2007/7

Documents

Andreas Fagereng

Documents

Exchange rate volatility and export performance

Evidence from disaggregated Norwegian data

Statistics Norway/Research Department

Preface

During my time as a student at the University of Oslo I have had the invaluable opportunity of periodically working as an intern at the Research Department (Unit for Macroeconomics) at Statistics Norway. Inspiration from this work and courses at the Econometrics Chair at the Humboldt University in Berlin during my stay as an exchange student has contributed to the content and outline of this thesis. The thesis work has been performed during my current period at Statistics Norway 2006/2007.

First and foremost I would like to thank my supervisor Pål Boug at Statistics Norway for his inspiring and encouraging guidance and support, and for always leaving his office door open. Without his instructive advices, the quality of this thesis as it is today would not have been possible. I also want to thank Statistics Norway for economic funding and the opportunity to work in a supporting environment. In particular Håvard Hungnes, Ådne Cappelen and Eivind Tveter have contributed with critical feedback and valuable suggestions. Finally I would like to thank family and friends for support throughout the whole process, and especially Rønnaug for her everlasting patience.

All estimations were performed using PcGive 10, see Hendry and Doornik (Modelling dynamic systems using PcGive, vol. II, Timberlake Consultants Ltd, London 2001) and Doornik and Hendry (Modelling dynamic systems using PcGive, vol. I, Timberlake Consultants Ltd, London 2001).

Oslo January 2007
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Abstract

Ever since the breakdown of the Bretton-Woods agreement in 1971 researchers and policymakers around the world have sought to answer the question whether uncertainty about the movements in the exchange rates affects international trade flows. Even with numerous empirical attempts over the last decades, no consensus seems to be found. This thesis seeks to provide further evidence to the area within the context of a demand type export model, multivariate cointegration techniques and disaggregated data from the Norwegian industry. Applying a measure of exchange rate volatility from a GARCH (Generalized Autoregressive Conditional Heteroskedasticity) model, we find no evidence of any connection between exchange rate volatility and export performance. An important aspect of the analysis is the discussion of the time series properties of the exchange rate volatility measure. We show that our conclusion is unaltered regardless of whether the exchange rate volatility is treated as a stationary or a nonstationary variable. Then, we provide a thorough empirical investigation of an estimated conditional equilibrium correction model (EqCM), which explains the export volume by relative prices and international demand conditions. We demonstrate that the estimated EqCM model is well-specified and reasonably stable in-sample and performs well in an out-of-sample forecasting exercise despite a major monetary policy regime shift in Norway.

Contents

1. Introduction

Since the breakdown of the Bretton-Woods agreement in 1971 and the transition to freely floating exchange rates in many of the worlds major trading countries both policymakers and researchers have been concerned about the impact of exchange rate volatility on international trade flows. Some argue from a theoretical point of view that increased exchange rate volatility following the move to floating exchange rates has an adverse effect on world trade due to risk averse exporters, see e.g. Hooper and Kohlhagen (1978), Artus (1983) and Brodsky (1984). A number of studies find evidence of such a hypothesis; see among others Chowdhury (1993), Arize (1995), Arize *et al.* (2000) and de Vita and Abbot (2004). Yet, others find no evidence suggesting that exchange rate volatility has any significant impact on trade; see e.g. Viaene and Vries (1992) and Aristotelous (2001). Considering today's well-developed financial markets, it can be argued that exporters at least to some extent should be able to hedge themselves from uncertainty associated with exchange rate volatility. The negative impact of trade may then be expected to decline, if not completely removed, see Choudhry (2005) for a thorough discussion. To complete the list of contradictory findings Asserry and Peel (1991) are among those who have performed studies showing that exchange rate volatility has a positive effect on trade. Finance theory may give an explanation for this positive relationship between exchange rate volatility and trade flows. Standard finance theory predicts that increased volatility of the underlying asset will increase the value of an option, see e.g. Hull (2005). Hence, assuming that a traded good is similar to an option, exchange rate volatility may lead to increased rather than decreased trade flows. Accordingly, despite numerous empirical attempts over the last decades, no consensus about the nature and magnitude of the relationship between exchange rate volatility and trade flows seems to be established in the literature.

McKenzie (1999) gives a thorough review of the literature over the last decades and discusses several issues that may be important when determining the impact of exchange rate volatility on trade. These issues are mainly related to which exchange rate volatility measure to use, which sample period to consider, which countries to study, which data, explanatory variables and aggregation level to employ and which estimation method to apply in each specific study at hand. Each of these issues and how they are handled in each case may be part of the explanations for the inconclusive findings about the empirical impact of exchange rate volatility on trade. Consequently, it is crucial to carefully consider relevant issues prior to performing an empirical analysis in order to provide reliable evidence.

When considering an export model and which data series to choose, the measure for volatility is the least obvious variable. Hence, ever since the initialization of the literature discussing possible effects from exchange rate volatility on trade, the method of how to measure exchange rate volatility has changed rather rapidly. McKenzie (1999) considers both the cases of nominal and real exchange rate volatility as well as the most commonly used statistical and econometrical techniques to measure exchange rate volatility. Also, the time series properties of the regressors in the export model have received increased attention. Graphical and formal investigations (by means of unit root testing) of the most customary included variables in export models (for instance prices and volume) typically reveal signs of nonstationarity. Thus, methods for testing of existence of long run relationships among nonstationary variables (i.e. cointegration relationships), e.g. Johansen's method (1988, 1995), become helpful and may be important to establish balanced models so as to avoid spurious regressions.[1]

Equally important when studying the relationship between exchange rate volatility and trade flows is to have available a sample period with a sufficiently large number of observations from managed or freely floating exchange rate regimes. Otherwise, it may be hard to find a significant relationship, if any, by means of econometric methods. Also, monetary policy regime changes or sudden heavy movements in the exchange rate due to speculations or devaluations possibly changing the behaviour of exporters would be suitable events for closer investigation of the volatility mechanism on trade. It is also important to select relevant trading partners for the country in study. For a country in Latin America it would be appropriate to include its closest neighbouring trading countries, and maybe also the US. For a country in Europe outside the monetary union, conducting an analysis at least including EU trading partners could be reasonable. When it comes to the level of aggregation, McKenzie (1999) points out that a majority of existing studies rely on the use of aggregate data. Doing so, one implicitly assumes that all industries behave similar, and one potentially

[1] The variables used in the analysis are often found to be integrated of order one, I(1). For a discussion of time series properties, including a discussion of the cointegration method, please refer to the chapters 3 and 4 below.

misses out important industry or sector specific relationships between exchange rate volatility and trade flows.

Finally, it is vital to choose a model specification suitable for the study at hand and to include the most relevant variables in addition to the exchange rate volatility. Different approaches have been taken in the literature. For instance, Thursby and Thursby (1985,1987) consider the gravity model, which specifies the trade value between two countries to depend negatively on the distance between them, and positively on their national income. Recently, however, it is common to apply demand specified models. Both volume and value have been chosen to be the variable modelled, and foreign income and relative prices between the trading partners are commonly used as explanatory variables in the model specification, see e.g. Arize (1995, 1997) and de Vita and Abbott (2004). Some studies also include foreign direct investment and international labor division as additional explanatory variables in order to investigate their potential impact on trade, see Égert *et al.* (2005) and Stephan (2006).

