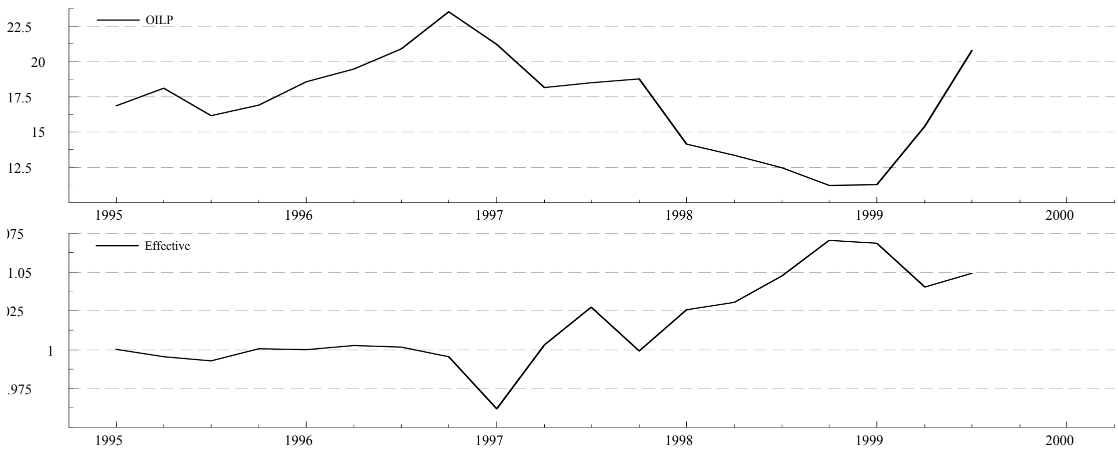


Figure 3.8: *Oilprice in US dollars (OILP) and the effective nominal exchange rate* $(\overline{E})$ *over the period 1995:1-1999:3.*

Post sample evaluation In the following, we undertake a post-sample evaluation of the model of $\Delta\overline{e}$. We are able to extend the quarterly data set by 7 quarters for all the variables in the model except $FI.Y$ for which we lack the data for the year 1999. $\Delta_4 FI.Y$ is anyway statistically insignificant at the 5% level so we do not expect that it will induce considerable omitted variable bias in other parameter estimates if the model is estimated without it.

The new observations cover the period 1998:1-1999:3 in which oil prices have fluctuated

Table 3.6: The EqCM of the effective exchange rate estimated on data extended by 7 quarters.

$$
\begin{aligned}
\Delta\widehat{\overline{e}}_t = & -\underset{(-3.988)}{0.110}\,[\overline{e} - (cpi - cpi^f)]_{t-1} + \underset{(2.053)}{0.149}\,\Delta\overline{e}_{t-1} \\
& -\underset{(-2.942)}{0.318}\,\Delta(RB_{t-1} - \Delta cpi_{t-2}) - \underset{(-2.900)}{0.277}\,\Delta cpi^f{}_t \\
& +\underset{(2.887)}{0.818}\,\Delta^2 RB^f{}_{t-1} + 0\Delta_4 FI.Y_t - \underset{(-3.765)}{0.031}\,id97q1 \\
& [-\underset{(-5.243)}{0.123}\,\Delta oilp_t - \underset{(-2.959)}{0.025}\,\Delta oilp_{t-2}] \times [1 + \exp\{4(OILP_t - 14.197)\}]^{-1} \\
& -\underset{(-3.577)}{0.031}\,\Delta oilp_{t-1} \times [1 + \exp\{OILP_t - OILP_{t-1}\})]^{-1}
\end{aligned}
$$

Sample: 1972:2-1999:3, $T = 110$, $k = 9$. *Method: OLS*

Diagnostics

$\hat{\sigma}$	=	0.011
R^2	=	0.54
DW	=	1.99
$AR\ 1-5\ F(5,96)$	=	0.376[0.864]
$ARCH\ (4)\ F(4,93)$	=	0.198[0.939]
$Het.Xi^2\ F(18,82)$	=	0.580[0.904]
$Het.Xi\ Xj\ F(46,54)$	=	0.697[0.894]
$Normality\ \chi^2(2)$	=	1.376[0.503]
$RESET\ F(1,100)$	=	0.289[0.592]

Note: Brackets below the estimates contain (ordinary) t-values while square brackets contain p-values. c_1 fixed at 14.197 while Δ_4FI.Y$_t$ was left out from the model. The value of R^2 is almost the same if a constant term is included in the model. The tests are explained in Table 3.1.

across a relatively wide range of levels and displayed both a downward and upward trend, see Figure 3.8. The figure also shows a relatively large depreciation in the value of the nominal effective exchange rate ($\overline{E}$) during 1998 followed by a marked appreciation in 1999. The new observations are therefore quite informative for the purpose of assessing the stability of the model, and of the non-linear oil price effects, in particular.

Table 3.6 presents the model of $\Delta\overline{e}$, reestimated by OLS on the extended data set with c_1 fixed at 14.197, and the associated diagnostics. The table shows that the model is intact; particularly the oil price effects, whose statistical significance has increased. Beside that, the t-value of the equilibrium correction term has become almost -4 which lends strength to the hypothesis of purchasing power parity in the long run. Furthermore, the model diagnostics have improved.

The model with linear oil price effects in Table 3.2 was also reestimated on the extended data set. However, the oil price effects remained almost as in Table 3.2. Moreover, zero restrictions on $\Delta oilp_t$, $\Delta oilp_{t-1}$ and $\Delta oilp_{t-2}$ were accepted by an F-test with a p-value of 6%. In addition, the RESET test still suggested a functional form misspecification at the 5% level.

A model with deterministic account of non-linearity: Figure 3.7 seemed to suggest that the higher explanatory power of the non-linear model of $\overline{e}$ stems from a proper representation of the devaluation in 1986, which coincided with the oil price fall in 1985/1986. One may suspect that a model with linear oil price effects but with dummy variables for the abrupt changes in the exchange rate can lead to a comparable improvement in the linear model and make the non-linear representation of oil price effects redundant. Moreover, oil prices may not even have linear effects on the exchange rate once the devaluation is controlled for by dummy variables, cf. Bjørvik et al. (1998).

A model with dummy variables would be preferable if the co-movement in the oil price and the exchange rate in 1985/86 was a unique event and not a stable feature; since otherwise, a model with non-linear oil price effects is likely to over-predict the exchange rate depreciation in a state of low and falling oil prices. In the following we test for: (a) whether oil prices have significant effects once the devaluation is controlled for, and in particular, whether use of dummy variables can make the non-linear representation of oil price effects redundant, and (b) whether the incident of low oil prices and a fall in the value of the currency in 1986 was just a unique event and unlikely to be repeated in the future.

The upper panel of Table 3.7 presents a model with dummy variables to account for the events in 1986, which seems to be the main cause of the observed non-linearity. Two impulse dummies that takes on a value of 1 in 1986:2 and in 1986:3, respectively, appeared sufficient to capture the fall in the value of the exchange rate, see Figure 3.9. Consequently, there is an increase in the explanatory power of the model relative to the linear model in Table 3.2, measured by R^2 and $\widehat{\sigma}$, and the RESET test does not reject the functional form at the 5% level.

Table 3.7: Models of the effective exchange rate with dummy variables to represent the non-linearity.

$$
\begin{aligned}
\Delta\widehat{\bar{e}}_t = & -\underset{(-3.420)}{0.107}\,[\bar{e}-(cpi-cpi^f)]_{t-1} + \underset{(2.802)}{0.233}\,\Delta\bar{e}_{t-1} \\
& -\underset{(-2.949)}{0.340}\,\Delta(RB_{t-1}-\Delta cpi_{t-2}) - \underset{(-2.488)}{0.258}\,\Delta cpi^f{}_t \\
& +\underset{(2.637)}{0.798}\,\Delta^2 RB^f{}_{t-1} + \underset{(1.432)}{0.091}\,\Delta_4 FI.Y_t - \underset{(-4.115)}{0.035}\,id97q1 \\
& -\underset{(-0.228)}{0.002}\,\Delta oilp_t - \underset{(-1.239)}{0.009}\,\Delta oilp_{t-1} + \underset{(0.592)}{0.004}\,\Delta oilp_{t-2} \\
& +\underset{(2.975)}{0.038}\,id86q2 + \underset{(1.684)}{0.021}\,id86q3
\end{aligned}
$$

Sample: 1972:2-1997:4, $T = 103$, $k = 12$. *Method: OLS*
$R^2 = 0.52$, $\hat{\sigma} = 0.0115$, $RESET\ F(1, 90) = 1.608$ [0.208]

$$
\begin{aligned}
\Delta\widehat{\bar{e}}_t = & -\underset{(-3.498)}{0.105}\,[\bar{e}-(cpi-cpi^f)]_{t-1} + \underset{(2.623)}{0.213}\,\Delta\bar{e}_{t-1} \\
& -\underset{(-3.261)}{0.358}\,\Delta(RB_{t-1}-\Delta cpi_{t-2}) - \underset{(-2.269)}{0.227}\,\Delta cpi^f{}_t \\
& +\underset{(3.301)}{0.974}\,\Delta^2 RB^f{}_{t-1} + \underset{(1.782)}{0.108}\,\Delta_4 FI.Y_t - \underset{(-3.822)}{0.031}\,id97q1 \\
& -\underset{(-0.091)}{0.001}\,\Delta oilp_t + \underset{(0.237)}{0.003}\,\Delta oilp_{t-1} + \underset{(1.971)}{0.020}\,\Delta oilp_{t-2} \\
& -\underset{(-0.536)}{0.012}\,id86q2 + \underset{(0.118)}{0.002}\,id86q3 \\
& [-\underset{(-2.478)}{0.122}\,\Delta oilp_t - \underset{(-3.050)}{0.044}\,\Delta oilp_{t-2}] \times [1+\exp\{4(OILP_t - 14.197)\}]^{-1} \\
& -\underset{(-1.981)}{0.037}\,\Delta oilp_{t-1} \times [1+\exp\{OILP_t - OILP_{t-1}\})]^{-1}
\end{aligned}
$$

Sample: 1972:2-1997:4, $T = 103$, $k = 15$. *Method: OLS*
$R^2 = 0.58$, $\hat{\sigma} = 0.0109$, $RESET\ F(1, 87) = 0.453$ [0.503]

Note: Brackets below the estimates contain (ordinary) t-values while square brackets contain p-values. Both models were estimated by OLS with c_1 fixed at 14.197. The value of R^2 is almost the same if a constant term is included.

In this model, the oil price effects are even weaker than in the linear model, numerically and statistically. Hence one could omit them altogether from the model for the sake of parsimony, and claim that the common perception of significant oil price effects on the Norwegian exchange rate just owes to the coincidence of devaluation and low oil prices in 1986. However, this would be a fallacy as shown below.

The lower panel of the table adds the non-linear oil price terms to the above model, with c_1 fixed at 14.197. The inclusion demonstrates that linear oil price terms and the oil price-

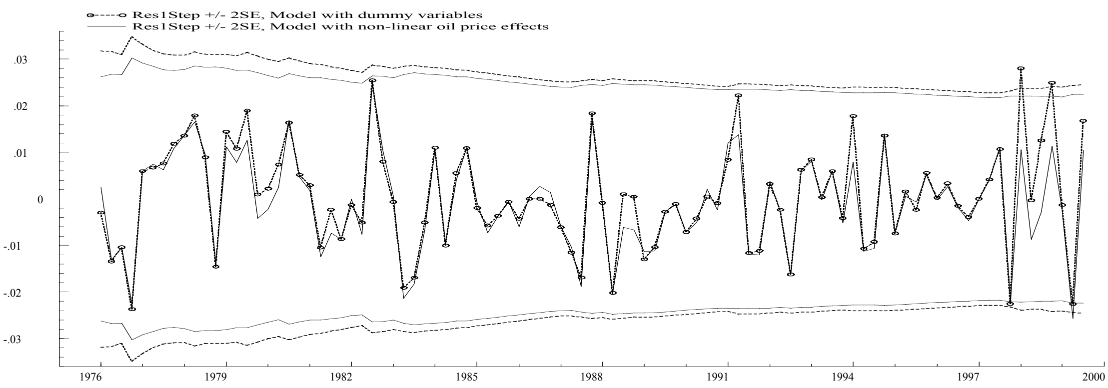


Figure 3.9: *Dashed lines:1-step ahead residuals* $\pm 2SE$ *of the model with dummy variables in the upper panel of Table 3.7, but recursively estimated on data over the period 1972:2-1999:3. Solid lines: depict 1-step ahead residuals* $\pm 2SE$ *based on the model with non-linear oil price effects of Table 3.7, but also recursively estimated on data covering the period 1972:2-1999:3. Initial estimation period 1972:2-1975:4, in both cases.*

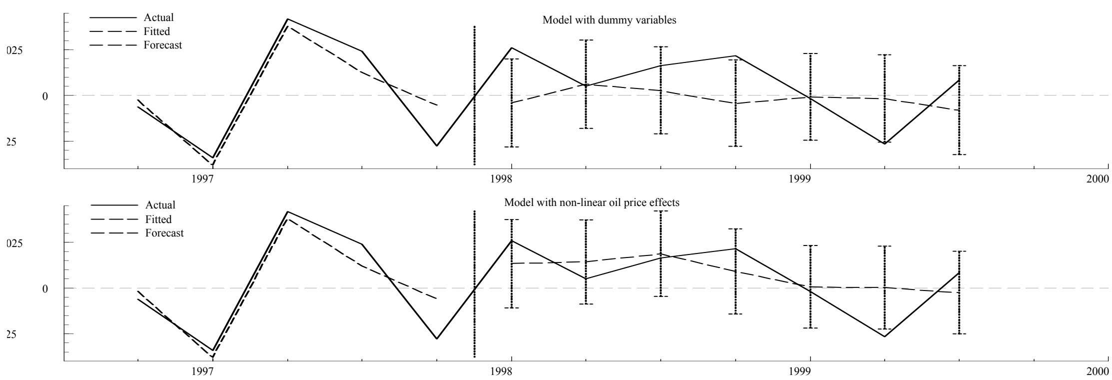


Figure 3.10: *Top: Actual, fitted and forecasted values (1-step ahead) of* $\Delta\bar{e}$ *based on the model with dummy variables, cf. the upper panel of Table 3.7. The dotted bars are 95% prediction intervals. 1-step ahead forecasts are for the period 1998:1-1999:3. Bottom: As above but based on the model with the non-linear oil price effects, Table 3.4.*

fall dummy variables become redundant while non-linear oil price terms are significant at the 5% level.

This exercise supplements the evidence in favour of the model with non-linear oil price effects, which also outperforms the model with the dummy variables in terms of explanatory power. The latter property is obvious in Figure 3.9, which shows that the estimated standard errors of the residuals of the model with non-linear oil price effects is lower than that of the model with the dummy variables, not only in the sample used to derive the models but also when the sample is extended by 7 quarters to 1999:3.

Figure 3.10 shows that the model with non-linear oil price effects is able to account for falls in the value of the krone beyond the devaluation in 1986. The figure depicts the 1-step ahead forecasts for $\Delta\bar{e}$ over the period 1998:1- 1999:3 based on the model with dummy variables and the models with non-linear oil price effects, as presented in the upper panel of Table 3.7 and Table 3.4. It appears that the model with dummy variables significantly under-predicts the depreciation of the exchange rate during 1998 while the model with non-linear oil price effects does not. Both models under-predict the appreciation of the exchange rate in 1999:2, however, which coincides with rising oil prices, see Figure 3.10. This points to oil price effects in the direction of those suggested by the bivariate analysis and the monthly model with non-linear oil price effects. As noted earlier, such effect were found statistically insignificant on the quarterly data, possibly due to the scarcity of such co-movements in oil prices and the exchange rate in the quarterly data sample ending in 1997:4. However, it remains to be investigated whether $\Delta_4 FI.Y$ can account for the poor forecasts for 1999:2.

Nevertheless, Figure 3.10 suggests that a fall in the value of the exchange rate in the face of low and falling oil prices is not a coincidental feature of the model. Hence, a model with dummy variables for the in-sample falls in the value of the exchange rate is likely to under-predict the exchange rate response in states of low and falling oil prices.

4. Conclusions

This essay has investigated whether imposition of a linear relation between oil prices and the Norwegian exchange rate leads to an underestimation of oil price effects and hence a failure to explain major changes in the exchange rate in the face of large fluctuations in oil prices. This is believed to be a feature of earlier work on this subject.

The essay has utilised samples of daily, monthly and quarterly observations of different lengths and a variety of techniques and models to investigate whether oil prices have non-linear or state dependent effects on the value of the Norwegian exchange rate. In particular, it has derived data consistent and interpretable equilibrium correcting models of both the krone/ECU exchange rate and the nominal effective exchange rate. Moreover, it has undertaken an extensive evaluation of the derived models to demonstrate the robustness of the obtained results, which have appeared fairly unanimous across the data samples and the different models. These results are:

There is a negative relation between the oil price and the Norwegian exchange rate: a rise in oil prices tends to raise the value of the krone while a fall tends to reduce the value of the krone.

The negative relation is however non-linear since the strength of this relation varies with the level and the trend in oil prices. A change in oil prices has a stronger impact on the exchange rate when the level of the oil price is below 14 US dollars than at higher levels. It also appears that the impact of a change in oil prices tends to increase at levels of oil prices around and above 20 US dollars, but this result has been found to be statistically insignificant at the 5% level. A change in the oil price has numerically and statistically insignificant effects when oil prices fluctuate within their normal range, which has appeared to be the levels between 14 and 20 US dollars in the data samples at hand. The effect of a change in oil prices on the exchange rate has also been found to depend on whether oil prices display a falling or rising trend. In the former case, the effect is quite strong while a change in oil prices does not have any significant effect on the exchange rate if oil prices are on a rising trend. Accordingly, changes in oil prices have negligible effects, if any, on the exchange rate if oil prices are at normal levels, unless they exhibit a falling tendency.

The effects of a change in oil prices on the exchange rate has been found to be much stronger in models with the non-linear representation of oil price effects than in models with a linear representation of oil price effects. For instance, at low levels of oil prices, the effect of a change in oil prices is about 10 times stronger than in the models with linear oil price effects. There is ample evidence to suggest that imposition of linear oil price effects tends to bring about a gross underestimation of the exchange rate response to a change in oil prices, especially, when oil prices are at low levels and are falling.

The models with non-linear oil price effects outperform the models with linear oil price effects in terms of explanatory power, especially during the major falls in the value of the krone. Moreover, the former models appear to be data consistent in contrast to the latter models, which are found to be misspecified. Furthermore, a model with the non-linear representation of oil price effects successfully predicts the out-of-sample depreciation of the Norwegian exchange rate in 1998. This is in contrast to a similar model but with dummy variables to account for in-sample excessive fluctuations in the exchange rate. Thus, the observed non-linearity seems to reflect a stable feature of the underlying relation between the oil price and the exchange rate rather than a coincidental or an unique in-sample event.

The reported non-linear oil price effects are only significant in the short run. In the long run, oil prices are found to have linear effects on the krone/ECU exchange rate but no effects on the nominal effective exchange rate. The model of the krone/ECU exchange rate, which is based on monthly data from the 1990s, implies that the Norwegian real exchange rate depends on the oil price. Accordingly, high levels of oil prices lead to a real exchange rate appreciation. However, the model of the nominal effective exchange rate, based on quarterly data over the period 1972-1997, implies that the nominal exchange rate only reflects the ratio between domestic and foreign prices, in strict accordance with the purchasing power parity (PPP) hypothesis. Accordingly, the Norwegian real exchange rate is constant and independent of oil prices in the long run. Thus the empirical evidence in this essay provides mixed support for the PPP hypothesis. However, given the long span of data used in deriving the model of the nominal effective exchange rate, we are inclined to attach more weight to the implications of this model than to those of the model of the krone/ECU exchange rate.

The absence of non-linear oil price effects in the long run and the strong support for the non-linear oil price effects in the short run is consistent with the view that non-linearities may arise due to institutional behaviour. That is, the findings are consistent with the view that a central bank is often less keen on stabilising the exchange rate when faced with depreciation pressure compared with appreciation pressure, especially when the exchange rate target is in conflict with other concerns such as unemployment, financial stability and competitiveness. Such an asymmetric response can itself lead to stronger pressure for depreciation than for appreciation and thereby make the depreciation self-fulfilling.

The empirical evidence in this essay indicates that a change in the oil price is likely to trigger larger capital movements and hence larger exchange rate fluctuations when the oil prices are below their normal range and are falling. Low and falling oil prices, due to their adverse effects on e.g. the activity level, may intensify the tension between a stable exchange rate policy and other objectives pursued by the monetary authorities and thus, increase the uncertainty associated with their commitment to the exchange rate policy.

Finally, we note that this essay has presented well specified multivariate exchange rate models with remarkably stable parameter estimates and fairly high explanatory power, which is encouraging. This is against the background of widespread pessimism in the literature regarding the possibility of deriving such exchange rate models, with or without non-linear effects of macroeconomic variables, see e.g. Meese and Rose (1991), Meese (1990) and Frankel and Rose (1995). Therefore, the essay not only suggests that one takes a new look at studies that have reported unstable oil price effects on exchange rates, but also offers results that can be utilised in further theoretical and empirical research on exchange rates.

Appendix A: Monthly data

Small letters indicate the natural logarithm of the variables listed below, e.g. $x = ln(X)$. Unless otherwise stated, the source is Norges Bank's database TROLL8.

The variables and some of their transformations are graphed in Figure 4.1 below. The unit root tests are reported in Table 4.1.

CPI: Consumer price index. Series no. M1600012.

CPI^f: Consumer price in Norway's trading partners. Series no. M4690512.

ECU: Index for NOK/ECU. Central value: 7.9940 =100. Series no. M9302432 for monthly and D9302432 for daily observations.

Fisc.Y: Public sector's net savings relative to the gross national product (Y).

i93p12: Impulse dummy that takes on a value 1 in 1993:12 and zero elsewhere. The other impulse dummies are constructed in a similar way.

Dcpi : Consumer price inflation in Norway. $Dcpi_t = cpi_t - cpi_{t-12}$

$Dcpi^f$: Consumer price inflation in Norway's trading partners, mainly western European countries. Calculated as *Dcpi*.

Decu: $= ecu_t$ - ecu_{t-12}.

$\Delta R3$: $= R3^f_t$ - $R3^f_{t-1}$.

$\Delta R3N$: $= R3_t$ - $R3_{t-1}$.

FI.Y: Foreign net financial assets in Norway relative to GDP. FI.Y = *(B70NOSS-BNO70SS)/Y*. Source: database TROLLBAL, Norges Bank.

OILP: Spot price of Brent Blend crude oil in US dollars. Series no. M2001712 for monthly and D2001712 for daily observations.

$R3^f$: Effective per annum eurorent on ECU denominated assets, maturity three month. Series no. M865135C.

$R3$: Effective per annum eurorent on NOK denominated assets, maturity three month. Series no. M901605C.

U: Registered (or open) unemployment rate, seasonally unadjusted. Series no. M0102322.

U^f : Registered unemployment in EU countries, seasonally adjusted. Series no. M4746023.

Y: Gross national product of Norway. The monthly series is calculated by even distribution of the quarterly series on each of the month (in the quarter). Source. Quarterly National Accounts, Statistics Norway.

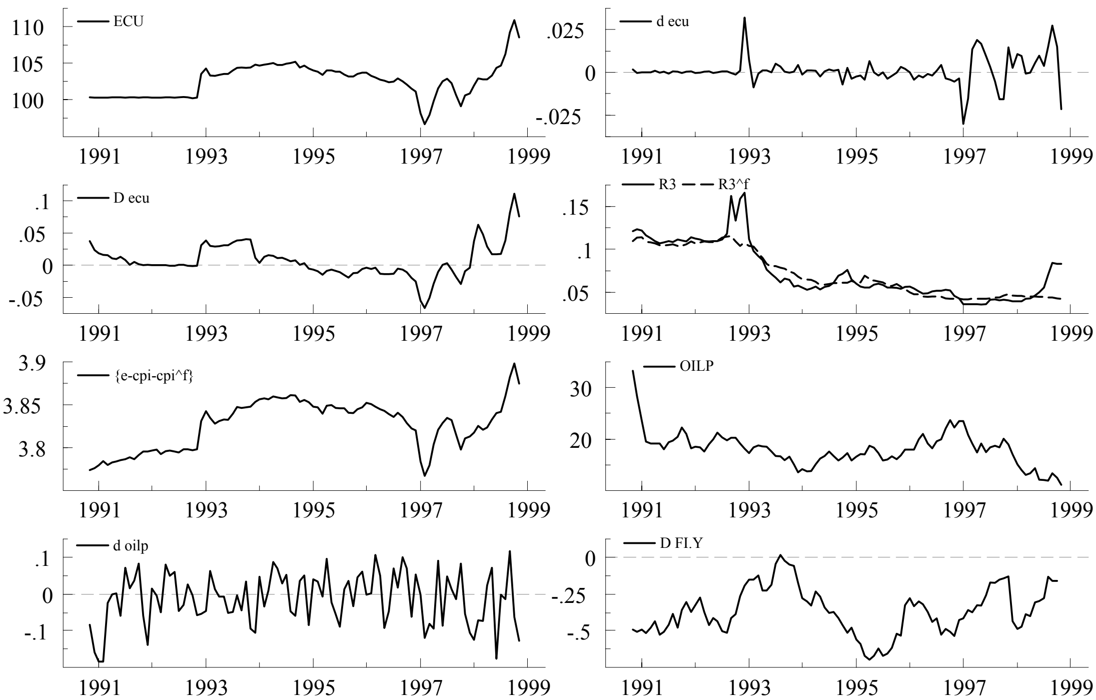


Figure 4.1: *Monthly observations of selected variables over the period 1990:11-98:11.* $d = \Delta$ while $D = \Delta_{12}$.

Table 4.1: Unit root tests, monthly observations 1989:7-1998:11.