This thesis aims to provide further evidence to the literature concerning the impact of exchange rate volatility on trade taking account of the issues addressed above. Specifically, we study exchange rate volatility and Norwegian exports to the major trading partners within a demand type model using disaggregated data for one of the main industries, namely machinery and equipment. We employ different measures of exchange rate volatility, multivariate cointegration techniques and equilibrium correction models using quarterly data for the period 1985 to 2005, hence covering periods of both fixed, managed floating and freely floating exchange rate regimes.[2]

Knowledge of the impact of exchange rate volatility on exports is of major importance for policymakers in a small open economy, like the Norwegian, which depends heavily on its trade with the outside world. For instance, in 2005 the value of the Norwegian exports amounted to 865 Billions Norwegian kroner (NOK), which made up almost 45 per cent of total GDP.[3] Also, Norwegian exporters have faced rather volatile exchange rates since the mid 1990s and it is thus likely that Norwegian exporters have behaved accordingly in some manner. As evidence of exchange rate volatility effects on Norwegian exports so far is rare, this thesis contributes to the existing literature on several areas.

[2] See Boug *et al.* (2006) for details of the Norwegian exchange rate policy over the last decades.
[3] Figures are taken from http://www.ssb.no.

First, the export equations in the macroeconomic model MODAG used by Statistics Norway[4] do not include effects from exchange rate volatility on trade. Boug *et al*. (2002) describe the previous model for the machinery and equipment industry, a demand specified equation based on work by among others Reymert (1984), Lindquist (1995) and Naug (2002). In our analysis we add to this model the possibility of exchange rate effects. To this end, we construct a trade-weighted exchange rate in which EU, Sweden, US, Japan and the UK constitute the main trading partners. Given the trade-weighted exchange rate we extract several volatility measures using different known techniques from the literature. We succeed by employing the conditional variance from a Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model as a final choice of a volatility measure applied in the analysis.[5]

Second, we pay special attention to the construction of a new relevant proxy for the competing price facing Norwegian exporters abroad. So far in MODAG competitive prices have been proxied by import prices (denominated in the Norwegian currency) for the relevant good. We argue that such a proxy has some inappropriate implications from a theoretical point of view. Third, we also pay special attention to the time series properties of the data used. As could have been expected prior to the analysis based on previous findings, c.f. McKenzie (1999), the determination of order of integration for the volatility measure turns out least straightforward. The test results yield contradicting conclusions, and we therefore choose to consider both the case of nonstationarity, where the volatility measure is assumed to be integrated of order one, I(1), and the case of stationarity, I(0). In the latter case, the discussion of short run effects becomes relevant. In addition, we consider some alternative techniques recently developed in order to treat exogenous I(1) and I(0) variables in the long run relationships, see Harbo *et al.* (1998), Rahbek and Mosconi (1999) and Pesaran *et al.* (2000).

Finally, we pay attention to monetary policy regime shifts and special exchange rate events during the sample period. During the selected sample period the monetary policy in Norway has gone through two main regime shifts. In 1992, the monetary policy switched from a fixed exchange rate regime to a managed floating regime in which the Central Bank aimed at stabilizing the NOK within given target bands to the currencies of our main trading partners.

[4] See http://www.ssb.no/vis/english/research_and_analysis/models/main.html for more information about models at Statistics Norway.

[5] Bollerslev (1986) developed this method. In chapter 3, we present the theoretical concepts behind the GARCH model.

Then, in 2001, the monetary policy changed to inflation targeting and a freely floating exchange rate regime. We test by means of out-of-sample forecasting weather the latter regime change did indeed have significant effects on the exporters behaviour and thus on the parameters of the estimated model.[6] Also, we investigate in the same manner the possible effects from the Asian financial crisis in 1997.

The remainder of this thesis is organized as follows: In chapter 2 we discuss the underlying theoretical export model for the empirical analysis. In chapter 3 we present the data used in the analysis and investigate time series properties graphically. Herein we give a detailed description on how the index for the competing price facing Norwegian exporters abroad and the chosen measure for exchange rate volatility are constructed. In chapter 4 we review key concepts of univariate time series analysis and formally investigate the time series properties by means of unit root testing. In chapter 5, we review key concepts of cointegration analysis and report results from cointegration tests among the selected variables. In chapter 6, we conclude the empirical analysis by reporting and discussing an estimated parsimonious equilibrium correction model for the export volume. The thesis is brought to conclusion in chapter 7, where we provide a summary of the most important findings and point out potential caveats in the analysis. We also address some suggestions for future research on the topic.

[6] Due to too few observations, we are prevented from doing the same exercise to analyse any potential impacts of the former change in monetary policy.

2. The Export Model

Boug *et al.* (2002) present the original export equations used in the macroeconomic models of Statistics Norway. These export equations are typically demand type models and build on work by e.g. Reymerts (1984), Lindquist (1995) and Naug (2002). However, as noted in the introduction, these equations do not include any effects from exchange rate volatility on exports. Drawing on among others Arize (1995) and de Vita and Abbott (2004), a general demand type export equation including exchange rate volatility as one of the explanatory variables can be set up as follows:

(2.1) $X^e{}_t = f(P_t / P^*{}_t, Y^*{}_t, VOL_t)$

Here, export demand (X^e) is a function of relative prices between Norway and the abroad countries (P/P^*), foreign demand conditions (Y^*) and the exchange rate volatility (VOL). Also, in line with previous studies the restriction of homogeneity of degree zero for the equation in prices is imposed a priori.

Parameterizing and applying the notation for the proxy variables in the subsequent analysis, equation (2.1) can be written on the following form:

(2.2) $a_t = \beta_0 + \beta_1 (pa - pk)_t + \beta_2 wmi_t + \beta_3 Vol_t$.

Here a is the natural logarithm of the export volume, $pa - pk$ is the natural logarithm of the price ratio (with Norwegian prices denoted as PA, and foreign competing prices as PK), wmi is the natural logarithm of world market conditions (demand pressure) for Norwegian goods and Vol is the conditional variance from a GARCH(1,1) model of the first differences of the trade weighted Norwegian exchange rate in natural logarithm.[7] The discussion of selection, theoretical basis and construction of this volatility measure is devoted to chapter 3, along with a presentation of the other time series just mentioned.

[7] In what follows, lower case letters indicate natural logarithm of a variable, unless otherwise noted. The term "growth rate" refers to the first difference of the natural logarithm of a variable. The growth rate for the volume of exports is then given by $Da_t = \ln(A_t / A_{t-1}) = a_t - a_{t-1}$. A natural logarithm of a ratio, for instance the ratio between two prices PA and PK, is denoted as $pa - pk = \ln(PA / PK)$.

In the literature, different approaches have been taken to approximate relative prices between domestic and foreign markets. Égert *et al.* (2005) use domestic producer prices as proxy for import and export prices due to data constraints. Others use consumer prices, but then it is clear that such prices have elements not directly of relevance for the exporting sector. In this thesis we aim at approaching the real underlying theoretical variable. Hence we focus on constructing a consistent approximation of the world market prices facing Norwegian exporters of machinery and equipment.

From an economic point of view one could prior to the analysis form some expectations about the signs of the coefficients in (2.2). Standard theory predicts that price ratio effects should be negative and foreign demand pressure should affect export volume positively, hence $\beta_1 < 0$ and $\beta_2 > 0$. Arize (1990) presents theoretical discussions of these matters. As commented upon in the introduction, the sign on the last parameter in (2.2) is more ambiguous. Both negative and positive significant effects are discussed and found in the literature concerning the effects of exchange rate volatility on export performance. Hence, in this analysis the a priori expectation of the volatility parameter should be open, both negative and positive effects are likely to appear in the conducted analysis.

3. Data and Variable Descriptions

McKenzie (1999) stresses the relevance of working with disaggregated data when performing analysis of the impact of exchange rate volatility on export performance. As previously noted, the analysis in this thesis is conducted on data from the sector of machinery and equipment, one of the main industries in Norway. From here on, all references to variables of price and quantities refer to those of the machinery and equipment sector. The data set consists of quarterly unadjusted observations for the period 1985(1) - 2005(4). A list with data description and sources are found in Appendix A.[8] Lagged and differenced variables limit the estimation period somewhat, and the standard period for the models estimated is 1986(1) - 2005(4).

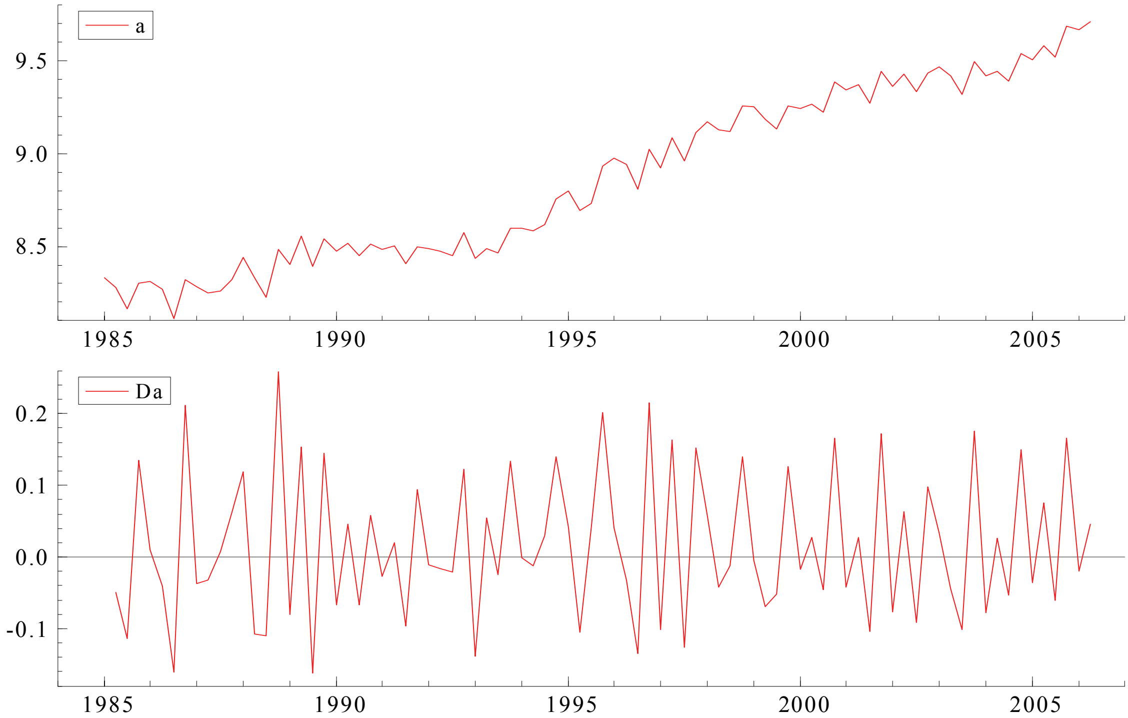
In the analysis, data for export volume, export prices for Norwegian exporters, a proxy for the world market price facing Norwegian exporters when competing abroad, the variable measuring the world demand conditions for Norwegian goods and an approximation for the volatility of the exchange rate are needed. Next we consider these variables in turn. This section also invites the reader to graphical inspection of the time series properties of the variables, whereas the formal treatment (by means of unit root testing) of these issues is left to chapter 4.

3.1 Export volume

Figure 3.1 plots the natural logartihm of the export volume and the quarterly growth. The level series develops with a positive trend throughout the period. In 1994 we might spot a break in the trend, leading to a steeper growth in the remaining period. The first difference of the series, on the other hand, looks stationary.

[8] A potential caveat in the analysis is the use of quarterly data. Most other related studies report the use of monthly data when investigating exchange rate volatility effects. In our analysis we are constrained to use quarterly observations, as the proxy variable for world demand conditions, *WMI* is only constructed quarterly by Statistics Norway. Replacing or reconstructing this variable to make it observable each month would make the analysis on monthly data possible.

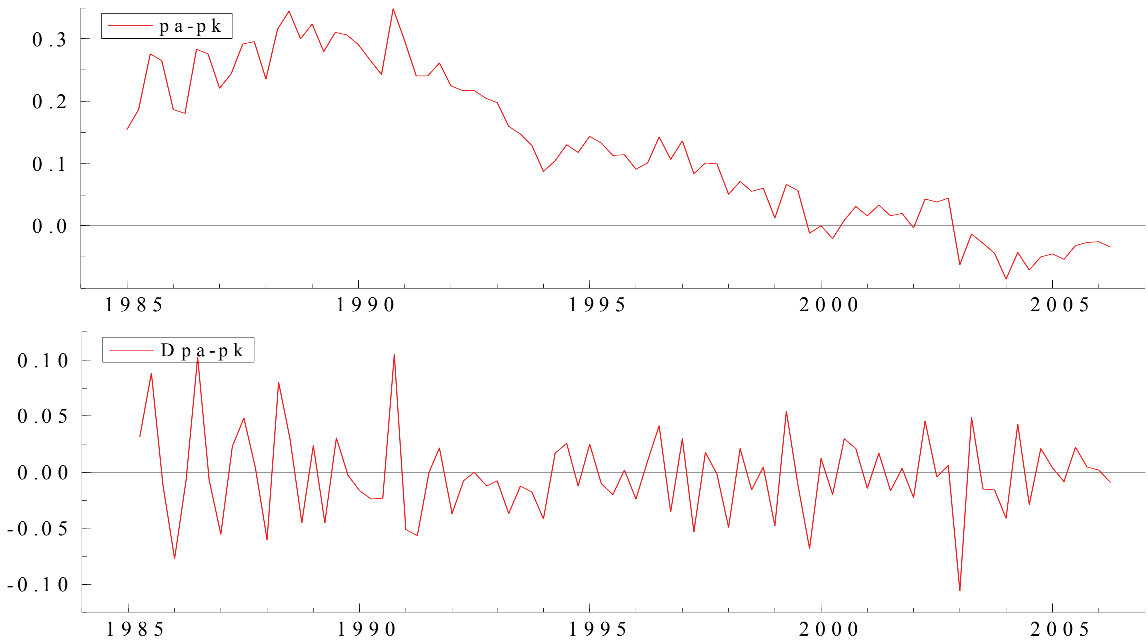
Figure 3.1: Export volume (*a*) and the quarterly growth (*Da*)



3.2 The ratio of Norwegian and world prices

Figure 3.2 plots the series of the relative price between Norwegian exporters and the competitive world market price (*PA/PK*), in level form and growth rate. The level series is characterized by a negative trend starting in 1991. Before this point, the series climbs slightly from its initial level, before hitting the top around 1991. The growth rate of the relative prices displays a quite stable pattern and seems to be stationary, although with possibly some larger fluctuations at the beginning of the period.