Variable	d	$\widehat{\varrho}$	t-$adf(d)$
$OILP$	1	-0.160	-4.263**
$oilp$	1	-0.123	-3.083*
$R3$	3	-0.043	-1.808
$R3^f$	1	-0.003	-0.350
$\Delta R3$	1	-0.797	-4.709**
$\Delta R3^f$	2	-0.536	-4.548**
$[R3\text{-}R3^f]$	3	-0.247	-3.595**
$[cpi\ \text{-}\ cpi^f]$	2	-0.046	-3.603**
ecu	2	-0.059	-2.364
ecu^D	2	-0.061	-2.964*
$[ecu\text{-}(cpi\text{-}cpi^f)]$	5	-0.050	-2.873*
$[ecu\text{-}(cpi\text{-}cpi^f)]^D$	5	-0.053	-3.749**
$DFI.Y$	5	-0.118	-3.796**

The following model is estimated to test the H_0: $\rho = 0$: $\Delta y_t = \mu + \varrho y_{t-1} + \sum_{i=0}^{d} \phi_i \Delta y_{t-1-i}$ $+\delta D + \varepsilon_t$,. The augmented Dickey-Fuller model: $\delta = 0$. * and ** indicate significance at 5% and are 1%, respectively. The respective critical values -2.863 and -3.435. ^{D} Indicates that an impulse dummy has been used to adjust for the break in the series in December 1992.

Appendix B: Quarterly data

Unless stated otherwise, the variables listed below are taken from the data base for RIMINI, the quarterly macroeconometric model used in Norges Bank. The main sources for RIMINI's data base are Quarterly National Accounts, FINDATR, TROLL8, OECD_MEI and IFS. These data bases are maintained by Statistics Norway, Norges Bank, Norges Bank, OECD and IMF, respectively. The RIMINI names of the variables are indicated in square brackets []. Note that this essay employs seasonally unadjusted quarterly data. Some of the variables are graphed in Figure 4.2 while the unit root tests are reported in Table 4.2.

CPI : Consumer price index for Norway, 1991 = 1. [CPI].

CPI^f: Trade weighted average of consumer price indices for Norway's trading partners. Measured in foreign currency, 1991 = 1. [PCKONK].

C_G : Public consumption expenditures, fixed 1995 prices, Mill. Norwegian krone (NOK). [CO].

CS : Centered seasonal dummy variable (mean zero) for the first quarter in each year. It is 0.75 in the first quarter and -0.25 in each of the three other quarters, for every year.

E : Trade weighted nominal value of NOK, 1991 = 1. [PBVAL].

$FI.Y$: A measure of foreigners (all), net financial investment in Norway, fixed 1991 prices, Mill. NOK. Constructed by taking the first difference of net foreign debts share of GDP, [LZ.Y], i.e. $FI.Y = \Delta$LZ.Y.

g : Public expenditures' share of GDP, i.e. $g = (C_G + J_G)/Y$.

id97q1 : Impulse dummy related to the oil price hike in 1996/97. It has a value of 1 in 1997:1, -1 in 1997:2 and zero elsewhere.

J_G : Public expenditures for gross real investment, fixed 1995 prices, Mill. NOK. [JO].

OILP : Price per barrel of Brent Blend crude oil in US dollars. Source TROLL8, series no. Q2001712.

Q : Value added per unit labour cost in Norway. The inverse of value added based unit labour costs. 1/[LPE.Y].

Q^f: Value added per unit labour cost in trading partners. The inverse of trade weighted average of value added based unit labour costs. 1/[M.LPE].

R : Trade weighted real exchange rate, defined as R = $(E \text{ x } CPI^f)/CPI$. [RPBVAL].

RB : Yield on 6 years Norwegian government bonds, quarterly average. [R.BS].

RB^f : NOK basket-weighted average of interest rates on long term foreign bonds. [R.BKUR].

RS : *3* month Euro krone interest rate. [RS].

U : Total unemployment rate, fraction of labour force exclusive self employed and those on labour market programs. [UTOT].

Y : Gross domestic product for Norway. Mill. NOK, fixed 1995 prices. [Y].

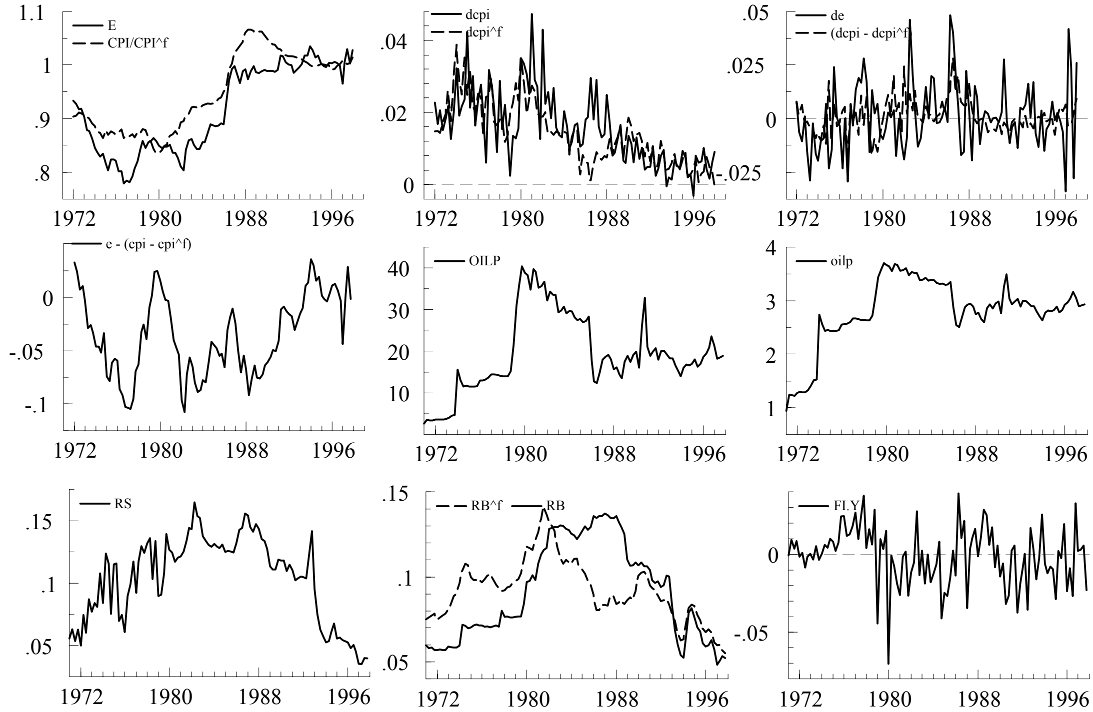


Figure 4.2: *Quarterly observations of selected variables over the period 1972:1-97:4. Prefix* $d = \Delta$. $E \equiv \overline{E}$ and $e \equiv \overline{e}$.

Table 4.2: ADF tests for unit root, constant and trend included.

Variables	$1-\widehat{\varrho}$	t-ADF	ADF(p)	k
e	-0.094	-3.169	1	4
Δe	-0.815	-8.128**	0	3
cpi	-0.001	-0.111	5	7
Δcpi	-0.521	-4.775**	4	5
cpi^f	-0.006	-1.371	8	11
Δcpi^f	-0.351	-3.864*	7	9
$[\overline{e}\text{-}(cpi\text{-}cpi)]^f$	-0.167	-3.945*	7	7
RB	-0.018	-1.151	8	7
ΔRB	-0.724	-6.573**	6	6
RB^f	-0.057	-2.623	3	6
ΔRB^f	-0.686	-6.748**	3	5
$FI.Y$	-0.619	-5.133**	7	7
$oilp$	-0.079	-2.424	5	5
$oilp^a$	-0.170	-3.850*	5	5
$OILP$	-0.059	-1.733	5	6
$OILP^a$	-0.085	-2.195	5	6

Note: Dickey -Fuller critical values: 5% = -3.457, 1% = -4.057. Constant and trend included. Sample 1972:2-1997:4. [a] When using sample 1974:1-1997:4. p denotes the largest significant lag and k denotes the number of regressors.

5. PPP DESPITE REAL SHOCKS: AN EMPIRICAL ANALYSIS OF THE NORWEGIAN REAL EXCHANGE RATE

Abstract

Despite the emerging consensus on the validity of purchasing power parity (PPP) between trading countries in the long run, empirical evidence in favour of the PPP theory is scarce in data predominantly exposed to real shocks. This essay tests for PPP between Norway and its trading partners using quarterly observations from the post Bretton Woods period, in which the Norwegian economy has been exposed to numerous real shocks such as frequent revaluations of oil and gas resources through new discoveries and price fluctuations. The essay undertakes an extensive examination of the behaviour of the Norwegian real and nominal exchange rates and shows that it is remarkably consistent with the PPP theory. Moreover, convergence towards the equilibrium level appears relatively fast; our estimate of the half life of a deviation from the equilibrium level is only six quarters. This is partly attributed to the Norwegian government's policies aimed at preserving the competitiveness of the economy and the system of centralised wage bargaining.

1. Introduction

A number of recent empirical studies observe convergence towards purchasing power parity (PPP) in the long run, see e.g. Froot and Rogoff (1994), Rogoff (1996), Isard (1995) and MacDonald (1995) and the references therein. Accordingly, changes in nominal exchange rates outweigh changes in domestic prices relative to foreign prices in the long run, and real exchange rates exhibit reversion towards their constant equilibrium rates. The speed of reversion is however relatively low. Consensus estimates of the half life of a deviation from an equilibrium level vary in the range of 2.5 to 6 years for industrial countries. Another common finding is that the support for long run PPP is stronger in data samples dominated by monetary shocks, e.g. in samples from high inflation periods, than in samples presumedly dominated by real shocks such as productivity changes and discovery of natural resources, see e.g. Patel (1990) and Cheung and Lai (2000). In the latter type of samples, real exchange rate behaviour is often indistinguishable from a random walk. Moreover, the support for PPP is often stronger in studies that employ wholesale prices, with a larger share of prices on tradables, rather than consumer prices. Against such a background, this study presents results which would seem to be exceptions within the literature.

This essay tests for PPP between Norway and its trading partners using consumer price indices. It employs quarterly data from the post Bretton Woods era over the period 1972:1-1997:4. This period covers numerous real shocks to the Norwegian economy such as revaluations of Norwegian oil and gas reserves through discoveries of new fields and fluctuations in (real) oil prices. Nevertheless, it finds strong support for PPP, upheld by e.g. a remarkably stable equilibrium real exchange rate over the sample period. Moreover, it is discovered that the half life of a given deviation from the equilibrium rate is only about 1 1/2 years!

The essay argues, however, that the Norwegian government's policies and the system of centralised wage bargaining need to be taken into account when explaining these findings rather than interpreting them as mere indications of relatively strong arbitrage pressure between Norway and its trading partners.

To underpin the PPP theory while taking into consideration the nature of shocks that have hit the Norwegian economy, the essay draws on the "Dutch disease" literature where the effects of revaluations of natural resources on nominal and real exchange rates are among the main issues, see e.g. Corden and Neary (1982) and Bruno and Sachs (1982). Accordingly, the PPP theory can be founded on an intertemporal optimising model that allows for shocks as e.g. revaluations of national wealth. It appears that such shocks, in theory, do not pose a greater challenge to the PPP theory than the Balassa-Samuelson theorem, which has received little support in tests of PPP between industrialised countries at comparable levels of income, see Balassa (1964), Samuelson (1964) and Froot and Rogoff (1994) *inter alia* for an overview of the empirical evidence.

This essay also considers factors that are likely to affect the adjustment of a real exchange rate towards its equilibrium level. In particular, it argues that centralised wage bargaining, which is a feature of a number of small and open industrialised economies, may speed up the adjustment of a real exchange rate towards its equilibrium rate, see e.g. Calmfors and Driffill (1988). For instance, the central wage bargainers may lower their wage claims to absorb adverse shocks to the profitability of the sector for tradables, and thereby restore its competitiveness relative to the abroad and the domestic sector for non-tradables. This contribution of centralised wage bargaining is well known in the literature on wage formation, cf. Aukrust (1977) and Rødseth and Holden (1990), but seems to have escaped attention in the PPP literature. For example, cross country differences in the persistence of real exchange rates do not seem to be classified along this dimension, see e.g. Cheung and Lai (2000).

By considering several factors that may affect the adjustment of a real exchange rate towards its equilibrium level, this essay also sheds light on the commonly observed non-neutrality of nominal exchange rate regimes with respect to the behaviour of real exchange rates, see e.g. Mussa (1986) and Stockman (1983). For instance, real exchange rates often follow random walks in samples from floating regimes, but tend to display mean reverting behaviour in samples from stable exchange rate regimes. This empirical regularity is often regarded as an anomaly and is rarely explained. Mussa (1986), however, suggests that economic policies conducted in conjunction with a fixed exchange rate regime can neutralise the effects from real shocks and might explain their failure to have permanent effects on real exchange rates. This essay elaborates further on how fiscal and exchange rate policies can be, or are, used to affect real exchange rates, but adds that such policies are more likely to be adopted by small and open economies that are more exposed to international arbitrage pressure, which itself is likely to speed up the convergence. Moreover, wages may be set through centralised or coordinated bargaining in small and open economies which may also contribute to mean reversion in real exchange rates. It follows that such additional attributes of an economy need to be controlled for when assessing the partial effect of a nominal exchange rate regime on the behaviour of a real exchange rate.

This essay tests the PPP theory by examining its implications for the behaviour of both the real and nominal exchange rates, see Subsection 2.1. More specifically, it tests whether: (i) the Norwegian real exchange rate is a stationary process with a constant equilibrium rate, (ii) domestic and foreign prices have symmetrical and proportional effects on the Norwegian nominal exchange rate in the long run while (iii), any other variable does not affect the nominal exchange rate. The essay also addresses the contentious issue of whether the PPP theory is for the nominal exchange rate or for prices or for both. That is to say that it is concerned with the issue of whether it is the nominal exchange rate that adjusts to deviations from the parity and brings about the convergence or whether it is prices, or both. If the theory is applied to e.g. domestic prices, implications similar to (ii) and (iii) apply to them as well.

To undertake the empirical analysis, we employ a wide range of univariate and multivariate models. Implication (i) is examined within the framework of a univariate model of the real exchange rate whilst implications (ii) and (iii) are tested within a multivariate system framework. The variable set in the system analysis includes the nominal exchange rate, domestic and foreign consumer prices, domestic and foreign interest rates, Norwegian financial investment abroad and oil prices. Most of these variables are commonly used to model nominal exchange rates and to test the PPP theory, see e.g. Frankel and Rose (1995), Juselius (1992) and Johansen and Juselius (1992). In the system framework, the empirical analysis is based on a full system where all of the variables are regarded as endogenous, and on a partial system where a subset of the variables are considered as given for the purpose of analysis. The system analysis (full and partial) applies the multivariate cointegration procedure developed by Johansen (1988) and (1995b), and previously used by e.g. Juselius (1992) and Johansen and Juselius (1992) in their investigations of PPP with reference to Denmark and the UK, respectively. The test for whether the PPP theory is a theory for the exchange rate and/or for prices can be carried out as an integrated part of this procedure.

Since the system framework treats both prices and the nominal exchange rate as endogenous variables, it avoids possible bias in the results due to invalid *a priori* assumptions, see Isard (1995) and Krugman (1978) *inter alia.* Beside that, if the prices and the exchange rate adjust to a given deviation from the parity, the real exchange rate may adjust faster towards its equilibrium level than if only one of them adjusts. Thus additional insight into the adjustment process of the real exchange rate can be gained by regarding these variables as endogenous. Furthermore, if both the domestic prices and the nominal exchange rate adjust to a deviation from the parity, the nominal exchange rate process cannot be considered as super exogenous for (the long run parameters of) the price process, see Engle et al. (1983) and Johansen (1995b, ch. 8). That is, the price process may not be invariant to a possible change from a stable to a floating exchange rate policy. The question of whether or not the Norwegian price process will withstand a switch to a floating exchange rate regime has been a subject of the Norwegian debate on the choice of a monetary policy target, albeit on theoretical grounds, see the articles in Christiansen and Qvigstad (1997), in particular Holden (1997) and Røedseth (1997a).

The essay proceeds as follows: Section 2 outlines the PPP theory and considers factors that may affect the adjustment towards the long run equilibrium. In particular, Subsection 2.1 defines the absolute and the relative version of the PPP theory and specifies its implications. Section 3 presents a brief overview of the Norwegian economy focusing on factors that seem relevant for the interpretations of results.

Section 4 undertakes the analysis in the univariate framework. However, it starts with a graphical analysis of the Norwegian consumer prices, the consumer prices in the trading countries and the trade weighted nominal exchange rate in the light of the PPP theory.[1]

[1] All empirical results and graphs are obtained using PcGive 9.10, PcFiml 9.10 and GiveWin 1.24, see Hendry and Doornik (1996) and Doornik and Hendry (1996) Doornik and Hendry (1996, 1997).

Section 5 contains the analysis in the system framework. Subsection 5.1 introduces the variables and investigates their time series properties. Subsection 5.2 discusses the choice between a vector autoregressive (VAR) model and a conditional VAR model which is not obvious given the possible stationarity of oil prices (and hence a cointegrating relation by itself).[2] Subsection 5.3 employs a VAR model to determine the number of long run relations between the variables, their identification and to test the validity of using a conditional VAR model in the further analysis. Subsection 5.4 derives a conditional VAR model and evaluates the robustness of the results obtained from the VAR model. In addition, possible endogeneity of prices is formally tested.

Section 6 reiterates the main conclusions and the appendix presents definitions and sources of variables and offers a sensitivity analysis of the main results in Section 5.

2. Purchasing power parity (PPP)

Subsection 2.1 defines the absolute and the relative versions of the PPP theory and lays out its implications for the behaviour of real and nominal exchange rates. Subsection 2.2 presents the economic rationale for the PPP theory together with a brief review of the empirical evidence on the different approaches to justify the theory. Next, Subsection 2.3 considers factors that may influence how fast a real exchange rate converges towards its equilibrium level. Factors brought into focus are centralised wage bargaining and fiscal and exchange rate policies that can be adopted by small and open economies to preserve their competitiveness, especially upon discovery of a natural resource. It is argued that such factors are likely to speed up the adjustment of the real exchange rate towards its equilibrium, through their effects on domestic prices and the nominal exchange rate. This elaboration offers a framework to assess the possible influence of the Norwegian wage and price setting system and that of the government's policies on the real exchange rate behaviour.

[2] In this essay, a variable is considered non-stationary if it is integrated; otherwise it is considered stationary although its process may have changed due to deterministic shifts.

2.1. Definitions and implications

The absolute version of the PPP theory relates the nominal exchange rate to the ratio between the domestic and foreign price levels as

$$E = P/P^f. \tag{2.1}$$

Here E is the nominal exchange rate, that is the price of domestic currency per unit of foreign currency, P is the domestic general price level and P^f is the foreign general price level. In the Casselian view of PPP the nominal exchange rate exhibits a tendency to converge towards the relative price level (P/P^f), see Cassel (1922) and e.g. MacDonald (1995). Thus the relative price level can be interpreted as the equilibrium nominal exchange rate, the rate to which the actual exchange rate (E) reverts in the long run.[3] The actual exchange rate may however be frequently away from its equilibrium level, for short and long periods, owing to changes in numerous economic and political factors.

The relative version of the PPP theory allows for a multiplicative term γ in equation (2.1):

$$E = \gamma P/P^f. \tag{2.2}$$

γ can also be interpreted as the equilibrium level of the real exchange rate. The relative version is defined by a constant equilibrium real exchange rate different from 1, opposed to the absolute version of the theory,

$$R \equiv EP^f/P = \gamma. \tag{2.3}$$

The PPP theory, in both versions, implies a constant real exchange rate in the long run. In Cassel's view, the nominal exchange rate adjusts to outweigh changes in the relative price level (P/P^f) and thereby maintains the level of the real exchange rate. Specifically, a

[3] "...the point of balance towards which, in spite of all temporary fluctuations, the exchange rate will always tend. This parity I call *purchasing power parity.* (Italics original), Cassel (1922, pp. 140).

change in e.g. P is not outweighed by a change in P^f alone or by a combination of changes in P^f and E, but only by a change in E. The opposing view is that the PPP relation determines domestic prices (P) for a given level of foreign prices and nominal exchange rate, especially for a relatively small economy with a fixed exchange rate, cf. Edison and Klovland (1987) and Giovannini (1988). Another view is that both the nominal exchange rate and the domestic prices are simultaneously determined in the PPP equation, see Isard (1995, pp. 59) and MacDonald (1995).

The PPP theory also implies a symmetry and a proportionality restriction on domestic and foreign prices in a nominal exchange rate equation, or alternatively on the nominal exchange rate and foreign prices in an equation for domestic prices. These restrictions follow from the implied constancy of the real exchange rate. These can be defined using equation (2.4), which is formulated by a slight generalisation of equation (2.2) and taking logs denoted by small letters.

$$e = \ln \gamma + \pi_1 p - \pi_2 p^f. \tag{2.4}$$

π_1 and π_2 denote constant elasticities; Greek letters without subscript t are used to denote constant parameters, throughout the essay. The symmetry restriction is defined by $\pi_1 = \pi_2 = \pi$, and the proportionality restriction by adding $\pi = 1$.

Obviously, the PPP theory also implies the absence of any other variable, say w, from a long run relation between the nominal exchange rate and domestic and foreign prices. For example, the coefficient of w should be zero if it is added to equation (2.4), though w may have short run effects on e.g. the nominal exchange rate.

This essay tests the PPP theory by examining each of its implications. Section 4 sheds light on the empirical validity of equations (2.1)-(2.3). There, equation (2.3) provides the foundation of the univariate analysis of the real exchange rate. Section 5 is concerned with testing the symmetry and proportionality restrictions, see equation (2.4), and the absence of long run effects of variables that are commonly regarded as important determinants of the Norwegian nominal and real exchange rates. Moreover, Section 5 formally tests whether or not both the nominal exchange rate and prices respond to deviations from the

PPP and contribute to re-establish the parity.

2.2. Economic rationale

This subsection outlines three different but complementary foundations of the PPP theory. First, it reviews the two main and oldest justifications for the PPP, namely the PPP as a long run neutrality proposition and the commodity arbitrage view. The neutrality proposition view does not address the effects of real shocks while the commodity arbitrage view can be criticised for an unsatisfactory treatment of non-tradables goods and services. These shortcomings are explicitly addressed by the third approach, here called "the structural approach", which divides the economy into a sector for tradables and a sector for non-tradables and examines the short run and long run effects of real shocks, as e.g. productivity shocks and discovery of natural resources. Presentation of the latter approach draws extensively on the "Dutch disease" literature.

A long run neutrality proposition: Within the classical quantity theory of money, the PPP theory is regarded as a long run neutrality proposition, see e.g. Samuelson (1964), Officer (1976) and McCallum (1996, pp. 32). Accordingly, an increase in e.g. the domestic money supply ultimately lead to an equiproportionate change in nominal variables leaving real prices such as the real exchange rate unchanged.

This view suggests a broad measure of the general price level to be used in tests of the PPP theory, see e.g. Officer (1976). This, because the nominal exchange rate is interpreted as the relative price of domestic and foreign money, which is defined by the broadest measure of the domestic and foreign general price levels.

The neutrality proposition view is often supported by empirical evidence, since the PPP theory is seldom rejected in data samples from high inflation periods when monetary shocks are likely to be predominant, see e.g. Froot and Rogoff (1994), Rogoff (1996) and MacDonald (1995).

The commodity arbitrage view: The PPP theory is commonly underpinned by assuming commodity arbitrage between open economies. Following this view, prices of tradables tend to equalise across countries when measured in the same currency:

$$P_i = EP_i^f, \qquad \forall\ i. \tag{2.5}$$

P_i and P_i^f denote the price of good i in the domestic and foreign currency, respectively. Equations like (2.5) are commonly referred to as the law of one price. The absolute version of the PPP theory is derived by assuming that P and P^f are aggregates of the same set of prices with equal weights, see e.g. Rogoff (1996).

A host of issues are often raised against the validity of the law of one price. For instance, (i) all goods are not traded, hence the possibility of arbitrage does not exist for nontraded goods, (ii) even traded goods are not homogenous, thus they need not command the same price, (iii) there are transaction costs that lead to a wedge between prices across countries and (iv), prices are set in accordance with the market conditions within a given economy (pricing to market) and are therefore likely to differ across markets, see e.g. Rogoff (1996), Krugman (1987) and Goldberg and Knetter (1997) and the references therein. Actually, numerous empirical studies find persistent deviations from the law of one price, even for highly traded commodities, see e.g. Engle and Roger (1996) and Goldberg and Knetter (1997).

Aggregation from prices at the micro level to general price levels is also subject to criticism. It is pointed out that aggregate prices, in general, are not composed of the same set of prices at micro levels due to variations in consumption and production patterns across countries. For a similar reason, even prices of the same set of goods need not enter the aggregates with constant weights over time. Therefore, aggregate price levels often measure the prices of different baskets of goods across countries and over time. Consequently, they are likely to differ from each other when measured in the same currency, even when the law of one price holds at the micro level.

Empirical studies that subscribe to the commodity arbitrage view employ price aggregates that contains a relatively large number of prices on tradables, for instance export prices, import prices or wholesale prices. Moreover, they use the relative PPP theory by (implicitly) assuming constant transaction costs, constant mark-ups and constant differences in commodity baskets across countries. The adoption of the relative PPP theory can also be justified by the lack of updated series for price levels while indices for general price levels are published regularly by governments. A few recent empirical studies have also allowed for non-linear adjustment in real exchange rates due to presumably large transaction costs, see e.g. Dumas (1992) and Taylor et al. (2000).

The commodity arbitrage view is supported by empirical studies that employ e.g. wholesale price indices and/or long samples of data, as they often reports results in favour of the PPP theory. These studies, however, suggest that arbitrage pressure is too weak to bring about rapid price convergence. Moreover, non-linear models suggest that the arbitrage may be effectively absent at small deviations from the equilibrium level due to large transaction costs.

The structural approach: Balassa (1964) and Samuelson (1964) take into account the effects of prices on non-tradable goods and services and point out that productivity differences between countries may induce persistent deviations from PPP between these countries. They show that prices of internationally non-traded goods (and services) and hence aggregate prices rise with an increase in productivity. Thus, higher domestic productivity growth lead to appreciation of the real exchange rate. Also, it follows that any other real shock that raises the growth in domestic prices on non-tradables relative to the foreign prices on non-tradables will initially have a similar effect on the real exchange rate.

For instance, a discovery of natural resources or any other real shock, which increases the consumption opportunities of an economy, will normally increase the demand for both traded and non-traded goods. The demand for traded goods can be satisfied by import at given international prices while non-traded goods have to be produced at home, at higher prices in general. As a result, the real exchange rate will appreciate, see e.g. Corden (1984). A rise in government expenditures is another popular example, since they are believed to be biased towards non-traded goods and services. Accordingly, an increase in government expenditures raises their price and thereby bring about a real appreciation.

However, real shocks other than shocks to the production technology only cause temporary changes in the real exchange rate under fairly standard conditions. Bruno and Sachs

(1982), among others, show within an intertemporal optimising framework that discovery of natural resources or their revaluation leave the long run real exchange rate unaffected, if: (i) there are constant returns to scale in the sector for tradables and non-tradables, (ii) the labour force is mobile between the two sectors and (iii), capital is mobile between the sectors and between countries. Under these conditions, the supply of non-tradables is perfectly elastic, thus, all demand shocks are absorbed by output changes at unchanged prices of non-tradables, see also Corden (1984). As a result, the ratio between the prices of non-tradables and tradables, say P_{nt}/P_{tr}, remains constant, and consequently the real exchange rate R too.

The implied constancy of the equilibrium real exchange rate, as in the relative version of the PPP theory, can be exposited as follow. Suppose that $P_{nt}/P_{tr} = \tau$, in the long run, where τ is determined by the production technology in both sectors. Assume that the general price levels at home and abroad are homogenous in sectoral price levels, as e.g. the domestic price level in equation (2.6), where ν is the weight of the price on tradables. Furthermore, assume that foreign countries have a similar economic structure such that $P^f_{nt}/P^f_{tr} = \tau_f$.

$$P = P^{\nu}_{tr} P^{1-\nu}_{nt} \tag{2.6}$$

Then, the equilibrium real exchange rate can be defined as the constant $\tau_f^{1-\nu^f}/\tau^{1-\nu}$ in equation (2.7) which can be interpreted as γ in equation (2.3).[4]

$$R \equiv EP^f/P = EP^f_{tr}\tau_f^{1-\nu^f}/P_{tr}\tau^{1-\nu} = EP^f_{tr}\tau_f^{1-\nu^f}/EP^f_{tr}\tau^{1-\nu} = \tau_f^{1-\nu^f}/\tau^{1-\nu} \equiv \gamma \tag{2.7}$$

Note that the equilibrium real exchange rate is equal to 1, as in the absolute version of the PPP theory, in the absence of cross country differences in production technologies and in the construction of general price levels, i.e. $\nu^f = \nu$. In the short run, the real exchange rate will appreciate or depreciate depending on whether $P_{nt}/P_{tr} > \tau$ or $P_{nt}/P_{tr} < \tau$.

The derived equilibrium real exchange rate is not inconsistent with long run current

[4] The denominator in the term after the second identity follows from the assumption of the law of one price in the sector for tradables: $P_{tr} = EP^f_{tr}$.

account balance. A rise in consumption opportunities, due to e.g. discovery of oil, and the ensuing rise in P_{nt}/P_{tr}, will lead to a transfer of production factors from the sector for (non-oil) tradables to the sector for non-tradables, and thereby to a decline in the sector for tradables, which is termed as the "Dutch disease", see e.g. Corden and Neary (1982). As a result, the non-oil trade surplus will fall but the total trade surplus need not deteriorate since the export of oil can make up for the fall in the non-oil trade surplus. Intertemporal optimising behaviour by consumers implies that P_{nt}/P_{tr} is above τ, only as long as the sector for tradables is larger than required to ensure long run current account balance. P_{nt}/P_{tr} will return to its long run level τ and R will return to $\tau_f^{1-\nu^f}/\tau^{1-\nu} \equiv \gamma$, when the sector for (non-oil) tradables has reached a level that does not compromise the economy's ability to pay for its imports, cf. Bruno and Sachs (1982).

The next subsection elaborates on the adjustment towards the long run equilibrium. In particular it focuses on the role of centralised or coordinated wage bargaining and of fiscal and exchange rate policies in counteracting a deviation from the equilibrium to e.g. avoid the "Dutch disease". Section 3 suggests that the features discussed below may have played an important role in determining the course of the Norwegian nominal and real exchange rates in the post Bretton Woods era.

2.3. Speed of adjustment towards the long run equilibrium

Beside the nature of a shock, market imperfections play an important role in determining how fast a real exchange rate (R) adjusts towards its equilibrium level. After an unexpected monetary shock, for example, the adjustment is commonly expected to take from 1 to 2 years because of menu costs and wage contracts of different lengths. Generally, the adjustment time will also depend upon the degree of international competition and on how fast production factors are transferred from sectors that are adversely affected by a shock to those that are experiencing a boom following the shock.

As noted earlier, international competition is believed to be weak due to tariff and non-tariff costs and, according to some accounts, effectively absent at small deviations from the equilibrium level due to large transaction costs. The intra-country reallocation of resources is also believed to be slow, e.g. because of slow labour mobility across sectors.

Thus, the product prices in these sectors can deviate from their equilibrium levels for a long period and prevent the real exchange rate reaching its equilibrium rate for the same period. Note that a sluggish adjustment of the prices of non-tradables can delay the adjustment of the real exchange rate, even when the law of one price holds continuously in the sector for tradables.

Sluggish adjustment to the equilibrium level of the real exchange rate can be formalised by equilibrium correction models of domestic prices and the nominal exchange rate, as:

$$\Delta p_t = +\alpha_p(e - (p - p^f))_{t-1} + \textit{residual}, \tag{2.8}$$

$$\Delta e_t = -\alpha_e(e - (p - p^f))_{t-1} + \textit{residual}. \tag{2.9}$$

Where Δ defines the first difference operator. These models are derived under the assumption of no cross country differences in production technologies and in the construction of general price levels, cf. equation (2.7). Consequently, the law of one price holds for aggregate prices in the long run, i.e. absolute PPP. The term $(e - (p - p^f))$ represents the deviation from the absolute PPP, while α_p and α_e denote to what extent Δp_t and Δe_t respond to the deviation in the subsequent period. In general, foreign prices may also adjust, but for simplicity they are assumed to be unresponsive, which is not unlikely in the case of a small domestic economy. The residuals represents lagged and contemporaneous effects of Δe_t, Δp_t^f and of other relevant factors.

Centralised wage setting and government policies

Centralised or coordinated wage bargaining between labour and employer unions is a market imperfection that may nevertheless contribute to fast adjustment towards the equilibrium level. It has often been pointed out that centralised wage bargainings tend to take into account the effects of adverse shocks to the overall economy by e.g. wage moderation, see e.g. Calmfors and Driffill (1988) and Jackman (1990). Especially, effects of shocks that may affect the viability of the sector for tradables given its limited opportunities, due to the arbitrage pressure, to offset higher factor costs by raising product prices. Such restraints on wages are built into the Scandinavian model of inflation, see Aukrust (1977) and Calmfors (1977). This model offers a blueprint to preserve international competitiveness and to avoid a transfer of resources from the sector for tradables to the sector for non-tradables following a shock. Accordingly, the wage settlement in the sector for tradables is adopted by the sector for non-tradables. An exposition of this model may be useful in clarifying its implications for the behaviour of prices and the real exchange rate.

The Scandinavian model of inflation is a two-sector model derived under conditions similar to those in e.g. Bruno and Sachs (1982), noted above. Consequently, the wage share, or alternatively the return on capital, is the same across sectors over time. If we add the condition that the aggregate price level is defined as above by equation (2.6), the model can be presented in the growth form as follows:

$$\begin{aligned}
\Delta p_{tr,\,t} &= \Delta e_t + \Delta p^f_{tr,\,t} && (2.10)\\
\Delta w_{tr,\,t} &= \Delta p_{tr,\,t} + \Delta q_{tr,\,t} && (2.11)\\
\Delta w_{nt,\,t} &= \Delta w_{tr,\,t} && (2.12)\\
\Delta p_{nt,\,t} &= \Delta w_{nt,\,t} - \Delta q_{nt,\,t} && (2.13)\\
\Delta p_t &= \nu \Delta p_{tr,\,t} + (1-\nu)\Delta p_{nt,\,t} && (2.14)
\end{aligned}$$

Equation (2.10) follows from the law of one price in the sector for tradables, while (2.11) states that wage growth in this sector equals the sum of growth in prices ($\Delta p_{tr,\,t}$) and labour productivity ($\Delta q_{tr,\,t}$). This sum is referred to as the "main course" or the "Scandinavian path" of wage growth. Equation (2.12) indicates that the wage growth in the sector for non-tradables follows the wage growth in the sector for tradables. This property is often considered the hallmark of this model, see e.g. Nymoen (1991). Prices in the sector for non-tradables are set as a constant mark up on unit labour costs, hence their growth equals wage growth in excess of productivity growth, see equation (2.13). Finally, equation (2.14) defines the growth rate in aggregate prices (inflation).

The model implies that domestic inflation depends on nominal exchange rate depreciation (Δe_t), growth rate in international prices on tradables ($\Delta p^f_{tr,\ t}$) and the productivity difference between the sector for tradables and non-tradables. The effect of the productivity difference is influenced by the share of prices of non-tradables in aggregate prices, since productivity growth in the sector for tradables affects the wage growth and thereby the price growth in the sector for non-tradables. In a relatively open economy, defined by a large ν, the effects of productivity differences on Δp_t can be negligible.

$$\Delta p_t = \Delta e_t + \Delta p^f_{tr,\ t} + (1-\nu)(\Delta q_{tr,\ t} - \Delta q_{nt,\ t}) \tag{2.15}$$

The Balassa-Samuelson theorem follows directly from this model if we assume a similar inflation process in foreign countries. Consider the real exchange rate in growth rate terms:

$$\begin{aligned}
\Delta r_t &\equiv \Delta e_t + \Delta p^f_t - \Delta p_t \\
&= \Delta e_t + \{\Delta p^f_{tr,\ t} + (1-\nu^f)(\Delta q^f_{tr,\ t} - \Delta q^f_{nt,\ t})\} - \{\Delta e_t + \Delta p^f_{tr,\ t} + \\
&\quad (1-\nu)(\Delta q_{tr,\ t} - \Delta q_{nt,\ t})\} \\
&= (1-\nu^f)(\Delta q^f_{tr,\ t} - \Delta q^f_{nt,\ t}) - (1-\nu)(\Delta q_{tr,\ t} - \Delta q_{nt,\ t}) \equiv \xi_t \qquad (2.16)
\end{aligned}$$

Equation (2.16) suggests that the real exchange rate appreciates if the difference in the productivity growth between the domestic sectors for tradables and non-tradables is higher than that in abroad. In the absence of such differences across countries and $\nu^f = \nu$, the real exchange rate will be constant, i.e. $\xi_t = 0$. The real exchange rate may be constant for a number of other reasons too, as evident from equation (2.16).

The Scandinavian model of inflation describes the long run growth in wages and prices, or alternatively, offers a normative account of wage behaviour. Actual growth in wages and prices is however unlikely to accord with this model in the short run. Among others, cyclical factors and wage-wage effects across sectors may be present in the short run, cf. Rødseth and Holden (1990) and Nymoen (1991). Nevertheless, the model suggests that aggregate prices tend to adjust such that the real exchange rate converges towards its equilibrium level. Thus the real exchange rate is likely to adjust faster towards the equilibrium level if wage and price setters adhere to this model.

Appropriate fiscal policies have been mentioned among mechanisms that can encourage adherence to the wage and price behaviour suggested by the Scandinavian model, see e.g. Aukrust (1977). In this regard, tax cuts, subsidies, selective increase in government expenditures, can be traded for wage and price moderation in the face of e.g. demand shocks, see also Corden (1995). Such measures, denoted as income policies, have been employed in e.g. the Nordic countries in addition to legislative actions to counteract wage and price growth in excess of the Scandinavian path, see Calmfors (1990).

The additional parameters α_b and α_g in the equilibrium correction model recognise, respectively, the possible contribution of centralised wage bargaining itself, and of income policies in increasing the speed of adjustment towards the equilibrium.

$$\Delta p_t = +(\alpha_p + \alpha_b + \alpha_g)(e - (p - p^f))_{t-1} + residual \tag{2.17}$$

Nominal exchange rate policy can also contribute to maintain the competitiveness of the sector for tradables by affecting the behaviour of the nominal exchange rate. The Dutch disease literature discusses frequent devaluations, or prevention of appreciations that would otherwise have taken place in the face of revaluation of natural resources, as measures to avoid large fluctuations in the real exchange rate, see e.g. Corden (1995). A stable nominal exchange rate policy may itself have similar effects on the real exchange rate, even in the absence of such "exchange rate protection" policies, as they are termed.

The influence of a stable nominal exchange rate policy can be recognised by the additional term in the nominal exchange rate process:

$$\Delta e_t = -\alpha_e(e - (p - p^f))_{t-1} - \alpha_c(e - \overline{e})_{t-1} + residual \tag{2.18}$$

Here α_c indicates how fast the central bank eliminates a deviation from the central parity, which is usually set equal to the perceived equilibrium nominal exchange rate.

There is ample evidence to suggest that central banks quite often define the equilibrium nominal exchange rate in the light of the PPP theory, either by using aggregate consumer prices, wholesale prices, unit labour costs or other related measures, see e.g. Cassel (1922), Officer (1976), Isard (1995) and Hinkle and Montiel (1999). If $\overline{e}$ is set equal to $p - p^f$, presumed to be fixed for the sake of argument, the authorities can speed up the convergence of the nominal exchange rate to its equilibrium level, $p - p^f$, and thereby also the adjustment of the real exchange rate towards its equilibrium.

It is, however, more realistic to assume that $\overline{e}$ is set equal to e.g. $(p - p^f)$ in a chosen base period t^*. Thus, the actual nominal exchange rate e_t will be directed towards the central parity $\overline{e} \equiv (p - p^f)_{t^*}$, which can lead to persistent deviations from its time varying equilibrium $(p - p^f)_t$. But, if $\overline{e}$ is adjusted frequently, $\overline{e}$ and hence e_t will track $(p - p^f)_t$ more closely. Nevertheless, even if $\overline{e}$ is not adjusted frequently, a stable exchange rate policy is likely to contain excessive fluctuations in the nominal exchange rate and keep it close to, if not equal to, its equilibrium level. A choice of another measure related to $p - p^f$ is likely to have a similar effect on the nominal and real exchange rates, but probably with less strength.

Exchange rate protection policies can be interpreted as frequent changes in $\overline{e}$, and/or as avoidance of short term deviations between e and $\overline{e}$ in the face of appreciation pressure.

The effects on the real exchange rate behaviour of e.g. centralised wage bargaining, fiscal and exchange rate policies are summarised below.

The real exchange rate process To avoid more notation, assume that $\Delta p_t^f = 0$, then a change in the real exchange rate Δr_t depends on the level of the real exchange rate in the following way:

$$\begin{aligned} \Delta r_t &\equiv \Delta e_t - \Delta p_t - \Delta p_t^f = -(\alpha_e + \alpha_p + \alpha_b + \alpha_g + \alpha_c)(e - (p - p^f))_{t-1} + \text{residual} \\ \Delta r_t &= -(\alpha_e + \alpha_p + \alpha_b + \alpha_g + \alpha_c) r_{t-1} + \text{residual} \end{aligned} \tag{2.19}$$

Equation (2.19), which is obtained by using (2.17) and (2.18), suggests that the real exchange rate is a stationary process when $(\alpha_e + \alpha_p) > 0$, i.e. when arbitrage pressure exists. The adjustment towards the long run equilibrium will be slow, however, if $(\alpha_e+\alpha_p)$ is small, unless the country has a fixed exchange rate regime, centralised wage bargaining and/or active government policies aimed at maintaining the competitiveness. Such features speed up the adjustment towards the long run equilibrium. If $(\alpha_e+\alpha_p)$ is small, the nominal exchange rate is floating and neither the labour market nor the government are concerned about the competitiveness of the economy, i.e. $\alpha_c = \alpha_b = \alpha_g = 0$, the real exchange rate may appear to follow a random walk in small samples of data.

The implied effects of the nominal exchange rate regime on the real exchange rate behaviour are consistent with the observations that real exchange rates tend to display stationary behaviour within fixed exchange rate regimes and appear to follow random walks under floating regimes, see e.g. Mussa (1986) and Stockman (1983). It has been argued that a stable exchange rate regime has a stabilising effect on expectations about the future course of a nominal exchange rate, see Mussa (1986). Moreover, a stable exchange rate policy is often backed up by fiscal policies that have stabilising effects on the nominal exchange rate and prices.

The analysis in this section, however, has focused on the role of centralised or coordinated wage setting and income policies in affecting the real exchange rate behaviour through their influence on domestic prices. In addition, a stable nominal exchange rate policy can influence the real exchange rate behaviour by the choice of an appropriate central parity, or through rapid elimination of deviations from the parity. In general, centralised wage bargaining, income policies and stable exchange rate policies are all features of relatively small (industrialised) economies that takes international prices as given and have relatively large sectors for tradables, i.e. economies that are exposed to relative large arbitrage pressure. It follows that such additional features need to be taken into account when assessing the partial effect of a nominal exchange rate regime on the real exchange rate behaviour.

The next section sketches some of the main features of the Norwegian economy, focusing on the wage and price setting behaviour, income and exchange rate policies. It appears that these factors have contributed to preserve the competitiveness of the economy in the face of considerable demand shocks. This overview will be helpful when interpreting the empirical results. For example, it may be misleading to ascribe our evidence in favour of the PPP theory to the Norwegian stable exchange rate policies and/or the arbitrage pressure.

3. The Norwegian economy

The Norwegian economy is relatively small, open and commodity based relatively to most of the other (western) European economies. The sum of its export and import is more than 1/3 of its GDP, which is almost the twice of that for e.g. France, Germany and Italy, see Haldane (1997). Around 1/2 of the total export is commodity based which is sold at fluctuating world prices, while imports are mainly manufactured goods with more stable prices. It is therefore exposed to considerable terms of trade shocks. The economy's openness is also testified by the relatively large weight, about 40%, in the Norwegian consumer price index (CPI) of prices on imported goods and domestically produced goods exposed to foreign competition.

Norway fully dismantled regulations on capital flows between Norway and abroad in July 1990; the deregulation process started in the early 1980s, see Olsen (1990). Domestic financial markets were also deregulated in about the same period, see Bårdsen and Klovland (2000). These structural changes have contributed to large fluctuations in the saving rate and financial investments abroad. Moreover, they have amplified the effects of other monetary, fiscal and external shocks to the economy and thereby increased fluctuations in the activity level, see e.g. Rødseth (1997b).

Norway has been a net exporter of crude oil since 1975 and of natural gas since 1977. Oil and gas production has constituted a fairly large share of its GDP, roughly 10-20% since the mid 1970s, see Aslaksen and Bjerkholt (1986) and Statistics Norway (1998). Oil and gas exports make up a large share of Norway's total export of goods and services; more than 1/3 of it in the period 1991-97, when oil production rose from about 1.7 to 3 million barrels a day, see Statistics Norway (1998). The petroleum wealth has been revalued several times due to large oil price fluctuations and discovery of new oil and gas fields. Its size can be gauged by the government's share of permanent income, which was estimated at between 48 and 66 billions krone in 1997, see Olsen (1997). Oil (and gas) revenues have been mainly used to finance non-oil fiscal deficits, on average 36 billion krone a year over the period 1980-96. The excess revenues have been used to cover non-oil current account deficits and to acquire financial assets abroad. The latter in order to mitigate the adverse shocks to the sector for non-oil tradables and as precautionary saving, see e.g. Olsen (1997). Positive net effects of a rise in oil prices on the Norwegian GDP, i.e. after taking account of the negative effects on traditional export industries due to lower demand abroad, are documented in e.g. Eika and Magnussen (2000) and Mork et al. (1994).

The wage setting process is relatively centralised and to a large extent guided by the Scandinavian model of inflation, which also offers a quite good first approximation to the actual wage and price behaviour, see e.g. Rødseth and Holden (1990), Johansen (1995a) and Nymoen (1991). The central wage negotiations usually take place every second year but allows for local wage adjustment in the intermediate year. These adjustments are often small compared with the centrally determined wage growth and, arguably, to some extent accounted for in the central accord, see Rødseth and Holden (1990) and Holden (1990) for details. It has however been pointed out that the total wage growth, and even the centrally set wages, have often been above the main course. Nevertheless, the wage share in both sectors has been fairly stable over time. This has been partly ascribed to government interventions aimed at dampening the wage and price growth, see e.g. Calmfors (1990) and Rødseth and Holden (1990). Direct government interventions have mostly occurred since the early 1970s, which can be interpreted as a reflection of large pressure on wages due to the effects from oil wealth and revenues, see the history of wage bargainings since 1946 in Rødseth and Holden (1990, pp. 240).

Figure 4.2.b shows that (quarterly) inflation in Norway has largely been comparable to inflation in its trading partners in the post Bretton Woods period, except during the 1980s. During this decade, the Norwegian inflation was generally above that of its trading partners. However, neither in Norway nor in its trading partners has the inflation rate

been at sufficiently high levels to be termed hyperinflation, which would have indicated the predominance of monetary shocks.

Concern for the viability of the sector for tradables is not only reflected in the choice of the wage setting system, but seems to have been the guiding principle for the governments involvement in wage setting and for the Norwegian exchange rate policy, at least since the discovery of North Sea oil. The government plays an active role in the wage and price setting processes to maintain the competitiveness of the non-oil sector for tradables. A permanent government commission serves the central wage bargainers by providing background information and outlining the consequences of different wage settlements. As noted above, the government has also intervened directly in the wage setting process whenever the wage growth has been far in excess of the main course. In 1973 and 1975-80, tax cuts and/or increases in government expenditures were traded for wage moderation. The government has also reduced taxes in advance of wage bargainings in order to influence their outcomes.

Since the end of the 1970s, the government involvement in wage negotiations has been less direct and it has to a large extent resorted to legal actions or voluntarily agreements to contain growth in wages and prices. In the period September 1978-December 1979, 1981 and later in 1988-89, growth in wages and prices was regulated through legislative actions, see Rødseth and Holden (1990). These wage and price controls had a significant effect on the wage growth during these years, though their initial effects were to some extent outweighed by compensatory wage growth in the subsequent wage settlements, see e.g. Nymoen (1989). Since 1992 the government has entered an agreement with the largest confederation of workers union (LO) to provide nominal exchange rate stability in exchange for wage moderation which is termed "the solidarity alternative". Its success in terms of wage moderation is contended, however, see e.g. Evjen and Nymoen (1997) and Alexander et al. (1997).

Norway has mainly pursued a policy of exchange rate stabilisation against western European countries since the end of Bretton Woods system in 1971. Norway participated in the European "snake agreement" until 1978, but linked the krone to a trade weighted basket of currencies, mainly composed of western European currencies, when the ERM was established. Since October 1990, the krone has been stabilised against the ECU. Despite the stated aim of currency stabilisation, the value of the Norwegian krone has shown relatively large fluctuations however. The fluctuations have been allowed by changes in the currency basket and in the fluctuation margins across time. In addition, the krone has been devalued about 10 times since 1972, mostly during the 1970s. During the 1980s, the 9.2% devaluation of the krone in May 1986 stands out. This also marks a switch to a no-devaluation stance, since the krone has not been devalued since then. The devaluations were mainly carried out to correct for the weakening of the competitiveness of the economy due to excessive wage and price growth, see e.g. Norges Bank (1987), Skånland (1983) and Rødseth and Holden (1990).

Major changes in the value of the krone have often coincided with large fluctuations in oil prices. The devaluation in May 1986, was also preceded by a fall in oil prices. The krone was allowed to float in January 1997 and in the autumn of 1998 following appreciation and depreciation pressure, respectively. The value of the krone changed by about 10% relative to the ECU on both occasions, which coincided with unusually high and low oil prices, respectively, see Norges Bank (1997) and (1998). In addition to these possibly oil price related currency crises, the krone was also allowed to float against the ECU in December 1992, following strong depreciation pressure triggered by the crisis in the ERM. The practice of exchange rate stabilisation against the ECU was resumed briefly afterwards, but without formal fluctuation margins, see Norges Bank (1995) and Alexander et al. (1997).

To summarise, the Norwegian economy is small and open with a large export share of commodities and hence exposed to large fluctuations in the terms of trade. However, the Norwegian wage setting system, the fiscal and exchange rate policies have been aimed at preserving the international competitiveness of the sector for non-oil tradables and to avoid growth in the sector for non-tradables at the expense of the former, especially since the discovery of relatively large oil resources. Therefore, these and other possible real shocks to the economy may not have affected the real exchange rate beyond the short run. Indeed, if wage and price setters adhere to the Scandinavian model of inflation, real shocks affecting prices through demand are not expected to last beyond the contract period of two years, as for monetary shocks, see Subsection 2.3. The next section, however, draws a more nuanced picture of the real exchange rate behaviour.

4. Testing PPP in a univariate framework

The next subsection describes the data series for the nominal exchange rate and domestic and foreign consumer prices, focusing on their relation to each other in the light of the PPP theory. Thereafter, Subsection 4.2 formalises the PPP theory in a univariate framework and employs an augmented Dickey Fuller (ADF) test to examine whether the real exchange rate behaviour is consistent with the PPP theory. This subsection also plots the implied equilibrium real exchange rate over time together with a 95% confidence interval to expose possible changes over time. The speed of adjustment towards the equilibrium rate is another main issue in this subsection.

4.1. Prices, the nominal and the real exchange rate

Figure 4.1 displays the value of the Norwegian trade weighted (effective) nominal exchange rate (E) and the Norwegian consumer price index (CPI) relative to a trade weighted index of foreign consumer prices (CPI^f) over the period 1972:1-1997:4. The trade weights are tabulated in Appendix A and suggest that Norwegian trade is not confined to a single dominant trading partner. This makes it more relevant to focus on the trade weighted exchange rate and prices rather than on bilateral exchange rates and prices.