Figure 3.2: The ratio of Norwegian and foreign prices (*pa-pk*), and the quarterly growth (*Dpa-pk*)



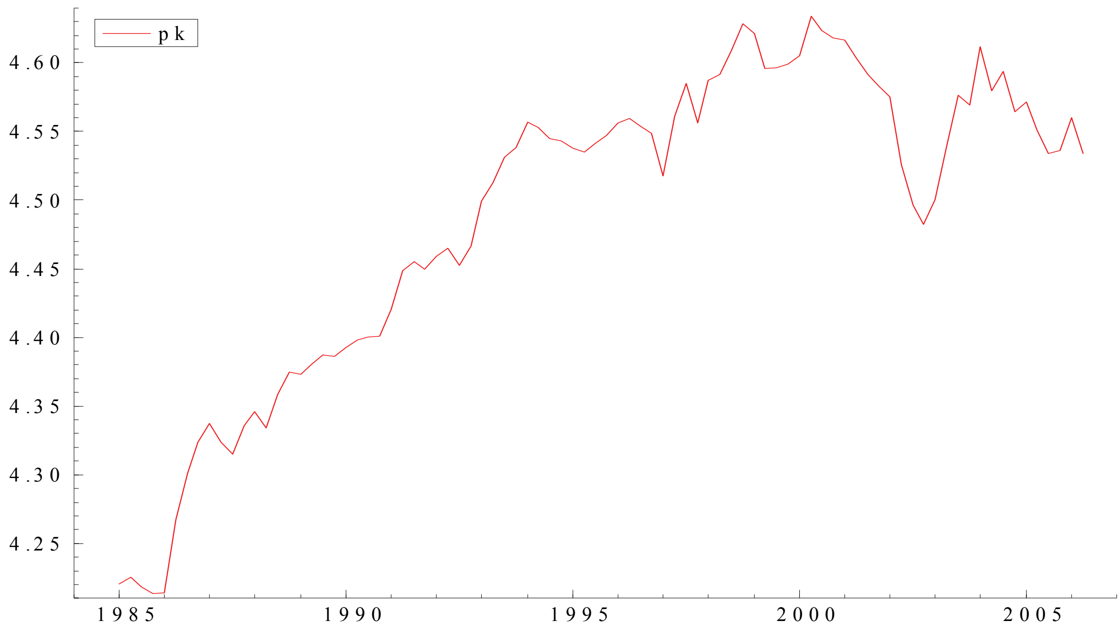
The two series underlying the price ratio are considered next.

3.2.1 The competitive world market price index

Figure 3.3 depicts natural logarithm of the constructed competitive world market price used in the denominator of ln(*PA*/*PK*), (*pa-pk*). It is steadily increasing throughout the first 15 years of the period. From 2000 onwards, the picture is changed towards a more stable pattern with a significant dip reaching the lowest point in the 3rd quarter of 2002. As there exists no official world market price for each and every good, the index *pk* has been constructed. As previously emphasised, attaining a high quality approximation of the theoretical competitive price for the specific good has been important prior to the estimation work, and may be considered as a contribution to existing materials related to the topic of export modelling of Norwegian industries. So far the import price in Norway has been used as an approximation for the competitive price in the macroeconomic models of Statistics Norway. Such practise is, however, not ideal from a theoretical point of view as it involves some problematic overall model properties. To be explicit, the import price for good *j* is modelled as a function of (among other variables) domestic costs (*c*) in Norway, $p^{i,j} = f(c,...)$. From standard theory an increase in Norwegian domestic costs,ceteris paribus, will increase the price on

Norwegian exported goods, thus reducing Norwegian exporters competitiveness as the ratio between export and world market prices deteriorates. But this cost increase also partially improves competitiveness for Norwegian exporters on the world market, as the world market price is approximated by the import price, and therefore directly affects the price ratio between Norwegian export prices and the world market price the opposite way. This is unreasonable, as a cost increase in Norway should only induce the pure negative effect on competitiveness from increased export prices, and not partially cancel out this negative effect through higher world market prices.

Figure 3.3: The competitive world market price index (*pk*)
for machinery and equipment



We now turn to the construction of the competitive price index and present some of its main underlying series and comment on some of their historical features. The index is based on the following general idea: Norwegian exporters face competition on the world market from exporters in other countries than the importing country. The world market price for Norwegian exporters for a good can then be considered to be a weighted sum of several import price indices in the countries to which Norwegian exporters are exporting goods. The theoretical formula for the competing world market price index for good *j* ($P^{w,j}$) may then be set up in the following way:

$$(3.1) \quad p^{w,j} = \sum_{i=Sweden}^{Japan} \left(pi^{i,j} \frac{NOK}{cur_i} \alpha_i \right),$$

where i ={ Sweden, the Euro zone, the UK, the US and Japan}, five of Norway's most important trading partners based on the α_i OECD-trade weights[9], and $pi^{i,j}$ is the import price index for good j in country i denominated in foreign currency. To convert the individual import price index to a common currency we multiply with the bilateral nominal exchange rates (NOK/cur_i).[10] The final competing price index is then constructed by weighting the different import price indices denominated in the Norwegian currency by the weigths a_i.

In the work of collection these import price indices one faces certain data constraints. Not always is it straightforward to find the relevant indices. For Sweden, the US, Japan and the UK this is more or less unproblematic, except for some complications due to different product specifications. For the Euro zone, however, we have not found the corresponding import price indices needed, and instead opted for the producer price index as the best alternative index. Norway and Sweden are among those countries using the official NACE classification system for economic activities in the European Community.[11] Thus, finding comparable data for Sweden is quite problem free.[12] Figure 3.4 depicts the import price index of machinery and equipment for Sweden denominated in both the Norwegian and the Swedish currency.

In November 1993 the Swedish Riksbank abandoned the currency peg and let the Swedish Krona (SEK) float with a huge depreciation as a result. From May 2000 to January 2003 the Norwegian Kroner (NOK) experienced a strong appreciation, followed by an almost equivalently large depreciation the years thereafter. Evidence from these events we

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[9] Excluding the smaller trading partners these weights are as follows: Euro zone: 49.5 per cent, Sweden: 21.4 per cent, UK: 13.3 per cent, Japan: 6.2 per cent and USA: 9.39 per cent. These are OECD weights used by Statistics Norway to construct the world market demand indicator for Norwegian goods, see Arbeids- og inkluderingsdepartementet (2006). Hence, there is consistency with respect to the weights used in the construction of both *wmi* and *pk* in this study.

[10] Noticeably, we transform all the price indices to the same base quarter, namely the first quarter of 2000, prior to converting the price indices to a common currency. All exchange rate data used in the analysis are taken from the Norwegian Central Bank, http://www.norges-bank.no.

[11] NACE: Statistical Classification of Economic Activities in the European Community, Eurostat: www.europa.eu.int/comm/eurostat/. Machinery and equipment is here labeled with the code "DK29".

[12] Statistics Sweden Online: http://www.scb.se and EcoWin (Reuters) database. Before the change to NACE in 1990, Statistics Sweden used their classification (SNI69). We have spliced the price index from the EcoWin database for the group of machinery and equipment with the price index from the previous system in the 1st quarter of 1990. In this previous system the group 38 "Verkstadvaror" seems to be the best alternative.

recognize in the NOK-denominated index. We can also see some signs of the large Norwegian devaluation in 1986 during the period of fixed exchange rate regime.