The overall impression from Figure 4.1 is that the actual exchange rate (E) does not evolve independently of the relative consumer prices (CPI/CPI^f), though it is more volatile than the relative consumer prices. One noticeable exception from this tendency is the relatively strong exchange rate appreciation until 1976/77, which is not matched by a comparable fall in the relative consumer prices in this period. From the late 1970s to about 1992, however, changes in the relative consumer prices appear to precede fluctuations in the exchange rate fluctuations. The large devaluation in May 1986 is commonly ascribed to the fall in oil prices in 1985/86, but the figure shows that it was also preceded by a rise in the relative consumer prices. The exchange rate fluctuations since 1992, in particular, seem to be driven by other factors than the relative consumer prices, since the latter are

Figure 4.1: *Trade weighted nominal exchange rate, E, (solid line) and relative consumer prices between Norway and trading partners, CPI/CPI^f, (boxed line).*

relatively stable in this period, see also Figure 4.2.d. However, the nominal exchange rate evolves around the relative consumer prices, which may be considered as the equilibrium level of the nominal exchange rate in the PPP framework, see Subsection 2.1.

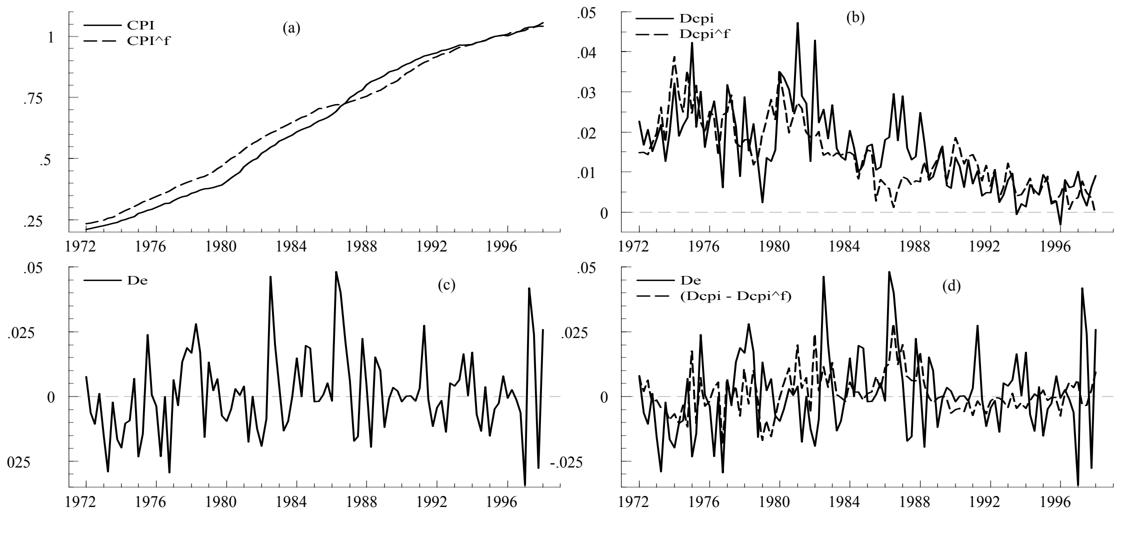


Figure 4.2: *The graphs display quarterly observations of different variables over period 1972:1-1997:4. (a) Consumer price indices for Norway and its trading partners, CPI and CPI^f, respectively. (b) Quarterly inflation rates, denoted by Dcpi and Dcpi^f. The prefix "D" denotes the first difference Δ. (c) Quarterly depreciation rate, De. (d) The depreciation rate and inflation differences.*

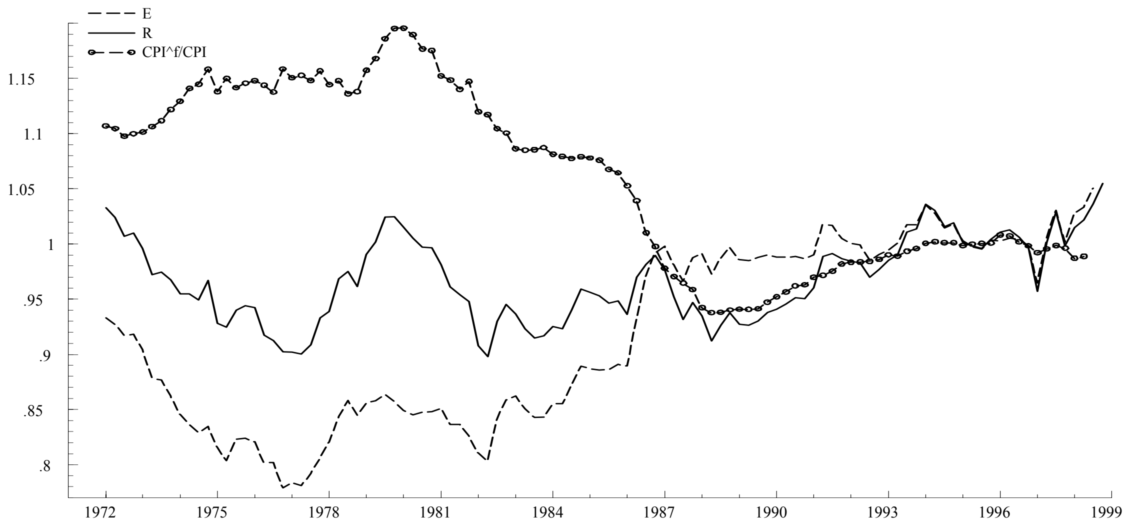


Figure 4.3: *Nominal exchange rate, E, (dashed line), real exchange rate, R, (solid line) and relative consumer prices between trading partners and Norway, CPI^f/CPI (circled line).*

Figure 4.3 shows that changes in the nominal exchange rate (E) and the (redefined) relative consumer prices (CPI^f/CPI) tend to outweigh each other, especially before the 1990s. Thus fluctuations in the real exchange rate, $R \equiv E(CPI^f/CPI)$, have a smaller range compared with the fluctuations in the nominal exchange rate and the relative consumer price level. In the late 1980s and early 1990s movements in the real exchange rate are mainly driven by the relative consumer prices since the nominal exchange rate is relatively stable in this period. The case is the opposite after 1992/93 when fluctuations in the real exchange rate can be almost entirely ascribed to fluctuations in the nominal exchange rate, as the relative consumer prices are quite stable. Figure 4.3 also indicates that the tendency of the nominal exchange rate and the relative consumer prices to outweigh each other is stronger when the fluctuations are large compared with when they are small. This is apparent when the period before e.g. 1987 is compared with the period afterwards.

Figure 4.3 gives the impression that the real exchange rate is likely to have evolved around a constant level and that the range of its fluctuations have possibly declined over time. At least, it is not obvious that the range has steadily increased over time, as implied by the random walk hypothesis of real exchange rate, cf. Mark (1990). Indeed, the behaviour of the real exchange rate resembles that of a (weak) stationary time series that displays fluctuations within a given range and reversion towards a constant mean level. The next subsection tests this impression formally.

4.2. Augmented Dickey Fuller (ADF) test

The PPP theory implies that the real exchange rate (R) evolves around a constant mean or equilibrium level, γ, over time, see equation (2.3). This proposition can be formalised as follows:

$$R_t = \gamma + \sum_{i=1}^{\mathrm{p}} \psi_i (R_{t-i} - \gamma) + \varepsilon_t. \tag{4.1}$$

Here ε_t is the error term, assumed to be identically, independently normally distributed with zero mean and constant variance, σ^2, i.e. IIDN(0, σ^2).

Under the PPP theory, the (absolute) value of $\sum_{i=1}^{\mathrm{p}} \psi_i$, hereafter denoted as ϱ, should be less than 1. Otherwise R will not revert to its equilibrium level γ, i.e. not be a stationary process. The speed of adjustment towards the equilibrium level is inversely related to the (absolute) value of ϱ. R follows a random walk process if $\varrho = 1$, in which case every shock has a permanent effect on R, making it drift away randomly from any level. If $\varrho = 0$, every shock has a transitory effect on R and it will fluctuate randomly around γ. This case can be interpreted as a formalisation of the PPP theory if it holds continuously. But, even Cassel recognised that purchasing power parity is frequently violated and a given violation can persist for long time periods, see Cassel (1922).[5] This notion of the PPP theory corresponds to a value of $\varrho \in (0, 1)$.

Restrictions on ϱ can be tested by employing the ADF test, see e.g. Banerjee et al. (1993, ch. 4) for details. The null hypothesis under the ADF test is that the real exchange rate is a random walk process, i.e. $\varrho = 1$. The alternative hypothesis is that ϱ is less than 1 in absolute value.[6] In order to separate out the unobservable γ from the actual rate R,

[5] For instance, "...this restoring of the equilibrium may take a long time, especially if the forces which keep the rate down are powerful and are continually at work." Cassel (1922, pp.158).

[6] The case with $-1 < \varrho < 0$, implies cyclical convergence towards $\overline{R}$, which seems implausible. This because it implies a systematic overshooting and undershooting of R relative to $\overline{R}$, as in the well known cobweb model.

equation (4.1) is rephrased in the ADF framework as follow, with the constant term α defined as $\gamma(1-\varrho)$:

$$\Delta R_t = \alpha - (1-\varrho)R_{t-1} + \sum_{i=1}^{\mathrm{p}-1} \psi_i \Delta R_{t-i} + \varepsilon_t. \tag{4.2}$$

This equation was estimated by OLS for a value of p equal to 6, which appeared sufficiently high to avoid any structure, e.g. autocorrelation, in the residuals. The coefficients of ΔR_{t-4}, ΔR_{t-5} and also of ΔR_{t-2}, however, appeared to be insignificantly different from zero with p-values, 0.84, 0.27 and 0.28, respectively. Joint zero restrictions on these coefficients were accepted by a F-test at a p-value of 0.44. Table 4.1 sets out a parsimonious version of the general model, obtained by (sequential) omission of these terms.

The diagnostics of the estimated equation do not indicate systematic structure in the residuals, increasing the reliability of the coefficient estimates. Due to some outliers, however, the normality assumption is violated at the 1% level of significance, cf. Figure 4.5.

The null hypothesis of, $\varrho = 1$, or equivalently of $1-\varrho = 0$ is rejected at the 5% level when using the Dickey-Fuller critical values. Since both the ADF model (4.2) and the presumed data generating process (4.1) contain the same type of deterministic term, a constant, one may argue for the use of critical values from the normal distribution. West (1988) shows that the t-statistic, e.g. $(1-\widehat{\varrho})/SE(\widehat{\varrho})$, will be asymptotically normally distributed under the null hypothesis of unit root, if both the model employed to test for unit root and the data generating process contain same type of deterministic terms, constant term and/or trend. Accordingly, the observed t-value -3.057 should be compared with the 95% critical value of -1.95 from the normal distribution. Banerjee et al. (1993, pp. 105), however, recommend the use of Dickey-Fuller distribution in finite samples to avoid overrejection of a possibly true null hypothesis.

The outcome of the ADF-test is consistent with the PPP theory. The real exchange rate does not seem to be a random walk process, but an equilibrium reverting process. The derived estimate of the equilibrium level is $0.126/0.131 \approx 0.96$, see Table 4.1. The speed

Table 4.1: A univariate model of the real exchange rate.

$$\underset{(3.048)}{\Delta\widehat{R}_t = 0.126} - \underset{(-3.057)}{0.131}\ R_{t-1} - \underset{(1.824)}{0.183}\ \Delta R_{t-1} + \underset{(1.266)}{0.134}\ \Delta R_{t-3}$$

Sample: 1972(2) -1997(4), 103 Quarterly observations.

$t - ADF = -3.057$, DF-*Critical values:* $5\% = -2.887$, $1\% = -3.489$

Diagnostics

R^2	=	0.101
Standard error of residuals: $\widehat{\sigma}$	=	0.015
Durbin Watson statistic: DW	=	1.95
Autocorrelation 1-5: $F_{ar,1-5}(5, 94)$	=	0.51[0.77]
ARCH 5: $F_{arch,1-5}(5, 89)$	=	1.11[0.36]
Normality: $\chi^2_{nd}(2)$	=	9.1[0.011]*
Heteroscedasticity: $F_{het,\mathrm{Xi}^2}\ (6, 92)$	=	1.86[0.10]
Heteroscedasticity: $F_{het,\mathrm{XiXj}}\ , (9, 89)$	=	1.27[0.26]
Model specification: RESET $F(1, 98)$	=	0.13[0.72]

Note: $\mathrm{F}_{ar,1-5}(df1, df2)$ tests for autocorrelation in the residuals up to 5 lags. df1 and df2 denote degrees of freedom. $\mathrm{F}_{arch,1-5}$ (df1, df2) tests for autoregressive conditional heteroscedasticity (ARCH) up to order 5, see Engle (1982). The normality test with chi-square distribution is that by Jarque and Bera (1980). $\mathrm{F}_{het,XiXj}$ (df1,df2) and F_{het,Xi^2} (df1, df2) are tests for residual heteroscedasticity due to omission of cross products of regressors and/or squares of regressors, see White (1980). RESET F(df1, df2) is a regression specification test. It tests the null hypothesis of correct model specification against the alternative hypothesis of misspecification, indicated by the significance of the square of the fitted value of the dependent variable in the model, see Ramsey (1969). The results in his table are based on the implementation of these tests in PcGive 9.10, see Hendry and Doornik (1996). Here and elsewhere in this study, a raised star * indicates rejection of the null hypothesis at 5% level of significance, while two stars ** indicate rejection of the null hypothesis at 1% level of significance. Furthermore, p-values are shown in square brackets.

of adjustment towards the long run equilibrium is 0.131, which is relatively fast. The half life of a given deviation from the equilibrium real exchange rate is about 6 quarters, while about 2/3 of a deviation is eliminated within a period of 10 quarters, see Figure 4.4.[7] We elaborate on the adjustment process towards the long run equilibrium after examining the constancy of the equilibrium real exchange rate.

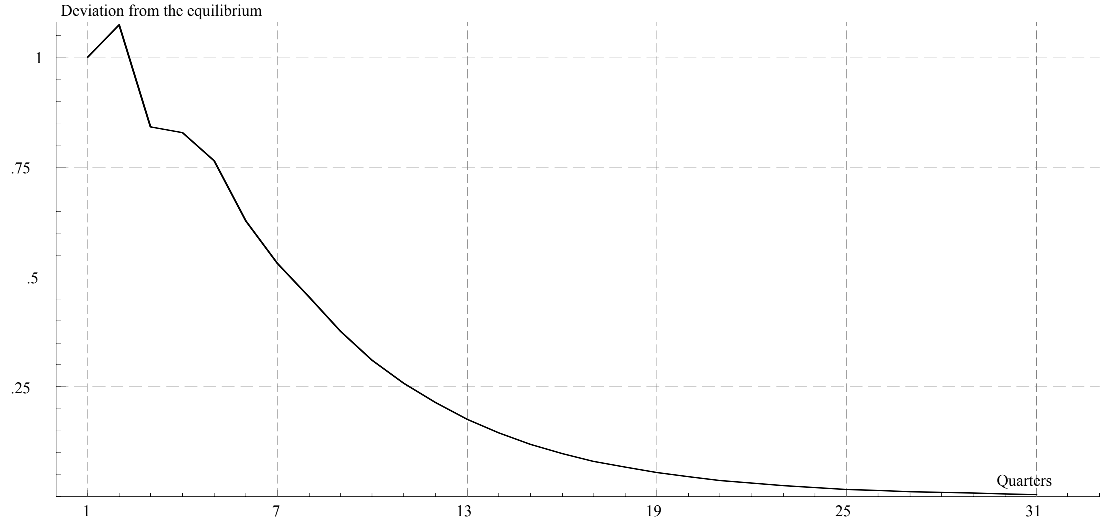


Figure 4.4: *Real exchange response when exposed to a shock of size 1 in quarter 1, when the real exchange rate is at its equilibrium level.*

4.3. Testing the constancy of the equilibrium real exchange rate

The real exchange rate model in Table 4.1 has been estimated recursively and subjected to a number of parameter constancy tests in order to evaluate the assumption of parameter constancy, in particular the constancy of γ. The upper panel of Figure 4.5 plots the recursive OLS estimates of α and $-(1-\varrho)$ together with ± 2 times their recursively estimated standard errors. The impression is of parameter stability, which is not contradicted by the Chow tests, used to test the overall stability of the model. However, due to an outlier in 1997, the 1-step up Chow test and the break point (Ndn) Chow tests indicate a parameter change in 1997.[8]

The relatively stable estimates of α and of $-(1-\varrho)$ imply that the recursively derived

[7] The estimate depends on whether one employs the commonly used formula for half life, i.e. $\ln(1/2)/\ln(\widehat{\varrho})$ where $\widehat{\varrho}$ is the estimated value of ϱ, or impulse response analysis, see e.g. Hallwood and MacDonald (1994, pp. 133). In the first case the half life is 4.94 or 5 quarters since $\widehat{\varrho} = -0.131 + 1 = 0.869$. However, if one takes into account the distribution of e.g. $\widehat{\varrho}$ on the different lags of R by impulse response analysis, the estimate of half life increases by 1.4 quarters in this case, see Figure 4.4. The figure is based on an AR(4) model for the real exchange rate, a reformulation of the model in Table 4.1.

[8] A dummy variable, *id97q1* with a value of 1 in 1997:1, -1 in 1997:2 and zero elsewhere was used to take account of the extreme values in 1997. This resulted in no-rejection of the parameter constancy assumption by the tests employed above, including the Chow tests, at the 5% level. Moreover, the dummy also helped to improve the model diagnostic, cf. Table 4.1. In particular, the normality assumption regarding the residuals was not rejected at 10% level. The t-ADF from this model was -3.147.

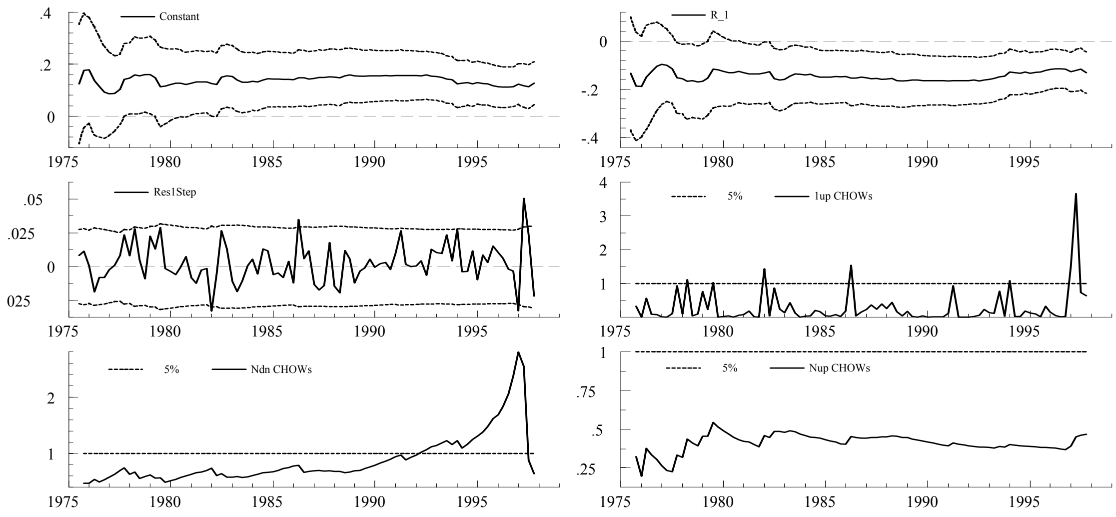


Figure 4.5: *Recursive statistics for the ADF model in table 4.1. Top panel: recursive estimates of the constant term* (α) $\pm 2SE$ *and of -(1-*ϱ*)* $\pm 2SE$*. Middle panel: 1-step ahead residuals* $\pm 2SE$ *and 1-step ahead Chow test statistics normalized by the critical value at 5% level of significance. Bottom panel: Break point Chow tests and N-step ahead Chow tests. Both test statistics are normalized by the accosiated critical values at 5% level.*

estimates of γ, the equilibrium real exchange rate, are stable too. Figure 4.6 shows the estimated values of $R\_Eq \pm 2SE$, where R_Eq is the estimated value of γ, over the period 1976:1 - 1997:4.[9] The standard errors have been (recursively) estimated by following the procedure suggested in Bårdsen (1989). Figure 4.6 also displays the actual real exchange rate, R, in the same period.

The estimates of γ are remarkably stable around 0.95 over the period 1978:1-1997:4. The γ-estimates before 1978 are relatively less stable, but then they are based on quite few observations. This is also reflected in the relatively large standard errors of the estimates from this period. If we disregard these, an estimate of γ at time $t+n$ where $n > 0$, remains well inside the $\pm 2SE$ band of a $\gamma-$estimate at any time t from 1978 and onwards.

[9] In fact, R_Eq has been obtained by recursive estimation of an autoregressive model for R with 4 lags, i.e. AR(4), which is just a reparameterisation of the model in Table 4.1 with p = 4. This, for our convenience in deriving the corresponding (asymptotic) standard errors. As noted earlier, the AR(4) model estimated on the full sample was also employed in the impulse response analysis. Recursive estimates of γ obtained by $\widehat{\alpha}/(1-\widehat{\varrho})$ are slightly different from R_Eq during the early 1970s owing to few observations, elsewhere they are almost indistinguishable as expected. As suggested by the graphs for the recursive estimates of α and -$(1-\varrho)$, recursive estimates of γ obtained by $\widehat{\alpha}/(1-\widehat{\varrho})$ are relatively more stable than R_Eq in the mid 1970s.

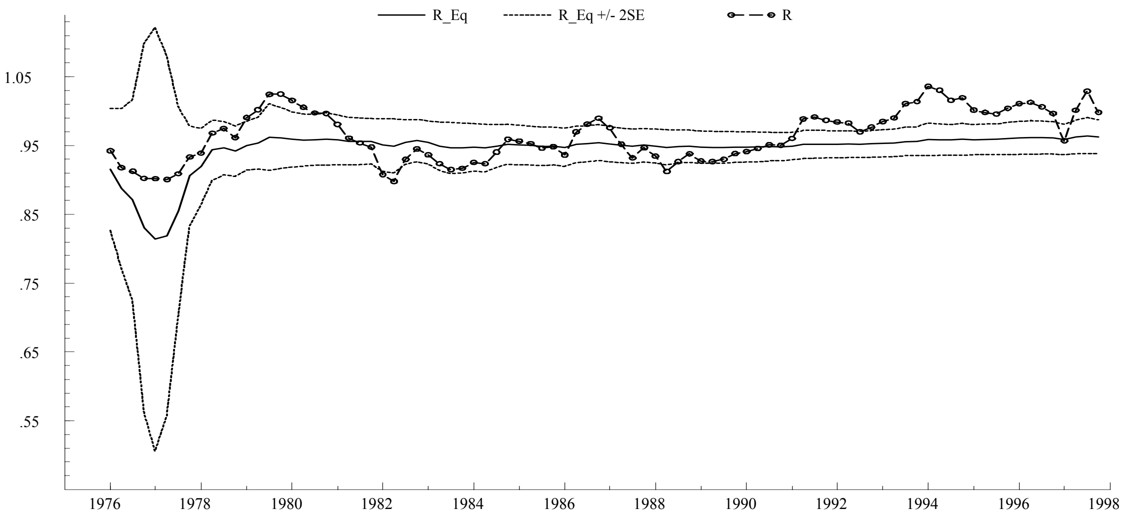


Figure 4.6: *Solid line shows recursive estimates of the equilibrium real exchange rate, γ, here denoted by R_ Eq. The recursive estimates have been derived for period 1975:2-1997:4. The initial estimate is based on 14 observations from period 1972:2-1975:4. The dashed lines represents the 95% confidence interval for R_ Eq. The dashed line with circles shows the actual real exchange rate, R, in this period.*

It follows that the constancy of γ cannot be rejected at the 5% level in this period.

Fluctuations of the (actual) real exchange rate, R, against the estimated equilibrium level in Figure 4.6, provide more information about how fast a given deviation from the equilibrium level is eliminated. In the figure, R fluctuates around the estimated equilibrium rate until about 1990. In this period, most of the deviations from equilibrium seem to be eliminated within 1-2 years. The exception is the late 1970s when the deviation lasts about 3.5 years. R is above its equilibrium rate, i.e. at a depreciated level relative to the equilibrium level, during this period. This may partly owe to a devaluation of the krone by 8% within the “snake” in February 1978 and thereafter, avoidance of a revaluation of about 2-4 % within the snake in October 1978, see Norges Bank (1987) and Section 3. The wage and price freeze from September 1978 to December 1979 appears to be another factor behind the real exchange rate depreciation during this period.

During the 1990s, R is again mostly above the estimated equilibrium level. The deviation is abruptly eliminated by the relatively large nominal appreciation in 1997:1, but R returns to its previous level shortly afterwards, see Figure 4.3. If we disregard this single quarter, the 1990s can be defined as a 7 years period with R above γ.

The persistent real exchange rate depreciation relative to the equilibrium rate during the 1990s may be related to the no-devaluation stance since the late 1980s and/or the "solidarity alternative" since 1992, according to which nominal exchange rate stability is offered for wage and price moderation, see Section 3. A value of $R > \gamma$ suggests that the real exchange rate is likely to appreciate. This can occur through nominal exchange rate appreciation and/or by a rise in CPI/CPI^f. The "solidarity alternative" appears to inhibit the real appreciation by obstructing both of these channels. Even if the wage and price moderation due to this agreement is insignificant, as argued by e.g. Evjen and Nymoen (1997), a stable exchange rate policy entails that the real appreciation has to occur mainly by way of a rise in CPI/CPI^f, which takes a longer time than nominal appreciation in general.

The coincidence of the relatively long period of real exchange rate depreciation with the no devaluation stance and the "solidarity alternative", also lends support to the proposition that the relatively fast adjustment toward the equilibrium before the 1990s, in particular before the late 1980s is to a large extent the result of frequent devaluations until the middle of 1986, cf. Rødseth and Holden (1990) and Calmfors (1990).[10] This is especially true, if the "solidarity alternative" has itself been ineffective in bringing about a wage and price moderation, i.e. the wage and price processes have remained as in the 1970s and 1980s.