Figure 3.4: Import price index of machinery and equipment for Sweden in NOK and SEK

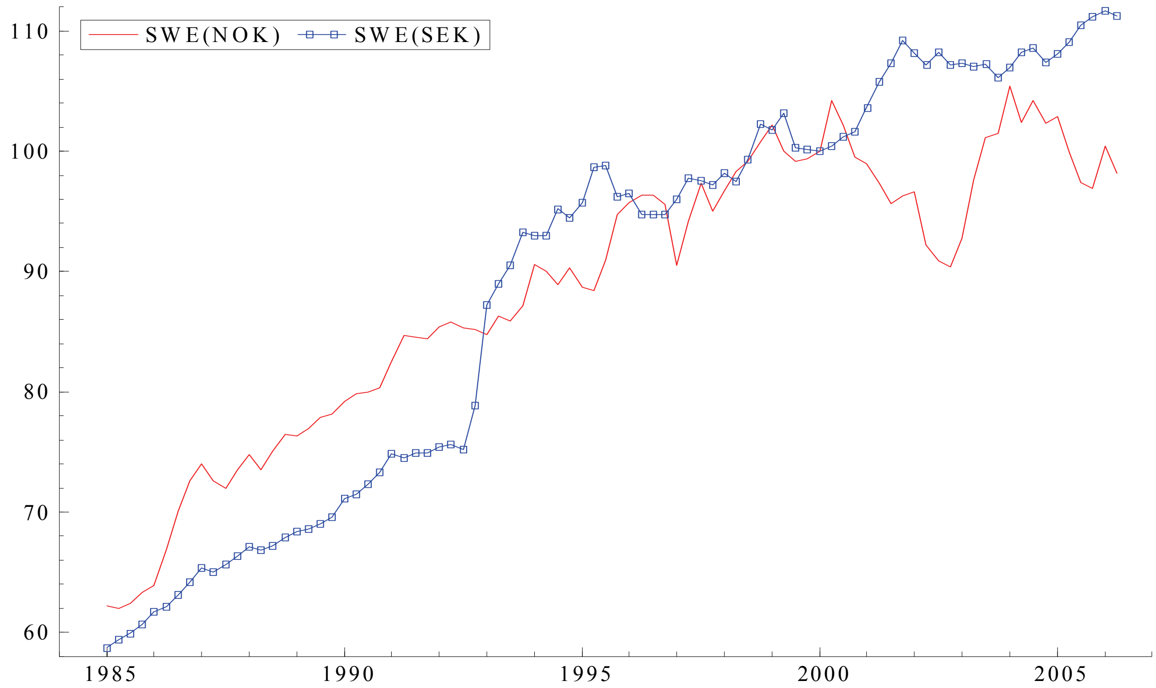


Figure 3.5 depicts the import price index of machinery and equipment for Japan denominated in both the Norwegian and the Japanese currency. The pattern of the Japanese import price index[13] is quite different from what was the case for the Swedish import price index . Throughout the period we recognize a sharp decline in the Japanese import price index. This becomes particularly clear when investigating the index denominated in Japanese Yen (JPY). Adjusting for the exchange rate between Norway and Japan, we see that much of the sharp decline early in the period is removed due to the Norwegian devaluation in 1986. From the 4th quarter of 2000 to 2nd quarter of 2003 the Yen depreciates by about 30 per cent against the NOK. This is also clearly identifiable from Figure 3.5.

[13] See the EcoWin (Reuters) database or Bank of Japan (www.boj.or.jp) for a more detailed description. They apply a slightly rougher definition of the sectors, but for machinery and equipment the definitions are almost exactly identical as those of Norway and Sweden.

Figure 3.5: Import price index of machinery and equipment for Japan in NOK and JPY

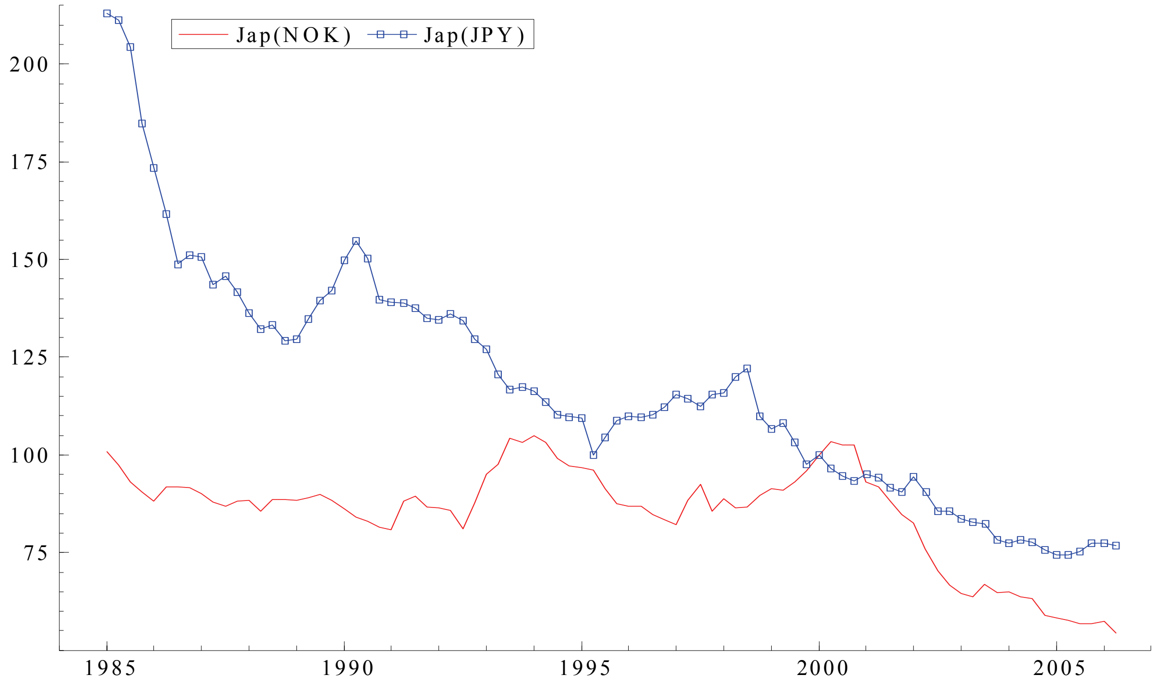


Figure 3.6 depicts the import price index of machinery and equipment for the UK denominated in both the Norwegian and the UK currency (GBP). In the UK the Standard International Trade Classification (SITC) generally is applied.[14] Similarly to the case of Japan, the import price index is declining towards the end of the period. Until 2001 the import prices in UK denominated in NOK develop with a positive trend. From 2001 onwards, however, the prices fall significantly resulting in a price index below its initial value at the end of the period. The NOK/GBP exchange rate evolves quite stable until 1997. In 1996 the import price index in UK starts to decline from its peak value. This may come from the increased share of computers and computer parts in the imports into the UK, goods that have had declining prices in recent years. When also adjusting for exchange rate movements, this price decline seems to be postponed by at least 5 years, thus the NOK-denominated index apparently lags the GBP-denominated index with some 5-6 years.

[14] SITC: Standard International Trade Classification, Revision 3, from the UN: http://unstats.un.org/unsd/cr/family2.asp?Cl=14. After comparing the classification tables of NACE and SITC, we have concluded that the SITC 71-77 index fits best for the purposes of the empirical analysis. Section 7 of the SITC is machinery and transport equipment, and a more detailed description of the sub selection 71-77 can be found at the UN web pages: http://unstats.un.org/unsd/cr/registry/regcs.asp?Cl=14&Lg=1&Co=7 .

Figure 3.6: Import price index of machinery and equipment for the UK in NOK and GBP

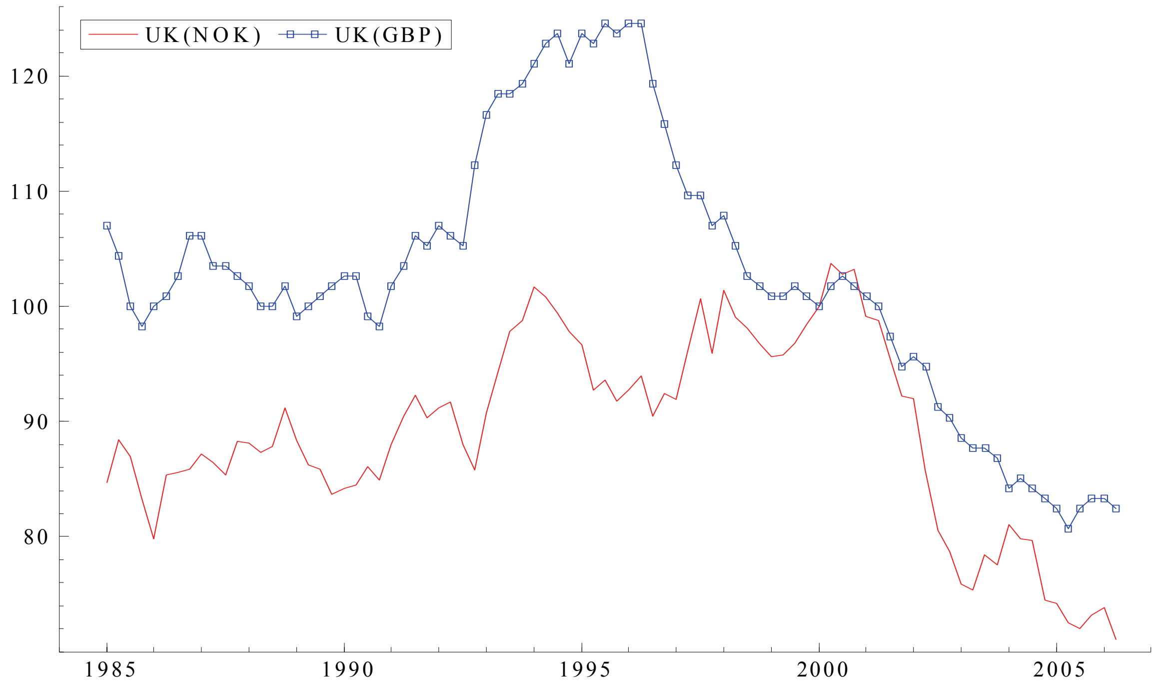
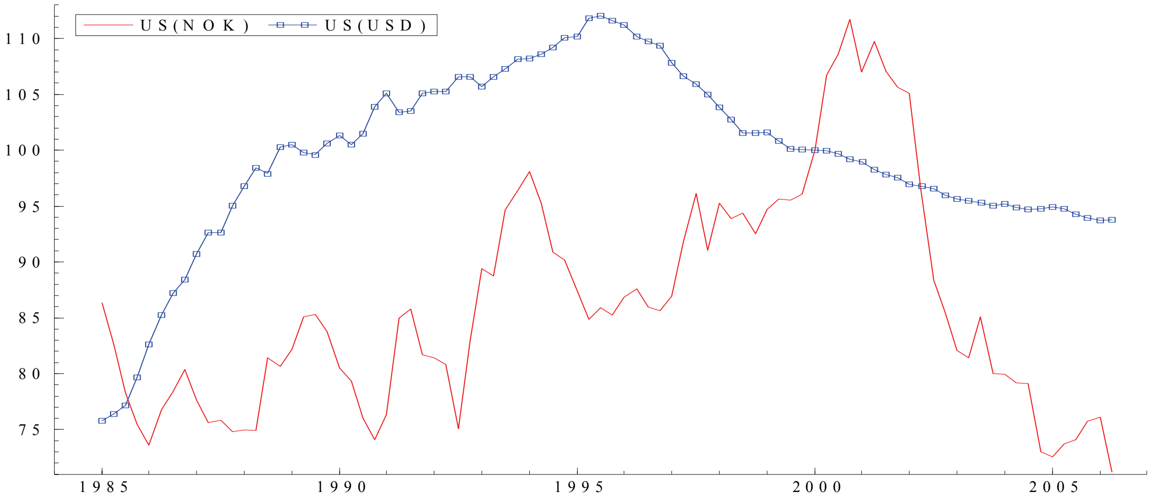


Figure 3.7 depicts the import price index of machinery and equipment for the US denominated in both the Norwegian and the US currency. From the Bureau of Labor Statistics we find the import price index for US imports.[15] In 1985 the USD peaks against the NOK, and in the following year the USD depreciates by almost 25 per cent against NOK. That explains the strong decline in the NOK-denominated index in the first year of the sample period. In the last quarter of 2000 the USD reaches another peak against the NOK (9,25 NOK/USD), but from here and until today the USD has continued to depreciate. These relatively heavy exchange rate fluctuations make the NOK-denominated index much more volatile than the original US import price index.

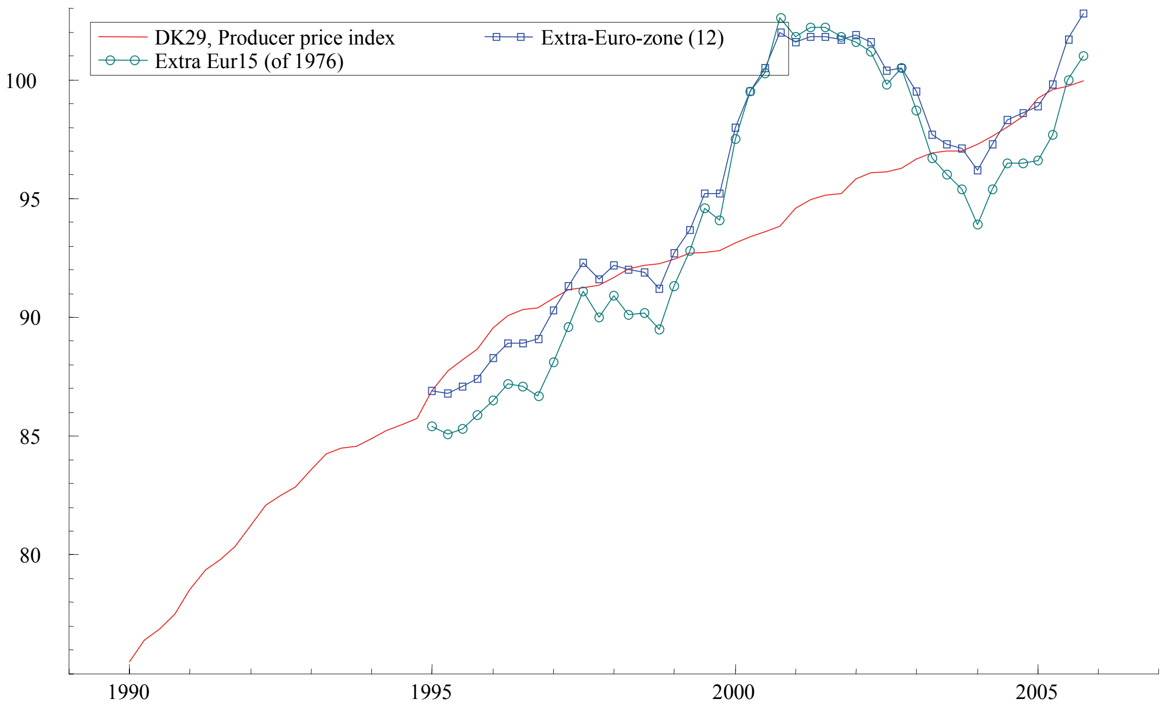
[15] See http://www.bls.gov. Here as in the UK the SITC is applied, and after comparing SITC with NACE classification we use the SITC 7 section, as the closest available dataset to machinery and equipment (DK29).