Summary and conclusions: Both the graphical analysis and a formal test conducted within a univariate model are consistent with the PPP theory. The test suggests that real exchange rate is a stationary process which converges towards its equilibrium level. Tests for parameter stability, in particular recursive estimates of the equilibrium real exchange rate have not revealed significant changes in the equilibrium rate over time, at least not over a 20 years period from 1978:1 to 1997:4. Theories that imply a change in the equilibrium

[10] A stable exchange rate policy under the prospects of nominal appreciation is also classified as an "exchange rate protection" policy, i.e. competitiveness preserving policy, see section 2.3 and e.g. Corden (1995). This may partly explain why it has been supported by the Norwegian labour and employers unions. Furthermore, why the monetary authorities, whose policies have traditionally been guided by the objective of preserving the competitiveness of the economy, have not needed to retreat from the no-devaluation stance, see the discussion in Section 3.

real exchange rate in the face of a discovery or a revaluation of natural resources, fail to explain this finding.

Deviations from the equilibrium rate are eliminated relatively fast. The half life is about 1.5 years, which is less than the consensus estimates from the PPP literature that vary from 2.5 to 6 years for industrial countries. The measure of half life has been shown to disguise fairly large variation in the periods of deviations from the equilibrium rate, however; periods of deviations from the equilibrium rate has varied in the approximate range of 1-7 years. Since comparable results are not available in the literature, as it usually focuses on the half life measure, it is not possible to assess how typical the observed range is.

It is also difficult to assess the exact contribution of each of the factors that may have played a role in bringing about a rapid adjustment. Yet, if we confine ourselves to the 20 years period for which we have more certain results, our impression is, that the fast adjustment mainly owes to the active use of devaluations until the middle of 1986. This is consistent with the persistent real exchange rate depreciation during the 1990s, when no devaluation did take place. Centralised bargaining itself appear to be another explanation, since most a given deviation from the equilibrium is, on average, eliminated within a 1-2 years period, which coincides with the length of wage contracts. But, given the relative openness of the Norwegian economy, the arbitrage pressure is also likely to be relatively strong. Accordingly, the 1-2 years period can also be interpreted as the time lag in price setting due to menu costs. However, if the arbitrage pressure is assessed in the light of the weak empirical support for the law of one price, suggested by e.g. Goldberg and Knetter (1997), then centralised bargaining seems to have had a stronger influence on the speed of adjustment than arbitrage pressure, see Subsection 2.3.

Income policy, interpreted narrowly as changes in fiscal policy to influence the outcome of wage settlements, is unlikely to have played a considerable role in bringing about a rapid convergence towards the equilibrium. This because it has not been actively used during this 20 years period, at least not as often and explicit as before the late 1970s. However, attempts to control growth in wages and prices through legal actions in the late 1970s, or voluntarily through the "solidarity alternative" during the 1990s, have apparently kept the real exchange rate at a depreciated level relative to its equilibrium for a longer period than it would have been otherwise.

5. Testing PPP in a multivariate system framework

The PPP theory implies symmetry and proportionality restrictions on e.g. domestic and foreign prices in a long run nominal exchange rate equation, see equation (2.4). In the previous section, these restrictions were imposed by the definition of the real exchange rate. A number of studies have, however, questioned the plausibility of these restrictions in the light of e.g. possible variation in the construction of aggregate prices across countries and measurement errors, see among others Froot and Rogoff (1994). This criticism finds support in numerous empirical studies that report considerable deviations from both restrictions, see e.g. MacDonald (1995).

The present section employs the multivariate cointegration procedure of Johansen (1988) to test the symmetry and proportionality restrictions, see e.g. Juselius (1992) and Johansen and Juselius (1992). Single equation models as equation (2.4) have been popular in tests of the symmetry and proportionality restrictions, using either the nominal exchange rate or domestic prices as the left hand side variable. Such tests are however potentially biased if conditioning on the right hand side variables is invalid. Traditionally, the criticism has been formulated in terms of a potential simultaneity bias and the suggested remedy has been to use more appropriate estimation methods, maintaining a single equation framework, see e.g. Krugman (1978). A single equation framework is however only justified for the purpose of drawing inference on parameters of interest if the right hand side variables are weakly exogenous for the parameters of interest, see Engle et al. (1983). Hence, joint modelling of the variables is required to avoid biased inference when this assumption is violated.

The multivariate cointegration procedure is employed within vector autoregressive (VAR) models that not only treats the nominal exchange rate and domestic and foreign consumer prices as endogenous variables, but also a number of other variables that are commonly regarded as important determinants of nominal and real exchange rates, in particular the oil price in our case. A well specified VAR model enables one to draw valid inference on e.g. the long run effects of a variable and, in addition, offers a convenient framework to test whether valid inference is feasible by conditioning on a subset of variables, see Engle et al. (1983) and Johansen (1995b, ch. 8). We exploit both of these capabilities in the following tasks:

Firstly, we examine the long run relation between the nominal exchange rate and the prices, and test for long run effects of the additional variables on these, especially of oil prices on the nominal exchange rate. The evidence in favour of a constant equilibrium real exchange rate in the previous section suggests that the additional variables are unlikely to have long run effects on the real exchange rate. Tests for their effects on the nominal exchange rate and the prices are not only interesting in their own right, but may also shed light on whether they indirectly affect the real exchange rate through the nominal exchange rate and the prices. For instance, one cannot exclude the possibility that e.g. oil prices have an appreciating effect on the nominal exchange rate which in the long run is cancelled out by a rise in the relative price, CPI^f/CPI, brought about by higher oil prices, leaving the equilibrium real exchange rate constant. If so, oil prices may have considerable influence on the real exchange rate in the short and medium run.

Second, we investigate whether, and to what extent, the consumer prices and the nominal exchange rate adjusts to a deviation from the parity, provided that it holds. The graphical analysis in the previous section has revealed that both the relative price, CPI^f/CPI, and the nominal exchange rate tend to adjust when there is a deviation from the equilibrium real exchange rate. This behaviour can be analysed more rigorously in the VAR framework by testing whether the consumer prices and the nominal exchange rate are weakly exogenous for the parameters defining the long run relation between them. The outcome of these tests may throw light on the speed of adjustment towards the long run relation. In particular, it can provide information on the relative contribution of the nominal exchange rate and the domestic prices in bringing about adjustment towards the long run, although it will not be possible to identify the partial contribution of the factors discussed in Subsection 2.3. And secondly, the outcome of the tests may suggest whether shifts in the nominal exchange rate process is likely to affect the long run behaviour of the domestic prices, i.e. whether the nominal exchange rate is super exogenous for the parameters determining the long run behaviour of the prices.[11]

The next subsection motivates the choice of the variables to be included in the information set and examines their time series properties. Subsection 5.2 discusses whether it is necessary to conduct the analysis in a full VAR model or whether a conditional VAR model may be employed and, based on this discussion, it offers a guide to the subsequent analysis.

5.1. Choice of variables

The choice of variables is mainly motivated by earlier studies of PPP, e.g. Juselius (1992) and Johansen and Juselius (1992). In addition to the (natural) log of E, CPI and CPI^f, which have already been defined, the following variables are included in the information set: the log of the oil price (*oilp)*, the domestic bond rate (RB), the trade weighted government bond rate (RB^f) and a measure of net financial investment abroad relative to growth in the Norwegian GDP, *FI.Y,* see Appendix B for precise definitions.[12] The data set consists of seasonally non-adjusted quarterly observations and the sample covers the period from 1971:1 to 1997:4.[13] The estimation period however starts in 1972:2 owing to lags in the variables. The variables, including the level of the crude oil price in US dollars ($OILP$), are plotted in Figures 4.1, 4.2, 5.1 and 5.2.

The domestic bond rate was chosen in preference to the money market rates as the latter display quite erratic behaviour until the end of the 1970s, probably because of a thin domestic money market and regulations of international capital flows, see Figures 5.1 and 5.2. The government bond rate of Norway's trading partners is used as the Norwegian exchange rate was stabilised against the currencies of a majority of these countries, see Section 3. This policy entails a close link between the domestic and foreign interest rates under free international capital movements. Figure 5.1, however, indicates large gaps

[11] Weak exogeneity will be defined more precisely in the main text. A variable is super exogenous for parameters of interest if: (i) it is weakly exogenous and (ii), all parameters in the conditional model are invariant to changes in the process for the weakly exogenous variable, see Engle et al. (1983).

[12] Log transformation of E, *CPI,* CPI^f and $OILP$ will make it possible to interpret all the coefficient estimates as elasticities. Note that RB, RB^f and $FI.Y$ are measured in rates, hence they do not require log transformation.

[13] Initially, RB^f, was not readibly available from 1971:1 but from 1972:1. To avoid losing observations when modelling, it was extended backwards with the same observations as in 1972, for convenience.

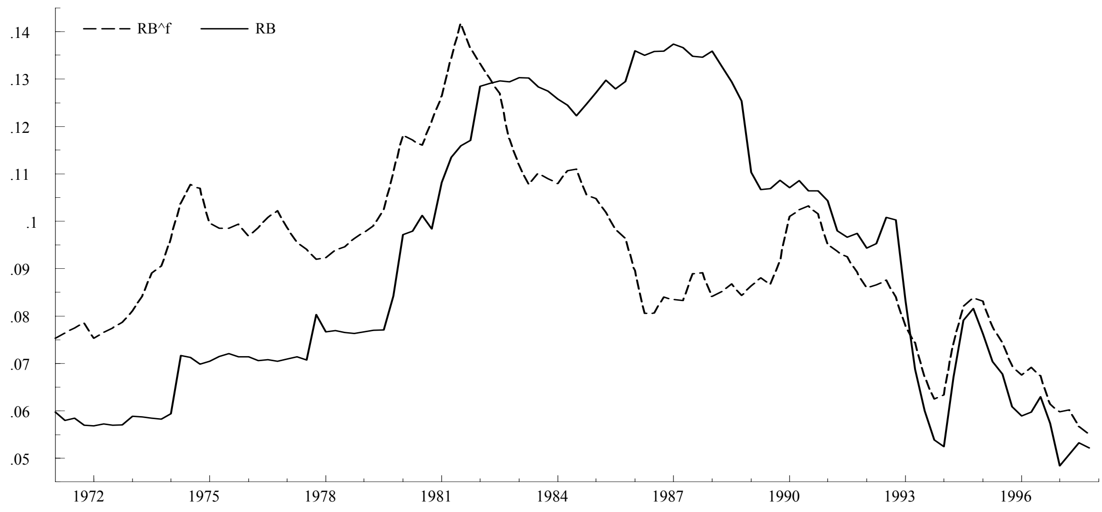


Figure 5.1: *Interest rates on government bonds in Norway and in its trading partners, RB and RB^f, respectively. The period is 1971:1-1997:4.*

between these interest rates in the 1970s and 1980s. These can be partly ascribed to the Norwegian capital regulations which were not fully dismantled before 1990 and were relatively tight during the 1970s, but may also reflect devaluation expectations and risk premia due to the frequent devaluations in this period.

The oil price is included to test for long run effects of oil prices on the exchange rate.[14] Oil prices, closely linked to gas prices, offer an indication of the oil and gas wealth and of their revenues on a regular basis. As shown in Skånland (1988), current oil prices tend to influence the estimates of future oil prices and thereby the estimates of permanent income from the oil and gas, which are not available on a regular basis. Thus a possible effect of oil prices on the nominal exchange rate and/or on prices can indicate whether revaluations of petroleum wealth and changes in the permanent income have affected the long run real exchange rate. Note that oil prices, as well as *FI.Y* (to be discussed below), are expected to have short run effects on the real exchange rate within the PPP framework, see Subsection 2.2. A number of authors have argued, however, that e.g. oil prices also affect the Norwegian nominal and real equilibrium exchange rates, see e.g. Alexander et al.

[14]This is in contrast to e.g. Johansen and Juselius (1992) where only relative changes in the oil price, $\Delta oilp$, are included.

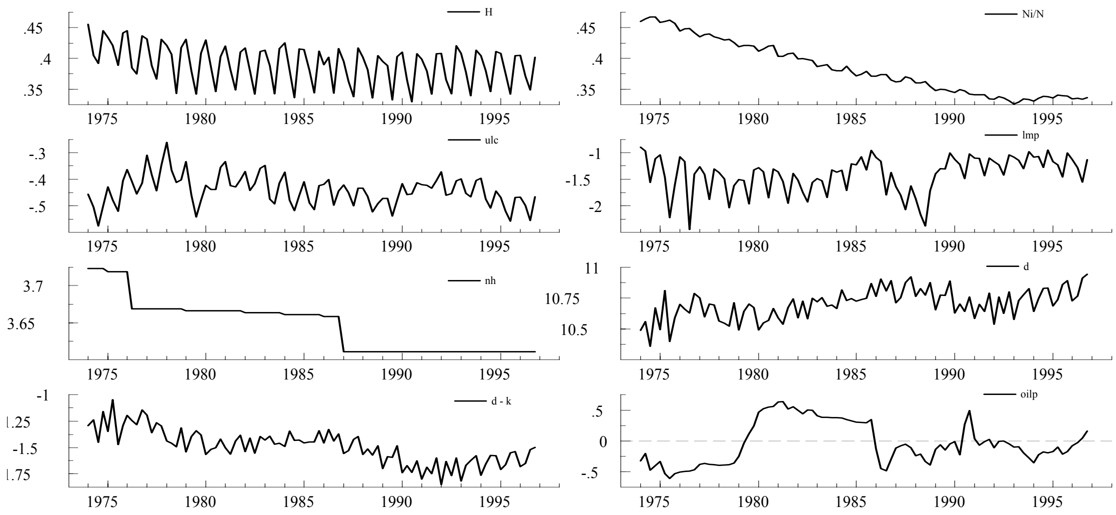


Figure 5.2: *From top left: (a) Three months money market rate in Norway, RS. (b) Price of Brent Blend crude oil in US dollars, OILP. (c) Natural log of OILP, oilp. (d) Foreigners financial investment in Norway scaled by the Norwegian GDP, FI.Y. The period is 1971:1-1997:4.*

(1997) and Haldane (1997). An assessment of long run oil price effects on the exchange rate can provide direct evidence on this issue.

Foreign net financial investment in Norway relative to growth in GDP ($FI.Y$), i.e. the negative of net financial investment abroad relative to growth in GDP, is another measure that can be used to assess the possible effects of changes in the petroleum wealth and revenues on the nominal and real exchange rates. As noted in Section 3, government savings made possible by the oil revenues have been used to acquire financial assets abroad. This investment, or saving, may reflect current and perceived future revenues from the petroleum sector and future import bills. It can thereby offer an indication of the required size of the sector for non-oil tradables, since import expenses not covered by future oil revenues and returns on assets abroad have to be met by the non-oil sector for tradables. Movements in $FI.Y$ may therefore influence the nominal and real exchange rates. Its relevance for the exchange rates is also suggested by the effects of capital flows on the (relative) demand and supply for the domestic currency, see e.g. Gibson (1996).

Productivity differentials and government spendings are omitted from the analysis to keep the analysis of the long run relations tractable. These variables have received mixed support in studies of PPP between developed countries, including those based on Norwegian data, see Froot and Rogoff (1994), Akram (2000b), Samiei (1998) and Edison and Klovland (1987). In Subsection 4.3, we found the long run real exchange rate to be remarkably constant over most of the sample period. This finding indicates that these variables are unlikely to have had long run effects on the nominal and real exchange rates.

Finally, it should be noted that the chosen set of variables is a reflection of one of the main objectives of this essay, namely to examine the short and long run behaviour of the Norwegian nominal exchange rate. The chosen set of variables is therefore likely to enable a more satisfactory characterisation of the nominal exchange rate behaviour than that of the other variables. However, it may still be adequate to sketch the long run behaviour of the other variables or, at least, illuminate aspects of it that have a bearing on our methods and results. The discussion on the choice of a full VAR model versus a conditional VAR model elaborates on this issue.

Time series properties

The time series properties of the chosen variables are tested to avoid combining variables with different order of integration which can lead to spurious relations, see Banerjee et al. (1993, ch. 3). Information about the degree of integration can also be useful when identifying possible long run relations between the variables.

Table 5.1 sets out the results from ADF tests of the order of integration. ADF models similar to model (4.2) were formulated, but with a trend term, t, in addition to the constant term. The trend was included to allow the null hypothesis of a stochastic trend to be tested against the alternative of a deterministic trend.

The ADF tests were conducted by formulating models with 8 lags of a dependent variable and sequentially omitting lagged terms that appeared to be insignificant, say at 20% level, since attention was also paid to the diagnostics of each of the ADF-models. Note that too few lags may result in overrejection of the null hypothesis when it is true, while too many may lead to underrejection of the null hypothesis when it is false, see e.g. Banerjee et al. (1993, ch. 4). Column 4 of Table 5.1 shows the highest retained lag, p,

Table 5.1: ADF tests for unit root, constant and trend included.

Variables	$1-\widehat{\varrho}$	t-ADF	ADF(p)	k
e	-0.094	-3.169	1	4
Δe	-0.815	-8.128**	0	3
cpi	-0.001	-0.111	5	7
Δcpi	-0.521	-4.775**	4	5
cpi^f	-0.006	-1.371	8	11
Δcpi^f	-0.351	-3.864*	7	9
RB	-0.018	-1.151	8	7
ΔRB	-0.724	-6.573**	6	6
RB^f	-0.057	-2.623	3	6
ΔRB^f	-0.686	-6.748**	3	5
$FI.Y$	-0.619	-5.133**	7	7
$oilp$	-0.079	-2.424	5	5
$oilp^a$	-0.170	-3.850*	5	5
$OILP$	-0.059	-1.733	5	6
$OILP^a$	-0.085	-2.195	5	6

Note: Dickey -Fuller critical values: 5% = -3.457, 1% = -4.057. Constant and trend included.Sample 1972:2-1997:4. [a]When using sample 1974:1-1997:4. p denotes the largest significant lag and k denotes the number of regressors.

while column 5 indicates the number of regressors, inclusive a constant and a trend. The data sample covered the period 1972:2-1997:4, except for *oilp* and *OILP* whose time series properties were additionally tested on data from the period 1974:1 to 1997:4, to control for the effect of OPEC I.

Table 5.1 indicates that the level of *e, cpi,* cpi^f, RB and RB^f are integrated of order 1 since their first differences seem to be integrated of order zero (stationary).[15] *FI.Y* appears to be a stationary variable, see also Figure 5.2. The evidence on the oil price is mixed. The initial test suggests that *oilp* is an integrated variable. The oil price series is however notorious for breaks, associated with e.g. OPEC I, OPEC II and the Gulf War, see Figure 5.2. Studies show that the ADF test suffers from low power when there are breaks in a series, see e.g. Perron (1989). Indeed, a number of studies argue that the oil price is not an integrated variable once infrequent changes in its mean are allowed for, see

[15]Note that stationary inflation rates, Δcpi and Δcpi^f, and non-stationary nominal interest rate, RB and RB^f, imply non-stationary real interest rates. This is not consistent with the Fisher hypothesis which implies a stationary real interest rate. A number of studies, however, regard the inflation rate to be integrated of order 1. But then order of integration is not an inherent property of a time series, cf. Hendry (1995). A time series can display a non-stationary behaviour in one sample and a stationary behaviour in another sample.

e.g. Perron (1989), Green et al. (1996) and Horsnell and Mabro (1993, ch. 11).

When *oilp* is tested on a shorter sample, starting from 1974:1 instead of 1972:2, to pass over the low oil prices in the pre-OPEC I period, the ADF test rejects the null hypothesis of integratedness at the 5% level, even though the other shocks are not accounted for. In the case of *OILP,* however, the null hypothesis is not rejected even when the sample starts in 1974:1. In general, the time series properties of a variable carry over to a non-linear transformation of the variable, see Granger and Hallman (1991). One possible explanation for this inconsistency may be that abrupt changes in *OILP* are to some extent suppressed in its log transformation, *oilp*, making it "easier" for the ADF test to reject the null hypothesis. The decision on whether or not oil prices should be treated as a non-integrated variable (with infrequent changes in its mean) is left to tests that are conducted within the multivariate cointegration framework.

5.2. Choice of models

This subsection discusses the choice between a full vector autoregressive (VAR) model and a conditional VAR model for the purpose of drawing valid inference on the number of long run relations (cointegrating relations) between variables and on parameters of interest.[16] Parameters of interest define particular long run relations between variables and measure their response to deviations from these long run relations. A conditional VAR model, obtained by conditioning on some of the variables, can simplify the analysis and increase the power of tests owing to its parsimony relative to a full VAR model. The discussion focuses on the conditions for a variable to be regarded as a valid conditional variable, i.e. weakly exogenous for the parameters of interest, see Engle et al. (1983).

A full VAR model: A full VAR model of vector $X = (e,\ cpi,\ RB,\ FI.Y,\ cpi^f,\ RB^f,\ oilp)'$, can be formulated in a reduced form vector equilibrium correction model (VEqCM) as follows:

[16] A given number of variables are said to cointegrate, if e.g. they are integrated of order d, but a linear combination of them is integrated of order $d-1$. A cointegrating vector denotes a vector consisting of the weights/coefficients that define a cointegrating relation between the variables. A stationary variable is by itself a cointegrating relation.

This study interchangeably use "long run relation", "equilibrium relation" for "cointegrating relation".

$$\Delta X_t = \alpha\beta' X_{t-1} + \sum_{i=1}^{\text{p}-1} \Gamma_i \Delta X_{t-i} + \delta + \varepsilon_t, \tag{5.1}$$

where

$$\alpha\beta' = \begin{bmatrix} \alpha_{e,\,1} & \alpha_{e,\,2} & . & . & \alpha_{e,\,r} \\ \alpha_{cpi,\,1} & \alpha_{cpi,\,2} & . & . & \alpha_{cpi,\,r} \\ \alpha_{RB,\,1} & . & . & . & \alpha_{RB,\,r} \\ \alpha_{FI.Y,\,1} & . & . & . & . \\ \alpha_{cpi^f,\,1} & . & . & . & . \\ \alpha_{RB^f,\,1} & . & . & . & . \\ \alpha_{oilp,\,1} & . & . & . & \alpha_{oilp,\,r} \end{bmatrix}_{7 \text{ x } r} \begin{bmatrix} \beta_{e,\,1} & \beta_{cpi,\,1} & . & . & \beta_{oilp,\,1} \\ \beta_{e,\,2} & \beta_{cpi,\,2} & . & . & \beta_{oilp,\,2} \\ \beta_{e,\,3} & . & . & . & \beta_{oilp,\,3} \\ . & . & . & . & . \\ . & . & . & . & . \\ \beta_{e,\,\mathrm{r}} & \beta_{cpi,\,\mathrm{r}} & . & . & \beta_{oilp,\,r} \end{bmatrix}_{r \text{ x } 7} \tag{5.2}$$

The VEqCM form is chosen to highlight the long run relations governing the variables, $\beta' X$, and their speed of adjustment towards these long run relations, α. β', the transpose of β, is a r x 7 matrix of coefficients which defines the r linearly independent long run relations between the seven variables in X, where $r \leq 7 - 1 = 6$, since all variables are not stationary. Element $\beta_{j,\,i}$ of β' denotes the weight/coefficient of variable j in the ith long run relation, $j = e$, *cpi, RB, FI.Y,* cpi^f, RB^f and *oilp,* and i = *1, 2, ...r*. Deviations from the long run relations in a given period are (partially) corrected/adjusted in the subsequent period. Rates of adjustments are contained in the 7 x r matrix α. Element $\alpha_{j,\,i}$ of α denotes the adjustment rate of variable j to deviation from the ith long run relation in period $t{-}1$, . p is the maximum lag length, δ is a vector of constant terms and ε_t is a vector of residuals. Γ_i s' are coefficient vectors representing short run effects of the variables.

The number of cointegrating relations, r, can be determined by following the procedure suggested in e.g. Johansen (1988) and (1995b). Efficient inference on the basis of the Johansen procedure, however, requires that ε_t is IIDN(.), i.e. ε_t is independently and identically normally distributed, which generally presupposes a well specified model.

However, the information set X is unlikely to be sufficient to provide a reasonable approximation to the underlying processes determining all of the variables in X, in particular, cpi^f, RB^f and *oilp.* Increasing the lag length p can be remedial, but shocks and structural changes over time are likely to require extraneous information in the form of deterministic and/or stochastic stationary variables, cf. Johansen and Juselius (1992), among others. Extraneous stationary variables can, however, invalidate the asymptotic distributions of the relevant test statistics, which are derived by making allowance for a deterministic trend and a constant, at maximum, see e.g. Doornik et al. (1998) and Mosconi and Rahbek (1999).[17] Increasing the dimension of the VAR model by increasing the number of endogenous variables can be a solution, but with the possible drawback that identifying long run relations implied by economic theory may become more difficult.

The number of endogenous variables in X is already relatively high (7), so a reduction is desirable to make it less cumbersome to identify interpretable long run relations. In addition, the power of tests that are used to test theory restrictions on the β and α matrices may be low within a full VAR model, given the number of parameters to be estimated on our data sample of about 100 observations; in other words, overparameterisation may lead to the appearance of transient correlation between unrelated variables.

A conditional VAR model: The analysis can be simplified by conditioning on a number of variables, especially those that can be difficult to model using the given information set, and conduct the analysis within a conditional VAR model. A conditional model can, by reducing the dimension of the VAR, increase the power of tests and simplify the analysis in the light of economic theory. A valid conditional model, however, presupposes weak exogeneity of the conditioning variables for parameters of interest, see Engle et al. (1983) and Johansen (1995b, ch. 8). Otherwise, a conditional model can lead to inefficient or inconsistent inference on the parameters of interest.