Figure 3.7: Import price index of machinery and equipment for the US denominated in NOK and USD



We faced some constraints when collecting data of competing prices for the Euro zone as the definition of the Euro zone itself and EU as a whole have changed throughout our sample period. From the Eurostat Comext database one can certainly find import price indices from intra EU trade, but these data only go back to 1995.

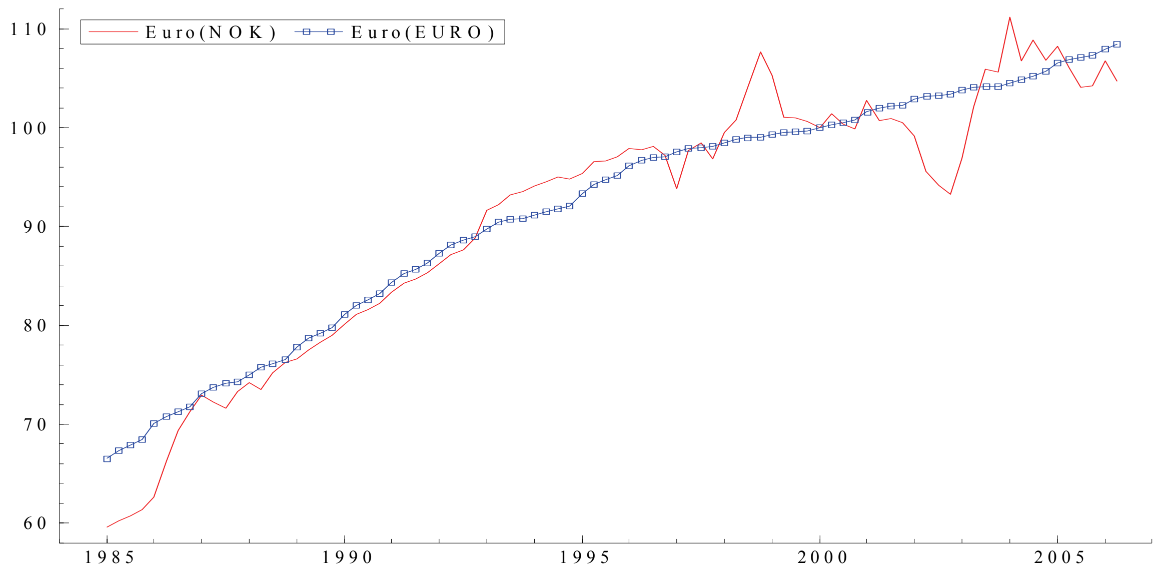
Figure 3.8: Import prices and producer prices of machinery and equipment in the EU area



Hence, we have chosen to use domestic producer prices as the closest possible approximation to import prices for the whole sample period in the case of the Euro zone.[16] Figure 3.8 depicts different available import price indices (EU 12 and EU15) and the producer price index of machinery and equipment for the entire Euro zone denominated in foreign currency. Although the import price indices exhibit larger fluctuations, they show evidence of a similar upward trend as the producer price index.

Figure 3.9 depicts the producer price index of machinery and equipment for the Euro zone denominated in both the Norwegian currency and the foreign currency. The EURO-denominated index shows a smooth positively trended path throughout the period, whilst the Norwegian currency denominated index exhibits some apparent fluctuations since the mid 1990s. Finding a proper nominal Euro/NOK exchange rate for the period of 1985 to 2006 is, however, not straightforward. In the analysis we use the variable called "KurvEcu" which is used as an exchange rate in the models of Statistics Norway. Before the introduction of the Euro it is based on the ECU and the currency peg used by the Central Bank of Norway. This is to some extent problematic as the Swedish Krona at early stages is counted into this ECU-exchange rate, and Sweden today is not part of the Euro zone. We choose, however, to use the KurvEcu variable for the whole period.

Figure 3.9: Producer price index of machinery and equipment in the Euro zone denominated in NOK and in Euro

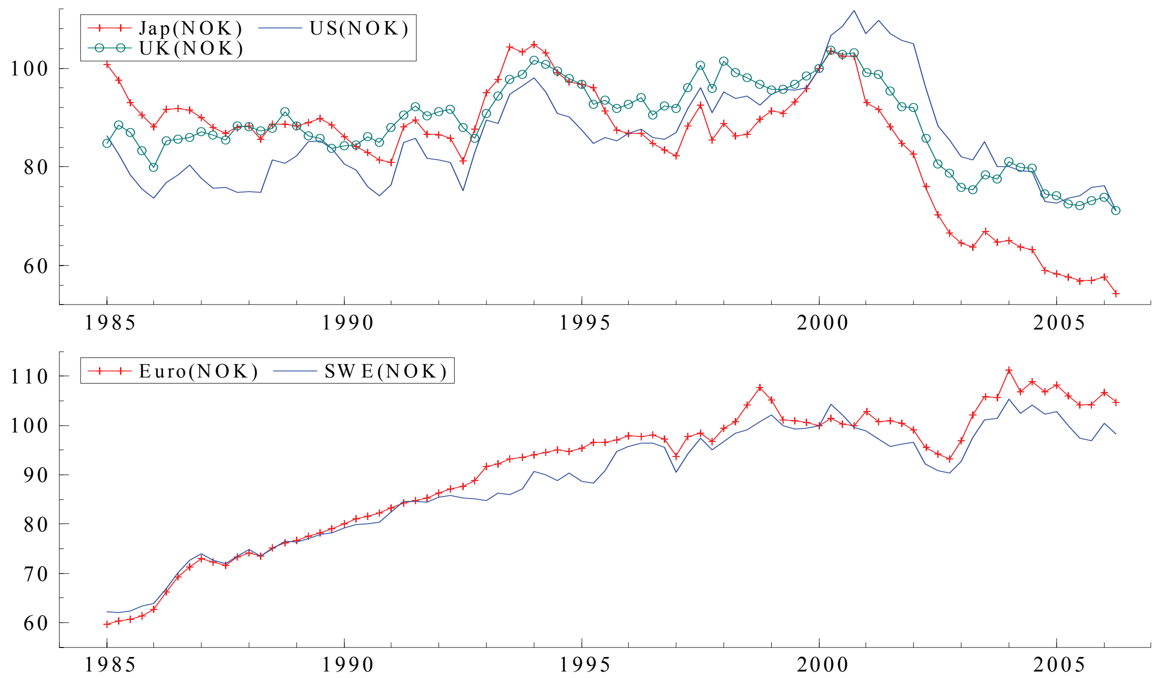


[16] Data are from the EcoWin (Reuters) database for NACE (DK29) machinery and equipment.

Some main features coming from exchange rate fluctuations stand out when inspecting the graphics of the Norwegian currency denominated producer price index. First the Norwegian devaluation in 1986 is apparent. Then, in the late 1990s and in the first years of the new millennium the exchange rate initially depreciated heavily, followed by an appreciation until 2002, before it moved towards average exchange rate levels at the end of the sample period.