The full VAR model (5.1) can be alternatively written as follows:

[17] The DisCo programme of Johansen and Nielsen (1993) can be employed when the analysis is conditioned on deterministic variables but not if conditioned on stochastic (stationary) variables.

$$\Delta Y_t = \alpha_1 \beta' X_{t-1} + \sum_{i=1}^{\mathrm{p}-1} \Gamma_{1,\, i} \Delta X_{t-i} + \delta_1 + \varepsilon_{1,\, t} \tag{5.3}$$

$$\Delta Z_t = \alpha_2 \beta' X_{t-1} + \sum_{i=1}^{\mathrm{p}-1} \Gamma_{2,\, i} \Delta X_{t-i} + \delta_2 + \varepsilon_{2,\, t} \tag{5.4}$$

This form is based on the following partition into vectors and submatrices: $X' = (Y',\, Z')$, $\varepsilon' = (\varepsilon_1',\, \varepsilon_2')$, $\alpha' = (\alpha_1',\, \alpha_2')$, $\Gamma_i' = (\Gamma_{1,\, i}',\, \Gamma_{2,\, i}')$ and $\delta' = (\delta_1',\, \delta_2')$.

A conditional VAR model of ΔY on ΔZ can be formulated as:

$$\Delta Y_t = \omega \Delta Z_t + (\alpha_1 - \omega \alpha_2)\beta' X_{t-1} + \sum_{i=1}^{\mathrm{p}-1} (\Gamma_{1,\, i} - \omega \Gamma_{2,\, i}) \Delta X_{t-i} + (\delta_1 - \omega \delta_2) + (\varepsilon_{1,\, t} - \omega \varepsilon_{2,\, t}) \tag{5.5}$$

Parameter vector ω is a function of the covariance matrix of residual vectors $\varepsilon_{1,\, t}$ and $\varepsilon_{2,\, t}$ and the variance matrix of $\varepsilon_{2,\, t}$.

If $\alpha_2 = 0$, a condition for the weak exogeneity of Z for β and α_1, valid inference on β and α_1 can be based on the conditional model (5.5) alone without loss of information relative to the full VAR model (5.1), see Johansen (1995b, pp. 122). For example, consistent estimation of α_1 based on the conditional model requires $\alpha_2 = 0$, even if β is known but $\omega \neq 0$. In particular, one may falsely draw the conclusion that Y does not respond to deviation from a long run relation, if $\alpha_1 \neq 0$ but $\alpha_1 = \omega \alpha_2$.

Since the Norwegian economy is small relative to its trading partners' one would *a priori* believe that cpi^f, RB^f and *oilp* do not respond to disequilibria in different Norwegian markets. This entails the following classification of the variables in X into endogenous and presumedly weakly exogenous variables: $Y = (e,\, cpi,\, RB,\, FI.Y)'$ while $Z = (cpi^f,\, RB^f,\, oilp)'$. Provided that this is a valid classification and the conditional model is well specified, a prerequisite for the residuals $\varepsilon_{c,\, t} \equiv \varepsilon_{1,\, t} - \omega \varepsilon_{2,\, t}$ to be IIDN(.), one can draw inference on r, α_1 and on the identity of the different cointegrating vectors in β by following Harbo et al. (1998).

Stationary oil prices imply $\alpha_2 \neq 0$: The possible stationarity of *oilp*, however, poses a problem for conducting the analysis within the conditional VAR model. The suggested classification of variables into Y and Z implies that α_2 can be defined as

$$\alpha_2 = \begin{bmatrix} \alpha_{cpi^f,\,1} & \alpha_{cpi^f,\,2} & . & . & \alpha_{cpi^f,\,r} \\ \alpha_{RB^f,\,1} & . & . & . & . \\ \alpha_{oilp,\,1} & . & . & . & \alpha_{oilp,\,r} \end{bmatrix}_{3\text{x}r} \tag{5.6}$$

The remaining elements of α can be defined as α_1. If *oilp* is a stationary variable then it is, by itself, a cointegrating term. Hence one of the rows, for instance row k, in β' may contain only one non zero coefficient associated with the $oilp_{t-1}$ in vector X_{t-1}, i.e. it would be sufficient that only $\beta_{oilp,\,k} \neq 0$ for row k to define a stationary relation. In which case, $\Delta oilp_t$ (in ΔZ_t) will respond to changes in $oilp_{t-1}$, implying $\alpha_{oilp,\,k} \neq 0$ and hence, $\alpha_2 \neq 0$. If $oilp_{t-1}$ simultaneously appears in the equations for the endogenous variables Y, then *oilp* is not a weakly exogenous variable for $\beta_{oilp,\,k}$ and for the coefficients representing the response of the endogenous variables to $oilp_{t-1}$, $\alpha_{j,\,k} \in \alpha_1$ where $j \in Y$. Moreover, the possibility of $oilp_{t-1}$ appearing in the equations for Δcpi^f and ΔRB^f cannot be neglected, e.g. in the light of the stagflation experience of industrialised countries following OPEC I, which also implies $\alpha_2 \neq 0$. Thus (even) cpi^f and/or RB^f will not be weakly exogenous for $\beta_{oilp,\,k}$ and $\alpha_{j,\,k}$, where $j \in Y$. Consequently, inference on the long run effects of oil prices on the endogenous variables, based on the conditional model alone, can be invalid.

Valid inference feasible on a subset of parameters: Weak exogeneity of a given variable is, however, defined relative to particular parameters of interest. Thus the failure of cpi^f, RB^f and *oilp* to be weakly exogenous for e.g. $\beta_{oilp,\,k}$ and $\alpha_{j,\,k}$, does not preclude that they can be treated as weakly exogenous for other cointegrating vectors in β' and speeds of adjustment in α_1.

The case of a stationary conditional variable resembles the case of cointegrating relation(s) between a set of conditional variables. The matrix β' can be partitioned as β' = (β'_y, β'_z), where β'_y denotes the $r_1 \leq r$ cointegrating relation(s) that either involve both the Y and the Z variables or only the Y variables, while β'_z denotes $r - r_1$ cointegrating relations composed exclusively of Z variables. The adjustment matrices can be partitioned as $\alpha_1' = (\alpha_{yy}', \alpha_{yz}')$ and $\alpha_2' = (\alpha_{zz}', \alpha_{zy}')$, where α_1 contains the weights of $\beta_y' X_{t-1}$ and $\beta_z' Z_{t-1}$ in the Y equation while α_2 contains their weights in the Z equation. A model of ΔY conditional on ΔZ and a marginal model of ΔZ can be formulated as:

$$\begin{aligned}
\Delta Y_t &= \omega \Delta Z_t + (\alpha_{yy} - \omega \alpha_{zy}) \beta_y' X_{t-1} + (\alpha_{yz} - \omega \alpha_{zz}) \beta_z' Z_{t-1} + \qquad (5.7) \\
&\quad \sum_{i=1}^{\mathrm{p}-1} (\Gamma_{1,\,i} - \omega \Gamma_{2,\,i}) \Delta X_{t-i} + (\delta_1 - \omega \delta_2) + (\varepsilon_{1,\,t} - \omega \varepsilon_{2,\,t}) \\
\Delta Z_t &= \alpha_{zy} \beta_y' X_{t-1} + \alpha_{zz} \beta_z' Z_{t-1} + \sum_{i=1}^{\mathrm{p}-1} \Gamma_{2,\,i} \Delta X_{t-i} + \delta_2 + \varepsilon_{2,\,t} \qquad (5.8)
\end{aligned}$$

Analysis of the conditional model (5.7) is without loss of information relative to the VAR model for the purpose of drawing inference on β_y' and α_{yy}, if

$$(1)\ \alpha_{zy} = 0 \text{ and } (2)\ \alpha_{yz} - \omega \alpha_{zz} = 0,$$

see Hendry and Mizon (1993). Under these conditions, $\beta_y' X_{t-1}$ only enters the conditional model while $\beta_z' Z_{t-1}$ only enters the marginal model (5.8). For instance, if $\beta_z' Z$ only consists of *oilp*, valid inference on the other r_1 possible cointegrating relations and the associated adjustment coefficients requires that these r_1 cointegrating relations are absent from the equations for the Z-variables, $Z = (cpi^f,\ RB^f, oilp)'$, and that *oilp* is absent from the conditional model of the Y-variables, $Y = (e,\ cpi,\ RB,\ FI.Y)'$.

Summary and a guide to the subsequent analysis: The full VAR model defined by the chosen variables set is unlikely to provide a satisfactory characterisation of all the variables unless deterministic and non-modelled stochastic variables are included in the model. This extension may, however, invalidate the test statistics to be employed and make it difficult to derive interpretable results. The latter concern even makes it desirable to reduce the number of variables to be modelled. A conditional VAR model may be a solution since it allows us to condition on variables whose processes are not of primary interest, but only under a certain condition. If oil prices are stationary, this condition will not be satisfied for the purpose of drawing valid inference on the parameters of interest. In particular, the effects of oil prices on the endogenous variables may not be consistently estimated from the conditional model. Inference on a subset of parameters is yet feasible under a set of weaker conditions.

A compromise would be to employ the full VAR model to test for the number of cointegrating relations, make an effort to identify their form, test for long run effects of oil prices on the endogenous variables and test whether or not a subset of variables are weakly exogenous for the parameters of interest; thereafter, if condition (1) for the weak exogeneity is fulfilled for a subset of parameters, to employ a conditional VAR model and check whether condition (2) is satisfied; and if it is, to proceed by conducting tests for the (remaining) hypotheses of interest within the conditional VAR model and by testing robustness of the results obtained from the full VAR model.

The analysis in Subsection 5.3 and 5.4 is based on this compromise. The main objectives of the analysis are to test for: (a) the symmetry and proportionality restrictions implied by the PPP theory, (b) possible long run effects of oil prices and (c) whether or not both the nominal exchange rate and the prices adjust upon deviations from the PPP, provided that it holds.

In Subsection 5.3: A full VAR model is employed to determine the number of cointegrating relations, r, and to test condition (1), $\alpha_{zy} = 0$, for the possible weak exogeneity of the Z variables, cpi^f, RB^f and *oilp,* for specific vectors in β, i.e. β_y, and the associated adjustment coefficients. An effort is made to identify their form in the light of the PPP theory, the uncovered interest rate parity (UIP) hypothesis and the results from the ADF tests. β is restricted to test whether *oilp* itself constitutes a cointegrating relation by being a stationary variable. The test is not rejected and *oilp* appears to be the only cointegrating term among the Z variables. The model is then used to test for possible long run effects of *oilp* on the Y-variables, i.e. to test the null hypothesis of $\alpha_{yz} = 0$.

In Subsection 5.4: Assuming that condition (1) holds, a conditional VAR model is formulated with $Y = (e,\ cpi,\ RB,\ FI.Y)'$ and $Z = (cpi^f,\ RB^f,\ oilp)'$. The number of cointegrating relations r is taken from the VAR model. The model is employed to test the validity of conditioning on Z by checking if condition (2), $\alpha_{yz} - \omega\alpha_{zz} = 0$, is satisfied, which turns out to be the case. Thus *oilp* appears to be redundant in the conditional model of Y. A more parsimonious conditional model is derived by excluding *oilp* (and some of the associated deterministic variables) from the conditional model. The exclusion of *oilp* reduces the number of cointegrating relations by one, to $r-1$. The resulting parsimonious conditional model is used to check the robustness of the results obtained from the full VAR model and to test the weak exogeneity of *cpi*, *RB* and *FI.Y* for the parameters of interest in the nominal exchange rate equation.

5.3. The VAR model

The full VAR model is presented in Table 5.2. The information set is extended by five dummy variables to take account of the oil price shocks, OPEC I, OPEC II, Gulf War and the oil price fluctuations in 1997, see Figure 5.2. In order to keep the cointegration space tractable, only a step dummy for OPEC II (*OP2*) is allowed to enter the cointegration space. This takes on a value of 1 from 1979:1 to 1985:4 and zero elsewhere. The three other oil price shocks are controlled for by impulse dummies, *OP1*, *id90q3* and *id97q1*, respectively, and entered unrestricted. *OP1* takes on a value of 1 in 1974:1, -0.3 in 1974:2 and zero elsewhere, while *id90q3* and *id97q1* take on a value of 1 in the indicated quarter, -1 in the subsequent quarter and zero elsewhere, cf. Figure 5.2. The dummy variables for the oil price shocks are primarily needed to take account of significant changes in the oil price process since the 1970s. A time trend t is restricted to the cointegration space while the constant term is entered unrestricted, as recommended by Doornik et al. (1999) and Harbo et al. (1998) to safeguard against invalid inference on the cointegration rank, r. Estimation of the unrestricted reduced form indicated the need to include at least 5 lags, p = 5, and at least two centered seasonal dummies for the first and the third quarter, *CS* and *CS2*, respectively, to avoid autocorrelation and violation of the normality assumption for the residuals, see Appendix C.

The single equation and system diagnostic tests in Table 5.2 show that the chosen model formulation is able to ensure almost IIDN(.) residuals, both when evaluated at

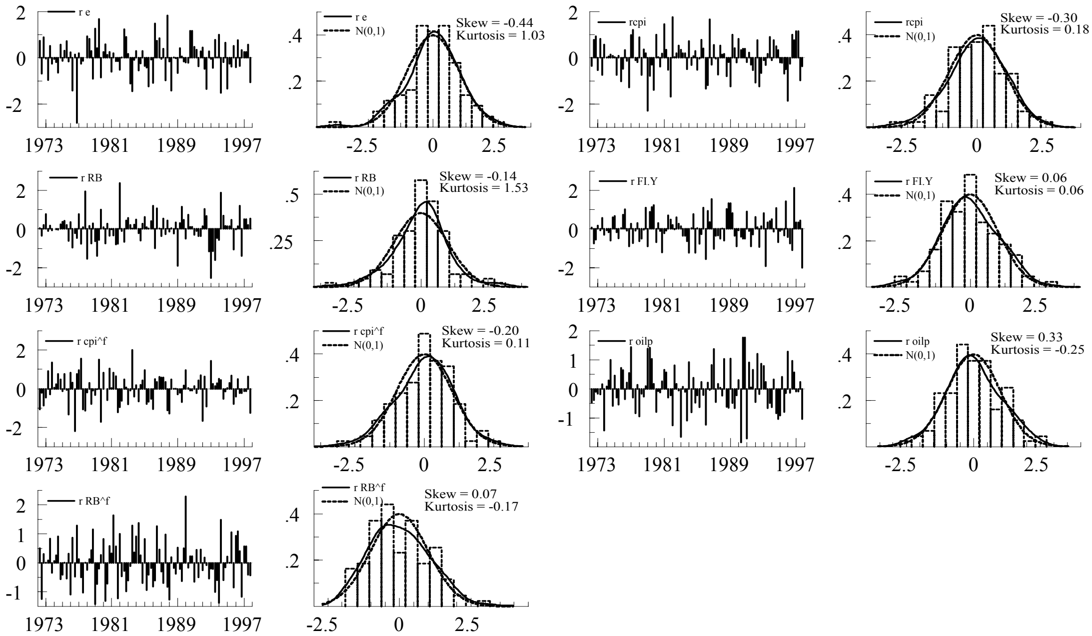


Figure 5.3: *Residual characteristics in the full VAR model. Scaled residuals (in the first and the third column) and the distribution of residuals (in the second and the fourth column) from each of the seven equations in the VAR model. The residual distributions are plotted against the standard normal distribution for comparison. The measures for skewness and excess kurtosis are also reported. The period is 1972:2-1997:4.*

the single equation level and at the system level. The diagnostics do not reject the null hypotheses of no autocorrelation up to 5 lags, ARCH type heteroscedasticity up to order 4 and the null hypotheses of normally distributed residuals at the 5% level. One exception is the domestic bond rate *(RB)* equation, whose residuals do not seem to be normally distributed even at the 1% level. Figure 5.3 suggests that the rejection of the normality hypothesis for the RB residuals mainly owes to excess kurtosis, the statistic is 1.53 whereas it should be close to zero under the null hypothesis of normality. There are also a few relatively large outliers that can be associated with the increase in the Norwegian bond rates in 1982 and with the ERM crisis in 1992, but these seem to matter less, see Figure 5.1. The distribution of the residuals from the nominal exchange rate equation, e-residuals, also appears to suffer from excess kurtosis and from one outlier in 1976/77. The normality assumption for e-residuals is however not rejected at the 1% level; the associated p-value

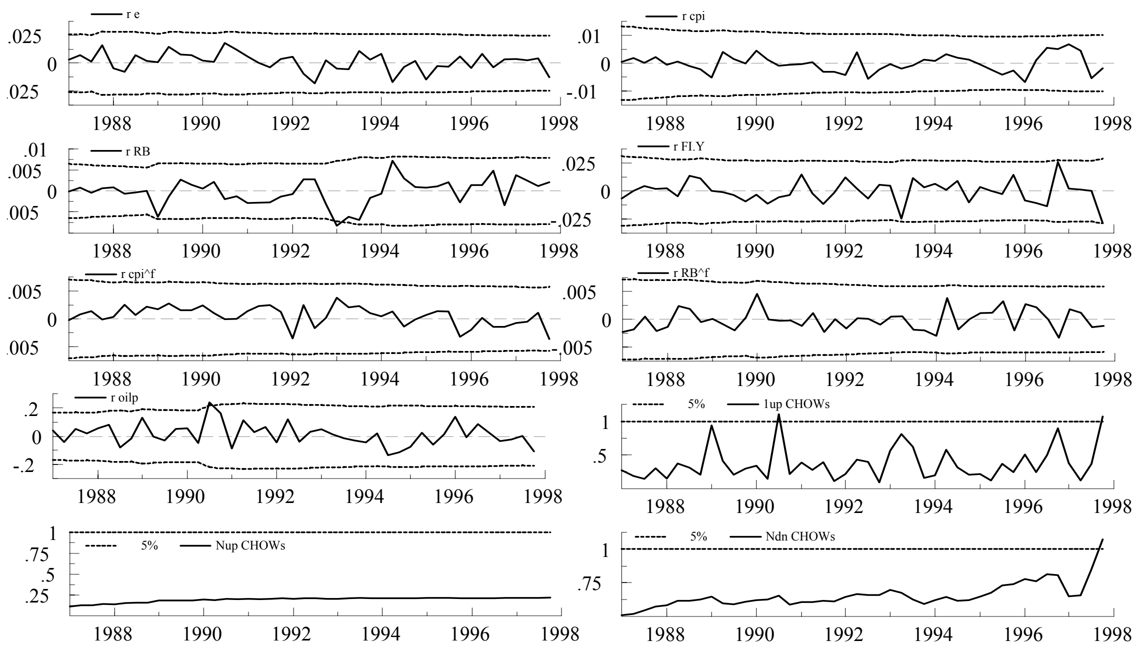


Figure 5.4: *Constancy statistics for the VAR model, obtained by recursive estimation of the VAR model over period 1986:4-1997:4. One-step residuals* $\pm 2SE_t$ *for each equation in the VAR model. One-step ahead Chow statistics (1up Chows), N-step ahead Chow statistics (Nup Chows) and break point Chows (Ndn Chows) for the VAR model. The Chow statistics are scaled by their (one-off) critical values at the 5% level of significance.*

is 4%.

The VAR model was also estimated without the dummy variables and a comparison of the standard errors of residuals with and without dummy variables, $\widehat{\sigma}$ and $\widehat{\sigma}|\text{ND}$, respectively, indicate that the oil price dummies are mainly relevant for the oil price process and to some extent for the exchange rate process. It is the dummy $id97q1$ which seems to matter for the exchange rate equation, cf. Figures 4.2 and 4.4.

Figure 5.4 displays the 1-step ahead recursively estimated residuals $\pm 2SE$ and forecast and break-point Chow tests for the VAR model, scaled by their critical values at the 5% level. The VAR model appears to have relatively constant parameters over time, at least from 1987 and onwards. There are a few exceptions. The outliers among the RB residuals seem to slightly increase the estimates of their standard error over time. The $FI.Y$ residuals display large fluctuations in 1996/97. The rejection of the null hypotheses

Table 5.2: The VAR model

Method: Johansen
VAR model of order: 5
Endogenous variables (Y): *e, cpi, RB, FI.Y*, cpi^f, RB^f and *oilp*
Conditional variables (restricted): t and *OP2*
Unrestricted variables: *Constant, OP1*, $OP1_{-1}$, *id90q3, id97q1, CS* and *CS2*
Sample: Seasonally non-adjusted quarterly data, 1972:2-1997:4.

I. Single equation and system diagnostics

Y	$F_{ar,1-5}(5,54)$	$\chi^2_{nd}(2)$	$F_{arch,1-4}(4,51)$	$\widehat{\sigma}$	$\widehat{\sigma}$ | ND
e	0.88[0.50]	6.28[0.04]*	0.43[0.79]	0.0123	0.0134
cpi	1.05[0.40]	1.86[0.40]	0.23[0.92]	0.005	0.005
RB	0.63[0.68]	12.05[0.00]**	0.42[0.80]	0.004	0.004
$FI.Y$	1.20[0.32]	0.50[0.78]	1.41[0.24]	0.015	0.015
cpi^f	1.99[0.10]	1.09[0.58]	0.08[0.99]	0.003	0.004
RB^f	1.01[0.42]	2.65[0.27]	0.28[0.89]	0.003	0.003
$oilp$	0.75[0.60]	0.10[0.95]	1.02[0.60]	0.104	0.167
	$F_{ar,1-5}(245, 136)$	$\chi^2_{nd}(14)$			
VAR	1.22[0.09]	17.32[0.24]			

II. Cointegration rank

r	0	1	2	3	4	5	6	7
likl	3569.9	3600.4	3622.7	3640.0	3656.5	3669.2	3675.2	3677.1
$\widehat{\mu}$		0.45	0.35	0.30	0.26	0.22	0.11	0.04
H_0 :	$r=0$	$r\leq 1$	$r\leq 2$	$r\leq 3$	$r\leq 4$	$r\leq 5$	$r\leq 6$	
Trace	214.3**	153.4**	108.8**	72.3**	41.2	15.9	3.7	
95%	146.75	114.96	86.96	62.61	42.20	25.47	12.39	
99%	157.53	124.61	95.38	70.22	48.59	30.65	16.39	
Max	61.0**	44.6*	36.5	31.1	25.3	12.2	3.71	
95%	49.4	44.0	37.5	31.5	25.5	19.0	12.3	

Note: See Table 4.1 for details. $\widehat{\sigma}$ | ND standard error of residuals when all restricted and unrestricted dummy variables are left out from the VAR model. *Trace* denotes the trace statistics and *Max* denotes the max-eigenvalue statistics. The critical values for the tests are from Johansen (1995b) and Osterwald-Lenum (1992), respectively.

of parameter stability in the VAR model as a whole, by the 1-step ahead Chow tests and break point Chow tests in 1997, is apparently caused by these fluctuations. The rejection is however at the strict 5% level and the N-step ahead forecast Chow tests do not reject the null hypotheses of parameter stability in the VAR model.

Cointegration analysis within the VAR model

Panel II of Table 5.2 reports the estimated eigenvalues ($\widehat{\mu}$), the trace test statistics (*Trace*) and max eigenvalue test statistics (*Max*). These provide information on the cointegration rank, r. Testing the cointegration rank amounts to testing the number of eigenvalues different from zero. In the trace test, the null hypothesis is that the eigenvalues $\mu_i = 0$, $i = r+1$, $r+2$, ..., $r+6$, while the first r eigenvalues are non-zero. It rejects the null hypothesis of $r \leq 3$ against the alternative hypothesis of $r \geq 4$ at the 1% level. The max eigenvalue test, which tests the null hypothesis of $\mu_i = 0$ against the alternative hypothesis of $\mu_{i+1} = 0$, for $i = 0, 1, ...r$, rejects the null hypothesis of $r = 1$ against the alternative of $r = 2$ at the 5% level. The max eigenvalue test also comes quite close to rejecting the null hypotheses of $r = 2$ and $r = 3$ against the alternative of $r = 3$ and $r = 4$, respectively. Both test statistics indicate the possibility of $r = 5$ since their values are only slightly below their 95% quantiles.

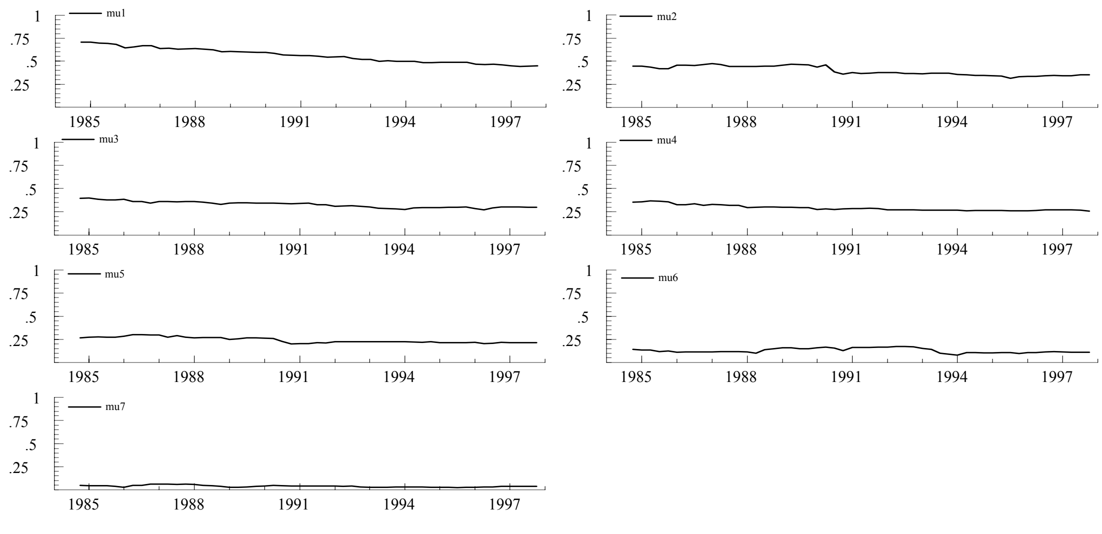


Figure 5.5: *Recursive estimates of eigenvalues based on the VAR model. For instance, mu1 denotes the recursive estimates of* μ_1. *The initial estimation period is 1972:2-1984:4.*

The tabulated critical values from Johansen (1995b) and Osterwald-Lenum (1992) may however be lower than the true critical values, since they have been derived under the assumption of at most two deterministic variables, the constant term and the time trend.

The inclusion of the step dummy *OP2* in the cointegration space and the unrestricted impulse dummies *OP1* and $OP1_{-1}$ may raise the (true) critical values. This is less likely to be the case with the unrestricted centered dummies, *id90q3, id97q1*, *CS* and *CS2* since they quickly converge to zero, cf. Doornik et al. (1999). Inclusion of additional deterministic variables whose effect do not die out asymptotically generally lead to higher critical values and more conservative tests for cointegration rank. However, given the strong rejection of $r \leq 3$ by the trace test and the fairly stable estimates of the eigenvalues, we proceed under the tentative assumption of $r = 4$, see Figure 5.5. Our confidence in the trace test rather than on the max eigenvalue test rests on Monte Carlo experiments suggesting that the former test is relatively more robust to excess kurtosis (and skewness) in the residuals than the latter, see e.g. Cheung and Lai (1993).

Table 5.3 tests restrictions on the β and α matrices when r is set to 4. Panel I tests restrictions on rows, β_i', $i = 1, 2, 3, 4$, of the β' matrix in order to identify the presumedly 4 linearly independent long run relations. Identification of 4 interpretable cointegrating relations can substantiate the choice of $r = 4$. The restrictions on β' are tested row by row, where each of the rows represents a cointegration vector under the null hypothesis. Since all variables in X are assumed to be integrated at most of order one, I(1), each cointegrating relation $\widehat{\beta}_i'X$ is stationary under the null hypothesis and I(1) under the alternative hypothesis. These tests are the multivariate alternative to the univariate ADF tests, but with the null hypothesis being that of stationarity and not of non-stationarity (in the unit root sense). The outcome of the tests is discussed below.

$\widetilde{\beta}_1'$ defines a relation between the nominal exchange rate and domestic and foreign consumer prices where the prices are assumed to be symmetrically related to the exchange rate. This restriction is accepted at a p-value of 0.15 with "correct" signs for both the domestic and foreign prices. The estimated coefficients are -0.57 and 0.57, respectively. In the second row of Table 5.3, $\widehat{\beta}_1'$ defines a relation in strict accordance with the PPP theory by imposing both the symmetry and proportionality restrictions on the domestic and foreign prices. The stationarity of this PPP term, $\widehat{\beta}_1'X = e - cpi + cpi^f$, can be accepted at a p-value of 0.07. This result appeared to be quite robust to a wide range of different specifications of the VAR model. Appendix C indicates this for changes in the

Table 5.3: Testing for long run relations and weak exogeneity, the VAR model.

I. Restrictions on rows of β'

	e	cpi	RB	$FI.Y$	cpi^f	RB^f	$oilp$	t	$OP2$	$\chi^2(.)$
$\widetilde{\beta}'_1$	1	-0.57	0	0	0.57	0	0	0	0	6.69[0.15]
$\widehat{\beta}'_1$	1	-1	0	0	1	0	0	0	0	10.16[0.07]
$\widetilde{\beta}'_2$	-0.69	0	1	0	0	-1	0	0	0	7.54[0.11]
$\widehat{\beta}'_2$	0	0	1	0	0	-1	0	0	0	19.01[0.00]**
$\widehat{\beta}'_3$	0	0	0	1	0	0	0	0	0.02	7.13[0.13]
$\widehat{\beta}'_4$	0	0	0	0	0	0	1	-0.03	-0.48	5.52[0.14]

II. Testing weak exogeneity, condition (1): $\alpha_{zy} = 0$.

j						$\chi^2(4)$	
RB^f :	$\alpha_{RB^f,\,1}$	$= \alpha_{RB^f,\,2}$	$= \alpha_{RB^f,\,3}$	$= \alpha_{RB^f,\,4}$	$= 0,$	9.27	[0.06]
cpi^f	$\alpha_{cpi^f,\,1}$	$= \alpha_{cpi^f,\,2}$	$= \alpha_{cpi^f,\,3}$	$= \alpha_{cpi^f,\,4}$	$= 0,$	9.35	[0.05]
$oilp$:	$\alpha_{oilp,\,1}$	$= \alpha_{oilp,\,2}$	$= \alpha_{oilp,\,3}$	$= \alpha_{oilp,\,4}$	$= 0,$	22.91	[0.00]**

					$\widehat{\alpha}_{oilp,\,4}$	$\chi^2(6)$	
$oilp$: $\widetilde{\beta}'$	$\cap$	$\alpha_{oilp,\,1} =$	$\alpha_{oilp,\,2} =$	$\alpha_{oilp,\,3} = 0,$	$\underset{(0.06)}{-0.17}$	12.67	[0.05]
					$\widehat{\alpha}_{e,\,1}$	$\chi^2(8)$	
e : $\widehat{\beta}'_1$	$\cap$	$\alpha_{e,\,2} =$	$\alpha_{e,\,3} =$	$\alpha_{e,\,4} = 0,$	$\underset{(0.04)}{-0.12}$	13.00	[0.11]

III. Testing $\alpha_{yz} = 0$.

j	e	cpi	RB	$FI.Y$	cpi^f	RB^f	$oilp$	t	$OP2$
$\widetilde{\beta}'_1$	$*$	$*$	$*$	$*$	$*$	$*$	0	$*$	$*$
$\widetilde{\beta}'_2$	$*$	$*$	$*$	$*$	$*$	$*$	0	$*$	$*$
$\widetilde{\beta}'_3$	$*$	$*$	$*$	$*$	$*$	$*$	0	$*$	$*$
$\widehat{\beta}'_4$	0	0	0	0	0	0	1	$*$	$*$
$\widetilde{\beta}' =$	$(\widetilde{\beta}'_1,$	$\widetilde{\beta}'_2,$	$\widetilde{\beta}'_3,$	$\widehat{\beta}'_4)'$	$: \chi^2(3)$	$=$	5.52	[0.14]	

e :	$\widetilde{\beta}'$	$\cap$	$\alpha_{e,\,4} =$	0,	$\chi^2(4)$	$=$	5.57	[0.23]
cpi	$\widetilde{\beta}'$	$\cap$	$\alpha_{cpi,\,4} =$	0,	$\chi^2(4)$	$=$	6.34	[0.18]
RB :	$\widetilde{\beta}'$	$\cap$	$\alpha_{RB,\,4} =$	0,	$\chi^2(4)$	$=$	7.16	[0.13]
$FI.Y$:	$\widetilde{\beta}'$	$\cap$	$\alpha_{FI.Y,\,4} =$	0,	$\chi^2(4)$	$=$	6.44	[0.17]
$\widetilde{\beta}' \cap$	$\alpha_{e,\,4} =$	$\alpha_{cpi,\,4} =$	$\alpha_{RB,\,4} =$	$\alpha_{FI.Y,\,4} = 0,$	$\chi^2(7)$	$=$	9.33	[0.23]

Note: The star $*$ indicates an unrestricted coefficient value. The cells heading $\widehat{\alpha}_{oilp,4}$ and $\widehat{\alpha}_{e,\,1}$ contain estimates of the unrestricted coefficients. A "$\sim$symbolises a less restricted vector or matrix. $\cap$ is used to indicate that restrictions are tested jointly.

number of lags of the VAR model and/or when we exclude the dummy variables. The appendix tabulates test results for the main hypotheses, for 10 different specifications of the VAR model.

$\widetilde{\beta}'_2$ defines a linear relation between the interest rate spread ($RB - RB^f$) and the nominal exchange rate, which is entered unrestricted. Stationarity of this relation is not rejected and it entails the following long run relation for the domestic interest rate: $RB = RB^f + 0.69e$. This relation resembles a central bank's interest rate response function under a policy of exchange rate targeting. The test of $\widehat{\beta}'_2$ shows that the interest rate spread is not stationary by itself, as implied by the UIP hypothesis. Experimentation with different model specification, however, suggested that this result is quite sensitive to the specification of the VAR model, (not reported).

The non-rejection of $\widehat{\beta}'_3$ as a cointegration vector, implying $FI.Y$ is a stationary variable, is in line with the result from the ADF test. Similarly, $\widehat{\beta}'_4$ implies that *oilp* is by itself a stationary term with a "broken" deterministic trend. The non-rejection of this hypothesis lends support to the proposition that the oil price process is not an integrated but a stationary process with infrequent changes in its mean level, see also Appendix C.

Identification of the 4 cointegrating relations substantiates the choice of $r = 4$ and suggests that there is only one cointegrating relation between the foreign denominated variables $Z = (cpi^f, RB^f, oilp)$; namely *oilp* itself, $\widehat{\beta}'_4 X$, which can be interpreted as $\widehat{\beta}'_z Z$, cf. equations (5.7) and (5.8).

Panel II of the table starts out by testing whether RB^f, cpi^f and *oilp* can be treated as weakly exogenous variables for all the cointegration vectors and the associated adjustment coefficients, that is, for β' and α. The weak exogeneity of RB^f and cpi^f is barely accepted at the 5% level but rejected of *oilp,* even at the 1% level.

Panel II then tests whether *oilp* can be treated as weakly exogenous for a partial β' and α. A 4 x 9 matrix $\widetilde{\beta}'$ is defined for this purpose, see panel III. The coefficient of *oilp* is set to zero in the first three rows while the fourth row is defined as $\widehat{\beta}'_4$ in panel I. The restrictions defining $\widetilde{\beta}'$ are accepted at a p-value of 0.14. The restrictions $\widetilde{\beta}' \cap \alpha_{oilp,\,1} = \alpha_{oilp,\,2} = \alpha_{oilp,\,3} = 0$ make sure that only $\widehat{\beta}'_4 X$ (or $\widehat{\beta}'_z Z$) enters the *oilp* equation, see panel II. Any other linear combination between the variables in X is barred from entering the *oilp* equation. These restrictions, which correspond to condition (1) for the weak exogeneity of *oilp* for a partial β' and α, are not rejected at the strict 5% level. The unrestricted coefficient estimate of $\widehat{\beta}_4'X$ in the *oilp* equation is -0.17, and differs significantly from zero. The test result adds to the evidence for treating *oilp* as a stationary variable in the sample at hand.

A joint test of the weak exogeneity of cpi^f, RB^f and condition (1) for the weak exogeneity of *oilp* was rejected at the 1% level (not reported). Yet economic intuition suggests that it is unlikely that these foreign denominated variables respond to deviations from the long run relations governing the relatively small Norwegian sub-markets, defined by e.g. the cointegrating relations above, $\widehat{\beta}_1'X$, $\widetilde{\beta}_2'X$ and $\widehat{\beta}_3'X$. A perhaps more plausible explanation for the rejection of the joint test may be that the VAR model is overparametrised relative to the number of observations. It is well known that overparametrisation often leads to the appearance of transient correlations between unrelated variables in small samples.

The last row of panel II tests the weak exogeneity of the nominal exchange rate for all the cointegration vectors in β' except $\widehat{\beta}_1'$, and for α_1 except $\alpha_{e,\,1}$. The imposed restrictions are accepted and implies that only the PPP vector ($\widehat{\beta}_1'X$) enters the exchange rate equation with an estimated weight of -0.12. Appendix C shows that this result is quite robust across the different specifications of the VAR model. It is shown that (the estimates of) the weight only varies in the range of (-0.135, -0.100), across the 10 different model specifications.

Panel III follows the same procedure as in panel II to test whether *oilp* has any long run effect on the Y variables, e, cpi, RB and $FI.Y$. Now, all other possible linear combinations of the variables in X are let into the equations for the Y variables, except *oilp*. For example, $\widetilde{\beta}' \cap \alpha_{e,\,4} = 0$ tests whether *oilp* can be excluded from the nominal exchange rate equation. These restrictions are accepted at the 5% level when placed equation by equation for all the Y variables, or jointly on all 4 equations. Appendix C indicates the robustness of this result.

To summarise, the results in Table 5.3 are consistent with $r = 4$. One of the identified cointegrating relation conforms with the PPP theory since the symmetry and proportionality restrictions are not rejected at the 5% level. The oil price (*oilp*) has appeared to be a cointegrating term by itself and changes in the oil price ($\Delta oilp$) in a given period have been found to respond to the level of the oil price in a stabilising way. More importantly, the oil price does not seem to have a statistically significant effect on either e or on *cpi, RB* and *FI.Y*. Table 5.3 also provides some support for condition (1) for valid conditioning on cpi^f, RB^f and *oilp,* in a conditional VAR model of e, *cpi, RB* and *FI.Y.*

However, the full VAR model is overparameterised, which may lead to underrejection of false null hypotheses and make room for transient correlations between unrelated variables. Hence, some of the obtained results are considered as indicative rather than firm evidence. In the next subsection, we test the robustness of e.g. the symmetry and proportionality restriction and other hypotheses of interest within the parsimonious conditional VAR model.

5.4. The conditional VAR model

This subsection proceeds under the assumption that condition (1) for the weak exogeneity of cpi^f, RB^f and *oilp* is satisfied for the parameters of interest. It starts out by formulating a conditional VAR model to check whether condition (2), i.e. $\alpha_{yz} - \omega\alpha_{zz} = 0$, for the weak exogeneity of these variables is fulfilled. Since there is only one cointegrating term among the Z variables, *oilp* itself, a test of this condition amounts to testing for the presence of *oilp* in the conditional VAR model of the Y variables. Thereafter, a parsimonious conditional VAR model is derived to test the hypotheses of interest.

Table 5.4 defines a conditional VAR model of the endogenous variables, *e, cpi, RB* and *FI.Y.* This model is derived by imposing zero restrictions on some of the parameters of a conditional model that corresponded directly to the full VAR model, cf. equation (5.7). The endogenous variables have four lags while relative changes in the conditioning variables, ΔRB^f, $\Delta oilp$ and Δcpi^f, enter unrestricted with up to 2 and 3 lags, respectively. Due to the conditioning on *oilp,* the impulse dummy for the Gulf War (*id90q3*) also became insignificant at the 5% level but not the dummies for OPEC I and the oil price rise in 1997, $OP1$, $OP1_{-1}$ and *id97q1,* respectively. The model was estimated with cpi^f_{t-1}, RB^f_{t-1}, $oilp_{t-1}$, $OP2$ and t restricted to the cointegration space while the remaining dummy variables and the constant term entered unrestricted.

The residuals from the conditional model appear to have the same properties as the

Table 5.4: The conditional VAR model.

Method: Johansen
Conditional VAR model of order: 4
Endogenous variables (Y): *e, cpi, RB* and *FI.Y*
Conditional variables, restricted: cpi^f_{t-1}, RB^f_{t-1}, $oilp_{t-1}$, t and *OP2*
Conditional variables, unrestricted: Δcpi^f_t,...Δcpi^f_{t-3}, ΔRB^f_t,...ΔRB^f_{t-2}, $\Delta oilp_t$, ...$\Delta oilp_{t-2}$, *OP1*, OP_{-1}, *id97q1*, *Constant*, CS_t and CS_{t-2}.
Sample: Seasonally non-adjusted quarterly data, 1972:2-1997:4.

I. Single equation and system diagnostics

	$F_{ar,1-5}(5,61)$	$\chi^2_{nd}(2)$	$F_{het,Xi^2}(41,24)$	$F_{arch,1-4}(4,58)$	$\widehat{\sigma}$
e	1.70[0.15]	1.36[0.51]	0.27[1.00]	0.20[0.94]	.0119
cpi	1.62[0.17]	1.26[0.53]	0.48[0.98]	0.15[0.96]	.0050
RB	1.42[0.23]	8.63[0.01]*	0.38[1.00]	0.47[0.76]	.0036
$FI.Y$	1.81[0.12]	0.50[0.78]	0.92[0.61]	0.49[0.75]	.0149

	$F^v_{ar,1-5}(80,172)$	$\chi^{2,v}_{nd}(8)$	$F^v_{het,\ Xi^2}(410,175)$
VAR	1.04[0.42]	10.28[0.25]	0.35[1.00]

II. Testing weak exogeneity, condition (2): $\alpha_{yz} - \omega\alpha_{zz} = 0$.

j					$\chi^2(4)$
e :	$\widetilde{\beta}' \cap$	$\alpha_{e,\,4} =$	0,	:	6.45[0.17]
cpi :	$\widetilde{\beta}' \cap$	$\alpha_{cpi,\,4} =$	0,	:	6.41[0.17]
RB	$\widetilde{\beta}' \cap$	$\alpha_{RB,\,4} =$	0,	:	10.30[0.04]*
$FI.Y$:	$\widetilde{\beta}' \cap$	$\alpha_{FI.Y,\,4} =$	0,	:	4.70[0.32]

	$\chi^2(9)$
$\widetilde{\beta}' \cap \alpha_{e,\,4} = \alpha_{cpi,\,4} = \alpha_{RB,\,4} = \alpha_{FI.Y,\,4} = 0$, :	16.54[0.06]

Note: $\widetilde{\beta}'$ is defined as in Table 5.3. See Table 4.1 for details about the tests.

residuals from the full VAR model. There are no significant violations of the hypotheses of no residual autocorrelation up to order 5, no heteroscedasticity up to order 4 and no violation of the normality assumption for most of the residuals except the *RB*-residuals, as earlier. Figure 5.6 indicates that the model has fairly stable parameters, in particular the equations for the nominal exchange rate and consumer prices. There are signs of changes in the *RB* and *FI.Y* equations, but not stronger than in the full VAR model, cf. Figure 5.4.

Panel II of Table 5.4 tests condition (2) for the weak exogeneity of cpi^f, RB^f and *oilp* by testing whether *oilp* enters any of the equations for the endogenous variables.[18] The

[18]Note that if condition (2) is satisfied, cpi^f, RB^f can be regarded as weakly exogenous for the para-

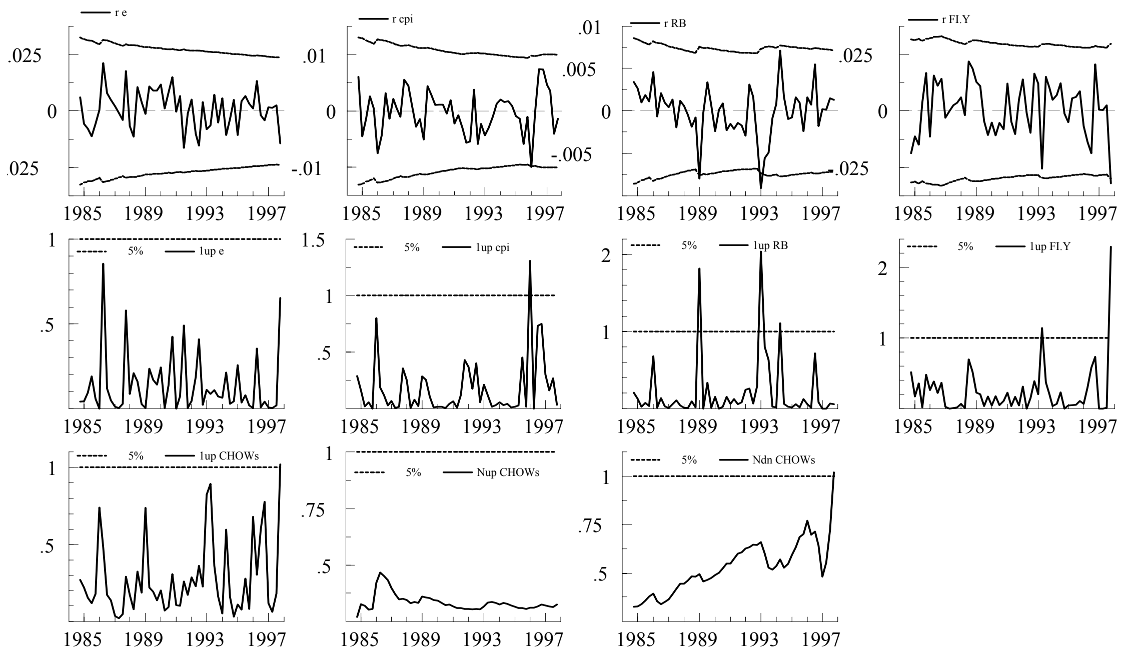


Figure 5.6: *Constancy statistics for the four equation conditional VAR model in Table 5.4. Top panel: One-step residuals* $\pm 2SE_t$. *Middle panel: Chow statistics for each of the four equations. Bottom panel: Chow statistics for the conditional system. All Chow statistics are scaled by their (one-off) critical values at the 5% level of significance. The initial estimation period is 1972:2-1984:4.*

tests are similar to the tests performed in panel III of Table 5.3 and are conducted under the assumption of $r = 4$. Condition (2) seems to be satisfied for all the equations when tested equation by equation and jointly for all equations, at around the 5% level.

oilp does not have statistically significant effect on any of the endogenous variables in the conditional model of Table 5.4, and is therefore left out from the model together with *OP2,* and the trend t. The trend appeared to be relevant only in the oil price vector, $\widehat{\beta}_4$, see panel I in Table 5.3. Table 5.5 presents a conditional model without these variables, denoted as the parsimonious conditional VAR model.

The residual properties from the conditional model in Table 5.4 seem to have carried over to the parsimonious conditional VAR model when judged at the 5% level. The residual standard errors have slightly increased for all equations, however. In addition, the tests of

meters of interest even if they respond to *oilp*.

Table 5.5: The parsimonious conditional VAR model.

Method: Johansen
Conditional VAR model of order: 4
Endogenous variables (Y): *e, cpi, RB* and *FI.Y*
Conditional variables, restricted: cpi^f_{t-1} and RB^f_{t-1}
Conditional variables, unrestricted: $\Delta cpi^f_t, ...\Delta cpi^f_{t-3}$, $\Delta RB^f_t, ...\Delta RB^f_{t-2}$, $\Delta oilp_t, ...\Delta oilp_{t-2}$, *OP1*, OP_{-1}, *id97q1*, *Constant*, CS_t and CS_{t-2}.
Sample: Seasonally non-adjusted quarterly data, 1972:2-1997:4.

I. **Single equation and system diagnostics**

	$F_{ar,1-5}(5,64)$	$\chi^2_{nd}(2)$	$F_{het,Xi^2}(36,32)$	$F_{arch,1-4}(4,61)$	$\widehat{\sigma}$
e	1.71[0.15]	1.24[0.54]	0.33[1.00]	0.05[0.99]	.0122
cpi	1.19[0.32]	2.99[0.22]	0.47[0.93]	0.11[0.98]	.0054
RB	1.82[0.12]	10.31[0.01]*	0.48[0.98]	0.40[0.81]	.0036
FI.Y	0.55[0.74]	1.28[0.53]	1.31[0.22]	1.22[0.31]	.0154
	$F^v_{ar,1-5}(80,183)$	$\chi^{2,v}_{nd}(8)$	$F^v_{het, Xi^2}(360,250)$		
VAR	1.03[0.44]	13.83[0.09]	0.48[1.000]		

Note: See Table 4.1 for details about the tests.

parameter stability show relatively stronger signs of parameter instability in 1997, mainly because of an outlier in the *FI.Y* equation, see Figure 5.7.

Cointegration analysis within the parsimonious conditional VAR model

The cointegration analysis proceeds under the assumption of three cointegrating relations, $r = 3$, since the 4th cointegrating term *oilp* has been left out from the conditional model. The rest of the β and α matrices can be interpreted as β_y and α_{yy} , as in equation (5.7). Table 5.6 reports the results when restrictions are placed on β'_y and α_{yy}.

Panel I shows that the PPP hypothesis, UIP hypothesis and the stationarity of *FI.Y* are not rejected at the 5% level, when tested individually or jointly. Figure 5.8 displays the test statistics when PPP restrictions defined by $\widehat{\beta}'_1$ are imposed recursively. The figure indicates that the PPP hypothesis is "easily" accepted by the data, at least from 1985. Notice also that the nominal exchange rate (e) need not be present in the second relation, which implies that the interest rate spread is stationary by itself.

Panel II of Table 5.6 reports the unrestricted estimates of the elements of the α_{yy} matrix when all rows of β'_y matrix are restricted, as shown in panel I. The first row contains the estimated weights of the three cointegrating vectors in the nominal exchange rate equation.

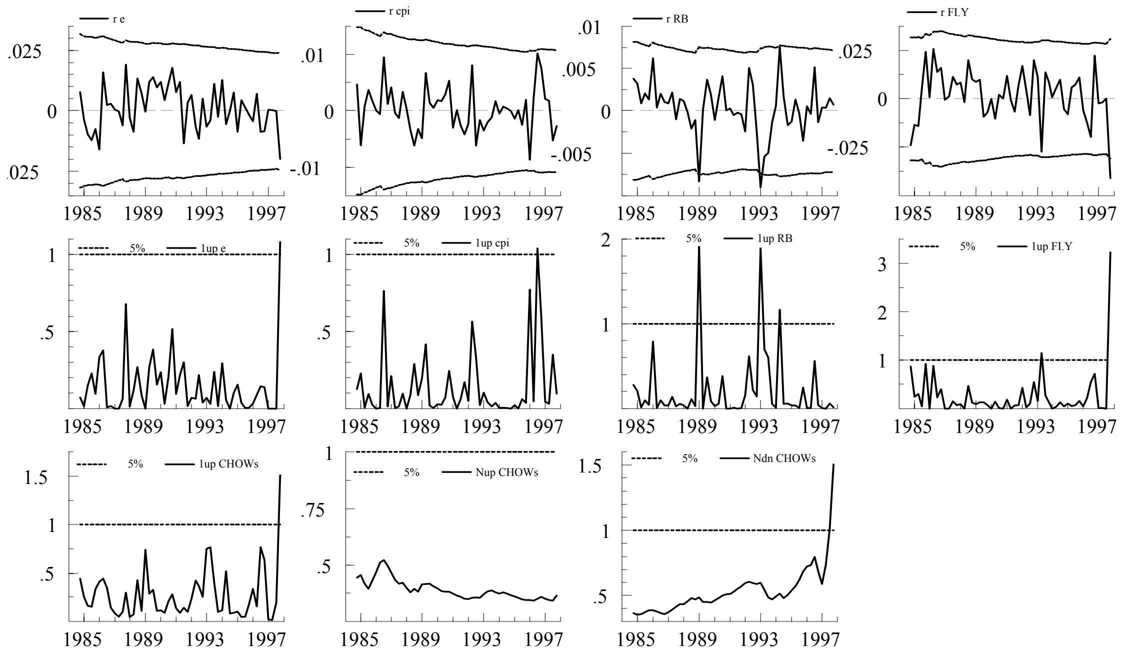


Figure 5.7: *Constancy statistics for the four equation conditional VAR model in table 5.5. Top panel: One-step residuals* $\pm 2SE_t$. *Middle panel: Chow statistics for each of the four equations. Bottom panel: Chow statistics for the conditional system. All Chow statistics are scaled by their (one-off) critical values at the 5% level of significance. The initial estimation period is 1972:2-1984:4.*

The estimated weights are consistent with the results from the full VAR model in Table 5.3. They indicate that only the PPP vector $\widehat{\beta}_1' X$ enters significantly in the nominal exchange rate equation with a weight of -0.11. This weight is quite close to the estimated weight within the full VAR model, -0.12, cf. also Appendix C. This supplements the evidence in favour of the presumed weak exogeneity of RB^f, cpi^f and *oilp,* at least for the parameters of interest in the nominal exchange rate equation. The second row of the estimated α_{yy} matrix indicates that Δcpi responds positively to deviations from the purchasing power parity and to the interest rate spread. The spread can be interpreted as the expected rate of depreciation, in accordance with the UIP hypothesis. The unrestricted estimates of the elements of the α_{yy} matrix also suggest that $\widehat{\beta}_1' X$ is not significant in the ΔRB equation nor in the $\Delta FI.Y$ equation.

The unrestricted estimate of the weight of $FI.Y_{t-1}$ in the $\Delta FI.Y$ equation is insignif-

Table 5.6: Tests for long run relations and weak exogeneity, conditional VAR model.

I. Testing hypotheses by restrictions on rows of β_y'

		e	cpi	RB	$FI.Y$	cpi^f	RB^f		$\chi^2(3)$
1	$\widehat{\beta}_1'$:	1	-1	0	0	1	0	:	2.23[0.53]
2	$\widehat{\beta}_2'$:	0	0	1	0	0	-1	:	7.11[0.07]
3	$\widehat{\beta}_3'$:	0	0	0	1	0	0	:	6.74[0.08]
	$\widehat{\beta}_y'$:	$\widehat{\beta}_1'$	$\cap$	$\widehat{\beta}_2'$	$\cap$	$\widehat{\beta}_3'$	:	$\chi^2(9)$ = 17.25	[0.05]

II. Unrestricted estimates of α_{yy}

$$\begin{array}{cccc} \widehat{\alpha}_{yy} & 1 & 2 & 3 \\ \Delta e: & \widehat{\alpha}_{e,1} & \widehat{\alpha}_{e,2} & \widehat{\alpha}_{e,3} \\ \Delta cpi: & \widehat{\alpha}_{cpi,1} & \widehat{\alpha}_{cpi,2} & \widehat{\alpha}_{cpi,3} \\ \Delta RB: & \widehat{\alpha}_{RB,1} & \widehat{\alpha}_{RB,2} & \widehat{\alpha}_{RB,3} \\ \Delta FI.Y: & \widehat{\alpha}_{FI.Y,1} & \widehat{\alpha}_{FI.Y,2} & \widehat{\alpha}_{FI.Y,3} \end{array} = \begin{array}{cccc} \widehat{\alpha}_{yy} & 1 & 2 & 3 \\ \Delta e: & \underset{(0.059)}{-0.112} & \underset{(0.104)}{-0.034} & \underset{(0.149)}{-0.081} \\ \Delta cpi: & \underset{(0.026)}{0.055} & \underset{(0.046)}{0.120} & \underset{(0.067)}{0.052} \\ \Delta RB: & \underset{(0.017)}{-0.017} & \underset{(0.031)}{-0.060} & \underset{(0.044)}{-0.087} \\ \Delta FI.Y: & \underset{(0.074)}{-0.019} & \underset{(0.131)}{-0.005} & \underset{(0.189)}{-0.234} \end{array}$$

III. Testing weak exogeneity

1.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{e,2} =$	$\alpha_{e,3} =$	0,		$\chi^2(11)$	17.69[0.09]
2.			$\alpha_{cpi,1} =$	$\alpha_{cpi,2} =$	$\alpha_{cpi,3} =$	0,	$\chi^2(3)$	= 20.39[0.02]*
3.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{cpi,1} =$	$\alpha_{cpi,2} =$	$\alpha_{cpi,3} =$	0,	$\chi^2(12)$	= 26.90[0.01]**
4.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{cpi,1}$	= 0,			$\chi^2(10)$	= 23.34[0.01]**
5.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{RB,1}$	= 0,			$\chi^2(10)$	= 18.68[0.05]
6.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{FI.Y,1}$	= 0,			$\chi^2(10)$	= 17.34[0.07]
7.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{RB,1}$	=	$\alpha_{FI.Y,1}$	= 0,	$\chi^2(11)$	= 18.86[0.06]
8.	$\widehat{\beta}_y'$	$\cap$	$\alpha_{RB,1} =$	$\alpha_{FI.Y,1} =$	$\alpha_{e,2} =$	$\alpha_{e,3} = 0,$	$\chi^2(13)$	19.28[0.12]

IV. $\widehat{\alpha}_{yy}\widehat{\beta}_y'X$

$$\widehat{\alpha}_{yy}\widehat{\beta}_y'X = \left[\begin{array}{cccc} \Delta e: & \underset{(0.039)}{-0.087} & 0 & 0 \\ \Delta cpi: & \underset{(0.021)}{0.051} & \underset{(0.041)}{0.108} & 0 \\ \Delta RB: & 0 & \underset{(0.023)}{-0.040} & \underset{(0.036)}{-0.071} \\ \Delta FI.Y: & 0 & 0 & \underset{(0.154)}{-0.205} \end{array}\right] \left[\begin{array}{c} e-(cpi-cpi^f) \\ RB-RB^f \\ FI.Y \end{array}\right],$$

$$:\chi^2(15) = 20.64[0.15]$$

Note: The results in this table are based on the model in Table 5.5.

icant at the 5% level, with a t-value of -1.24. The estimated weight of the interest rate spread in the ΔRB equation is also a borderline case with a t-value of -1.94. The insignificance of $FI.Y$ might owe to inadequate modelling of $\Delta FI.Y$, which generally leads to higher standard errors and biased coefficient estimates. This interpretation fits well with the results from the ADF test where up to 7 lags of $\Delta FI.Y$ are significant while the coefficient estimate of $FI.Y_{t-1}$ is -0.62, see Table 5.1. This suggests that a larger information set, and/or variables of more relevance for the behaviour of $\Delta FI.Y$ are required to obtain more precise estimates of $FI.Y_{t-1}$ and in modelling $\Delta FI.Y$. Similarly, the ΔRB equation might also benefit from an extension of the information set, e.g. by including short term domestic and foreign interest rates. Moreover, by taking proper account of regulations of capital flows, in particular during the 1970s and the early 1980s; for instance, by allowing for separate equilibrium terms before and after the dismantling of the Norwegian foreign exchange rate regulations in 1990, cf. Figure 5.1. A more satisfactory modelling of $\Delta FI.Y$ and ΔRB is however not among the main concerns of this study and is therefore not pursued further.

Panel III tests restrictions on the feedback coefficients (weights) jointly with the restrictions on the β_y' matrix (in all cases but one) and the results support the impression from panel II. The first row shows that the exclusion of the second and the third cointegrating relations from the nominal exchange rate equation is accepted at a p-value of 0.09. The tests in the second, third and the fourth row suggest that *cpi* adjusts to divergence from PPP, that is, $\alpha_{cpi,1}$ is not zero. The second and third row of panel III show that the weak exogeneity of *cpi* for all cointegrating relations is rejected at the 5% level. The fourth row tests explicitly for the absence of the PPP vector ($\widehat{\beta}_1'X$) from the consumer price equation and rejects this hypothesis at the 5% level. The remaining tests in panel III test whether the PPP vector is absent from the equations for *RB* and *FI.Y.* The tests are conducted equation by equation and jointly. In particular, the final row tests the valditity of excluding the interest rate spread and *FI.Y*, jointly with the exclusion of the PPP vector from the *RB* and *FI.Y* equations. The imposed restrictions are not rejected at the 5 % level.

Panel IV shows the three long run relations together with their weights in the four equations for the endogenous variables. Two additional zero restrictions have been imposed on the α_{yy} matrix, relative to the last row in panel III. The joint restrictions on the α_{yy} and β_y' matrices are accepted with a p-value of 0.15. The joint restrictions are also tested recursively from 1985 and onwards in Figure 5.8. The graphed values of the test statistics do not show rejection of these restrictions at the 5% level, in any period from 1985 to the end of the sample in 1997:4.

The elements of the α_{yy} matrix are more precisely estimated in panel IV since their estimated standard errors are slightly lower than in panel II. Taken at face value, the estimate of $\widehat{\alpha}_{e,1}$, now -0.087, is slightly lower than the unrestricted estimate of -0.11. The $\widehat{\alpha}_{yy}$ matrix shows that $FI.Y$ only adjusts to its own level and that only RB appears to respond to the level of $FI.Y$. A joint test with a zero restriction on $FI.Y$ in the ΔRB-equation is however accepted at the 5% level. The relevant $\chi^2(16)$ test statistic is 25.9 with at p-value of 0.06. This implies that $FI.Y$ can be treated as a weakly exogenous variable when modelling e, cpi and RB. This, however, does not apply to RB when modelling the nominal exchange rate. Even though $\widehat{\beta}_1'X$ is absent from the RB equation, cf. condition (1), it "shares" $\widehat{\beta}_2'X$ with the cpi equation that contains $\widehat{\beta}_1'X$. Hence RB is not weakly exogenous for $\widehat{\beta}_1'$. Consequently, valid inference on the long run parameters in the nominal exchange rate equation is not warranted on the basis of a conditional model of the nominal exchange rate alone. It requires a joint conditional model of at least e, cpi and RB.[19]

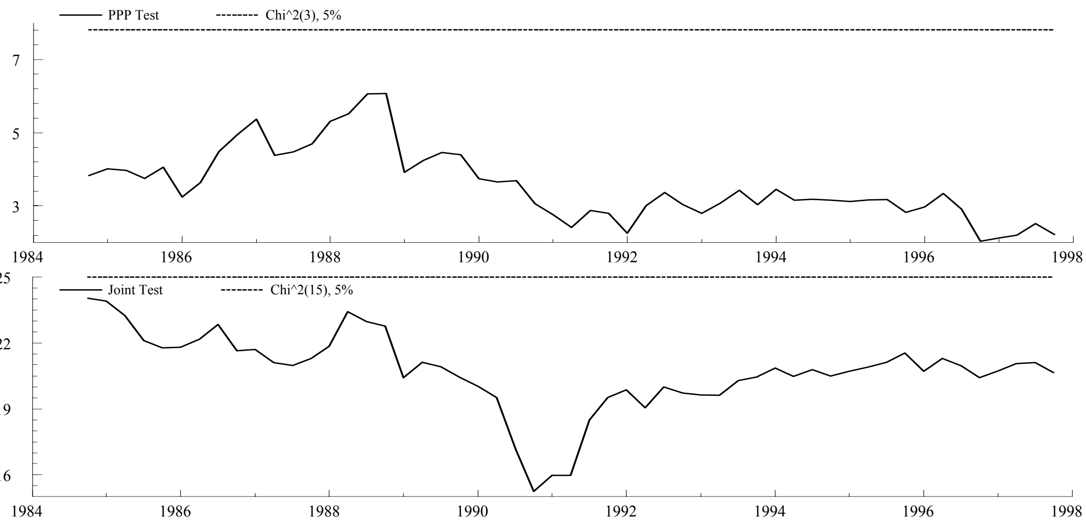


Figure 5.8: *Recursive encompassing tests for the conditional VAR model in table 5.5. The upper graph shows the recursive test statistcis when only the PPP restrictions are imposed. The lower graph shows the recursive statistcis when all restrictions defined in panel III of the table are jointly imposed. The initial estimation period is 1972:2-1984:4.*

[19] Alternatively, valid inference on the parameters of interest in the nominal exchange rate equation requires joint modelling of cpi. Valid modelling of cpi, however, cannot be pursued unless RB is modelled.

To summarise, the results from the (parsimonious) conditional VAR model are consistent with the results from the full VAR model. Accordingly, the symmetry and proportionality restrictions implied by the PPP theory are not rejected. Moreover, oil prices and financial investment abroad do not have long run effects on the nominal exchange rate. Their effects on the domestic consumer prices and interest rates have also been found to be insignificant. Domestic consumer prices respond positively to the interest rate spread, which can be interpreted as an indicator of expected depreciation. More importantly, they respond to deviations from the PPP in a statistically significant way, and contribute to re-establish the purchasing power parity in the long run. This is not only consistent with the view that the PPP is a theory for both the nominal exchange rate and prices, but also with the Scandinavian model of inflation, see Subsection 2.3.

6. Conclusions

Despite the emerging consensus on the validity of purchasing power parity (PPP) between trading countries in the long run, it is commonly rejected in data predominantly exposed to real shocks. This essay presents novel results against this background. The essay tests for PPP between Norway and its trading partners by examining its implications for the behaviour of the Norwegian real and nominal exchange rates. The empirical results are based on quarterly data from the post Bretton Woods period in which the Norwegian economy has been exposed to numerous real shocks such as discoveries of oil and gas resources and their revaluations through major shocks to oil prices. Yet, we find that the empirical evidence, obtained by employing a wide range of empirical models, is remarkably consistent with the PPP theory.

The evidence suggests that the real exchange rate is a stationary process which converges towards an equilibrium level that has been remarkably stable over most of the sample period. Moreover, the half life of a given deviation from the equilibrium rate is about 1 and 1/2 years, which is relatively fast when compared with the consensus estimates from the vast PPP literature. The symmetry and proportionality restrictions implied by the PPP theory are not rejected and the additional variables, including oil prices, are not found to have long run effects on the nominal exchange rate and prices. Furthermore, the

PPP theory is found to characterise the long run behaviour of both the nominal exchange rate and domestic prices. Both variables respond to deviations from the PPP and contribute to re-establish the parity in the long run. The response of the nominal exchange rate is however found to be almost twice the size of that for the domestic prices.

The relatively fast convergence of the real exchange rate towards its equilibrium rate is interpreted as a reflection of the active use of devaluations until the mid of 1986, centralised wage bargaining and of possibly strong international arbitrage pressure owing to the openness of the Norwegian economy. The relatively strong response of the nominal exchange rate to deviations from the parity, compared with that of domestic prices, is consistent with this interpretation.

The empirical analysis shows that the processes determining the exchange rate, consumer prices and interest rates are interdependent (in the sense of not being weakly exogenous for each others long run parameters). It follows that a change in the process of one of the variables is likely to induce a change in the processes of the other variables. For instance, a change from exchange rate targeting to inflation targeting may imply a simultaneous change in the exchange rate process, from stable to floating, and in the interest rate process. In the latter case, interest rates will no longer shadow the foreign interest rates to keep the exchange rate stable but become more attuned to the domestic activity level in order to achieve the inflation target. It is therefore not unlikely that a change from exchange rate targeting to inflation targeting affects the process that determines prices, as pointed out by Holden (1997) and Rødseth (1997a) *inter alia*.

This essay argues that one needs to take into account institutional features of an economy in order to assess the partial effect of e.g. a nominal exchange rate regime, or centralised wage setting, on the real exchange rate behaviour. An empirical assessment is however left to future studies since it is likely to involve a cross country data set. One possible way to proceed would be to model e.g. the half life of a deviation from a real exchange rate equilibrium using data on the degree of centralisation and/or coordinations of wage setting, degree of openness and by taking into account the nature of fiscal and exchange rate policies of each country; the latter, possibly by taking into account the degree of central bank independence, which usually makes a central bank less disposed to undertake devaluations in order to make up for deterioration in the competitiveness of an economy. A cross country study along the sketched lines might add weight to the interpretation of results in this essay and throw more light on the empirical regularities encountered in the PPP literature.

Appendices

Appendix A lists the trade weights of Norway's main trading partners, used in constructing the trade weighted exchange rate. The countries listed in the table are also known as basket countries. The trade weighted exchange rate is a weighted average of the value of these countries' currencies measured in NOK. Appendix B provides definitions and sources of the data used in this study. Appendix C tests some of the main hypotheses in different specifications of the full VAR model. This exercise is partly meant as an illustration of the robustness of results, to different types of model (mis)specifications, even though the usual distributions of the test statistics are unlikely to be valid.

Appendix A: Trade weights

Table 6.1: Trade weights based on import shares of basket countries

Austria	Belgium	Canada	Denmark	Finland	France	Germany
0.014	0.033	0.020	0.077	0.049	0.044	0.175
Italy	Japan	Netherland	Sweden	Switzerland	UK	USA
0.033	0.063	0.041	0.206	0.019	0.134	0.092

Note: Each country is assigned a weight equal to its average import share, in the total import from the 14 countries in this table. These 14 countries were the basket countries until October 1990. The average import shares are based on data for period 1978-1987, see Naug (1990, pp. 103).

Appendix B: Data definitions

Unless stated otherwise, the variables listed below are taken from the data base for RIMINI, the quarterly macroeconometric model used in Norges Bank. The main sources for RIMINI's data base are Quarterly National Accounts, FINDATR, TROLL8, OECD_MEI and IFS. These data bases are maintained by Statistics Norway, Norges Bank, Norges Bank, OECD and IMF, respectively. The RIMINI names of the variables are indicated in square brackets []. Note that this essay employs seasonally unadjusted quarterly data.

CPI : Consumer price index for Norway, 1991 = 1. [CPI].

CPI^f: Trade weighted average of consumer price indices for Norway's trading partners. Measured in foreign currency, 1991 = 1. [PCKONK].

C_G : Public consumption expenditures, fixed 1995 prices, Mill. Norwegian krone (NOK). [CO].

CS : Centered seasonal dummy variable (mean zero) for the first quarter in each year. It is 0.75 in the first quarter and -0.25 in each of the three other quarters, for every year.

E : Trade weighted nominal value of NOK, 1991 = 1. [PBVAL].

$FI.Y$: A measure of foreign net financial investment in Norway, fixed 1991 prices, Mill. NOK. Constructed by taking the first difference of the net foreign debts' share of GDP, [LZ.Y], i.e. $FI.Y = \Delta$LZ.Y.

g : Public expenditures' share of GDP, i.e. $g = (C_G + J_G)/Y$.

id86q2 : Impulse dummy related to the oil price fall in 1986. It has a value of 1 in 1986:2 and zero elsewhere.

id90q3 : Impulse dummy for the Gulf War. It has a value of 1 in 1990:3, -1 in 1990:4 and zero elsewhere.

id97q1 : Impulse dummy related to the oil price hike in 1996/97. It has a value of 1 in 1997:1, -1 in 1997:2 and zero elsewhere.

J_G : Public expenditures for gross real investment, fixed 1995 prices, Mill. NOK. [JO].

OILP : Price per barrel of Brent Blend crude oil in US dollars. Source TROLL8, series no. Q2001712.

OP1 : Impulse dummy for OPEC I. It has a value of 1 in 1974:1, -0.3 in 1974:2 and zero elsewhere.

OP2 : Step dummy for OPEC II. It has a value of 1 over period 1979:1-1985:4 and zero elsewhere.

Q : Value added per unit labour cost in Norway. The inverse of value added based unit labour costs. 1/[LPE.Y].

Q^f: Value added per unit labour cost in trading partners. The inverse of trade weighted average of value added based unit labour costs. 1/[M.LPE].

R : Trade weighted real exchange rate, defined as R = $(E \text{ x } CPI^f)/CPI$. [RPBVAL].

RB : Yield on 6 years Norwegian government bonds, quarterly average. [R.BS].

RB^f : NOK basket-weighted average of interest rates on long term foreign bonds. [R.BKUR].

RS : *3* month Euro krone interest rate. [RS].

U : Total unemployment rate, fraction of labour force exclusive self employed and on labour market programs. [UTOT].

Y : Gross domestic product for Norway. Mill. NOK, fixed 1995 prices. [Y].

Appendix C: Sensitivity analysis of the main results

Here some of the main results in Section 5 are tested in different specifications of the full VAR model. Specifically, we reexamine the key hypotheses in more restricted versions of the full VAR model employed in Section 5. These different specifications might be preferred by some analysts on the basis of their parsimony, higher degrees of freedom or because of distaste for the use of dummy variables in general, or for the way they have been specified and used in this study. For convenience and transparency, the VAR model is only changed in two different ways. First, the hypotheses are tested within VAR models that only differ from each other in the number of lags. Next, all dummy variables including seasonal dummies are excluded from the model and hypotheses are tested by changing the number of lags in the VAR model. Another simplification is that only those hypotheses are tested which are of direct relevance for the theme of this study. This exercise serves two purposes:

Firstly, it substantiates the choice of the VAR model in Table 5.2 as an approximation to the data generating process. Note that none of the other models in the table have residuals which are normally distributed and free of autocorrelation.

Secondly, it lends some credibility to the results reported in this study. The main hypotheses seem to be accepted by the tests even in models that display signs of misspecification, see the test statistics for the autocorrelation tests and the normality tests. Note however, that the results are only indicative, since the residuals are in general not IIDN(.).

Table 6.2: Tests of hypotheses within different specifications of the VAR model.

I. Defining hypotheses

		e	cpi	RB	$FI.Y$	cpi^f	RB^f	$oilp$	t	$OP2$
a)	$\widehat{\beta}'_1$	1	-1	0	0	1	0	0	0	0
b)	$\widehat{\beta}'_4$	0	0	0	0	0	0	1	$*$	$*$
c)	$\alpha_{cpi} = 0:$		$\alpha_{cpi,\,1} =$		$\alpha_{cpi,\,2} =$		$\alpha_{cpi,\,3} =$		$\alpha_{cpi,\,4} =$	0
d)	$\widehat{\beta}'_1 \cap \widetilde{\alpha}_e = 0:$			$\widehat{\beta}'_1 \cap$	$\alpha_{e,\,2} =$		$\alpha_{e,\,3} =$		$\alpha_{e,\,4} =$	0
e)	$\widetilde{\beta}' \cap \widetilde{\alpha}_{oilp} = 0:$	$\widetilde{\beta}' \cap$	$\alpha_{e,\,4} =$		$\alpha_{cpi,\,4} =$		$\alpha_{RB,\,4} =$		$\alpha_{FI.Y,\,4} =$	0

II. Changing lag length p

H_0	VAR(5)	VAR(4)	VAR(3)	VAR(2)	VAR(1)
$\widehat{\beta}'_1$	10.16[.07]	12.10[.04]	9.65[.09]	15.43[0.01]	6.54[.26]
$\widehat{\beta}'_4$	5.52[.14]	6.32[.10]	5.28[0.15]	2.35[0.50]	7.74[.05]
$\alpha_{cpi} = 0$	18.61[.00]	24.95[.00]	18.65[.00]	28.70[.00]	95.40[.00]
$\widehat{\beta}'_1 \cap \widetilde{\alpha}_e = 0$	13.00[.11]	15.50[0.05]	11.26[.19]	15.99[.04]	21.91[.01]
$(\widehat{\alpha}_{e,1})$	(-0.121)	(-0.112)	(-0.104)	(-0.122)	(-0.100)
$\widetilde{\beta}' \cap \widetilde{\alpha}_{oilp} = 0$	9.34[.23]	14.60[0.04]	9.12[.24]	13.28[.07]	14.48[.04]
$F^v_{ar,1-5}(.)$	1.23[.09]	1.34[.01]	1.67[.00]	1.50[.00]	1.70[.00]
$\chi^{v,\,2}_{nd}(14)$	17.32[.24]	24.41[.04]	31.20[.01]	45.17[.00]	52.34[.00]

III. Excluding all dummies and changing p

H_0	VAR(5)	VAR(4)	VAR(3)	VAR(2)	VAR(1)
$\widehat{\beta}'_1$	4.47[.35]	4.26[.37]	4.58[.33]	10.99[.03]	6.16[.19]
$\widehat{\beta}'_4$	3.01[.39]	1.21[.75]	3.26[.35]	11.80[.01]	8.70[.03]
$\alpha_{cpi} = 0$	23.56[.00]	12.19[.02]	17.62[.00]	31.55[.00]	73.70[.00]
$\widehat{\beta}'_1 \cap \widetilde{\alpha}_e = 0$	9.46[.22]	8.19[.32]	8.59[.28]	13.98[.05]	18.45[.01]
$(\widehat{\alpha}_{e,1})$	(-0.132)	(-0.130)	(-0.103)	(-0.135)	(-0.113)
$\widetilde{\beta}' \cap \widetilde{\alpha}_{oilp} = 0$	4.61[.71]	4.30[.75]	5.27[.63]	16.83[.02]	12.11[.10]
$F^v_{ar,1-5}(.)$	1.17[.15]	1.07[.31]	1.40[.00]	1.34[.01]	1.77[.00]
$\chi^{v,\,2}_{nd}(14)$	50.52[.00]	57.62[.00]	67.99[.00]	70.52[.00]	75.89[.00]

Note: Panel I: a) the PPP relation is stationary, b) the *oilp* is a cointegrating term by itself. $*$ denotes an unrestricted coefficient. c) *cpi* is weakly exogenous in the system, i.e. does not respond to any of the disequilibria in the VAR model. d) only the PPP relation enters the exchange rate equation. e) oilp does not enter any of the equation for the domestic variables, e, cpi, RB and FI.Y, i.e. oil price does not have long run effects on any of these variables. Panel II: The hypotheses defined above are tested within the full VAR model for different number of lags. Here VAR(5) is as in Table 5.2. $\widehat{\alpha}_{e,1}$ is the adjustment coefficient associated with the PPP term in the exchange rate equation. The numbers in each of the columns are χ^2 tests statistics under the different null hypotheses. The square brackets contain the *p*-values under the null hypotheses. These are not valid since the residuals do not have IIDN(.) properties. The (system) tests for auto-correlation and normality, strongly rejects the hypotheses that the residuals in this system are in possession of such properties in most of the cases. $\widetilde{\beta}'$ is defined in panel III of Table 5.3. Panel III: Tests the same hypotheses as above, but within VAR models which do not include dummy variables at all. Note that $\widehat{\beta}'_4$ is defined without OP2 in this case.

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