Figure 3.10 depicts the different price indices for the main trading partners underlying the overall competing price index (*PK*), denominated in the Norwegian currency. As one may remark by reviewing the graphics, the five different indices can be recognized to follow two different patterns.

Figure 3.10: Unweighted price indices of machinery and equipment denominated in NOK



The indices from Sweden and the Euro zone behave similarly throughout the sample period. The remaining three indices from Japan, the UK and the US appear to follow a different pattern. These indices are initially declining, before they catch up and reach a last peak somewhere around the beginning of 2001. From here on and throughout the period, they experience a significant decline. These import price indices contain goods such as computers and computer parts, which have experienced a fall in prices in recent years. This can be one possible explanation of the two different patterns.

To weight the five NOK-denominated indices together we use the constant OECD trade weights discussed earlier. One may argue that each country's relative share is changing over time due to increased globalisation and fast growing economies, for instance Asian countries taking part in world trade. Using constant weights, however, seems like the most reasonable solution in the present context for two main reasons. The world market demand indicator constructed by Statistics Norway is based on these constant weights. Accordingly, we apply the same weights in the construction of the overall competing price index in order to achieve two consistent trade weighted explanatory variables in the regression models below. Also, using time varying weights may cause a data problem. If the trade weights change substantially from one period to another, this may result in changes in the price index not necessary caused by changes in the individual price index, rather the change in weights. To avoid this potential data problem one has to look at the change in each index in every quarter in order to weight them together properly. Thus, using constant weights in our context simplifies matters considerably.

Figure 3.11: The constructed world market price (*PK*) and the import price (*PI*) of machinery and equipment denominated in NOK

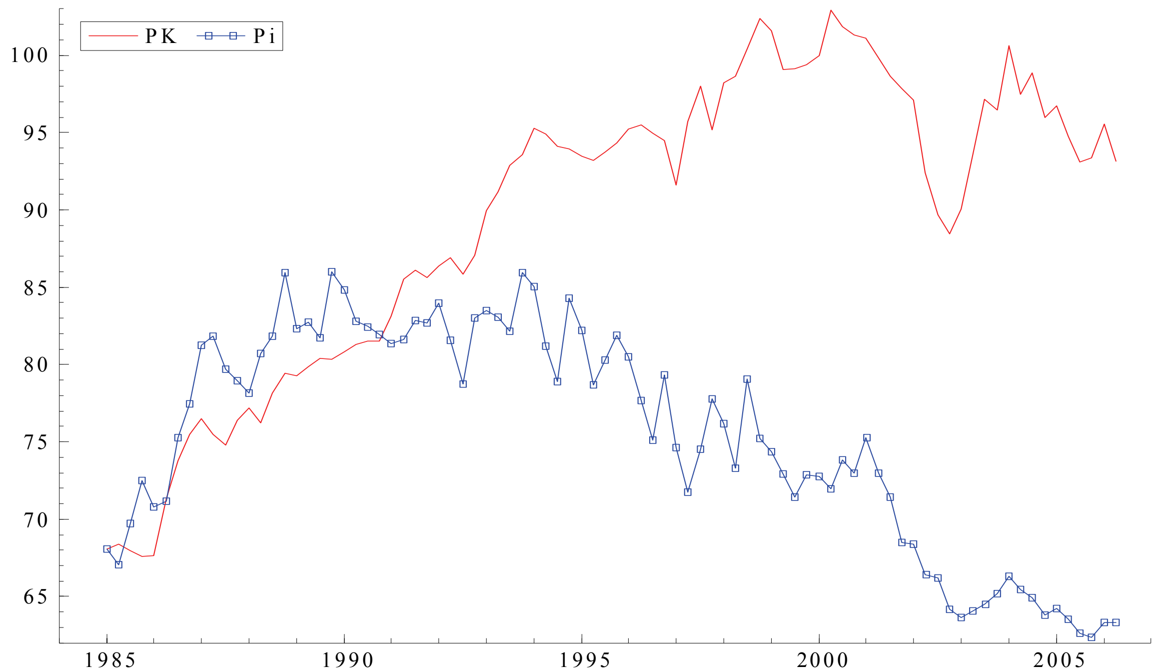


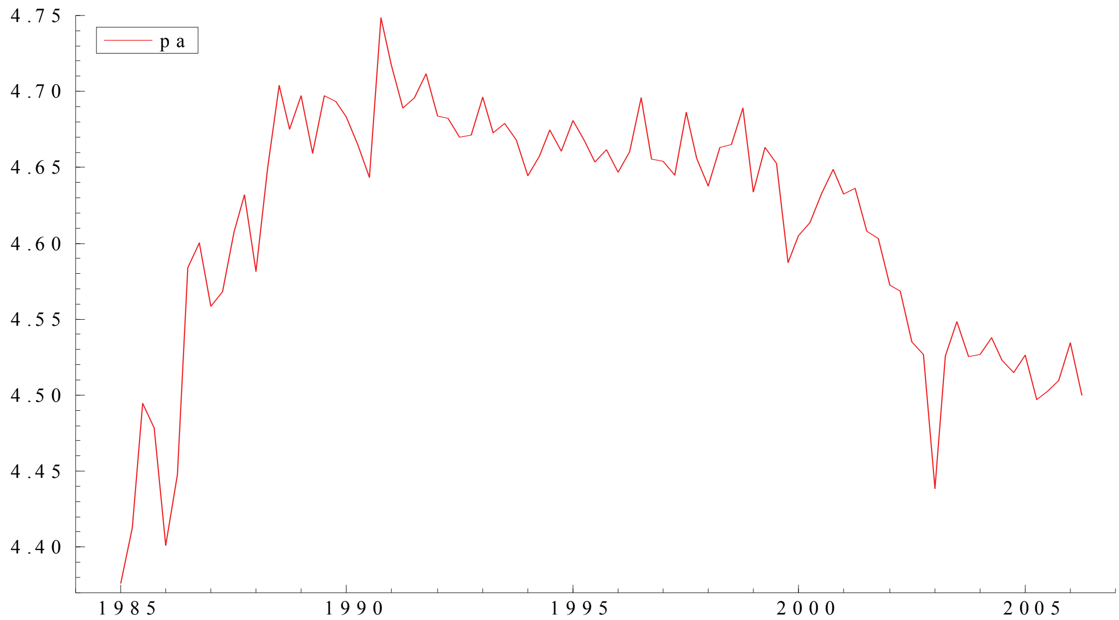
Figure 3.11 depicts the constructed overall competing price index (*PK*) together with the import price index (*PI*) previously used as a proxy for the competing price in the export equation for machinery and equipment. Both series are denominated in the Norwegian

currency. As Sweden and the Euro zone amounts to over 70 per cent in the weighting of the overall competing price, *PK* inherits the graphical features from these two trading partners individual price indices. The two indices develop quite similarly at the beginning of the period. Around 1992, however, the two indices split apart. From here on the import price index shows a clear downward sloping trend, while the competitive price index continues its upward trend. One explanation may be that the import price index for machinery and equipment contains some goods that are not relevant for Norwegian exporters. These may include computer and computer parts from low cost producing economies as Asian countries and others, goods that are known to have experienced falling prices over the last decade or so. This suggests that the import price index may not be a good approximation for the world market price, as Norwegian exporters generally do not trade a lot of such goods. Hence, the Norwegian export industry of machinery and equipment has experienced less deteriorating world market prices when using the constructed *PK* series instead of the import price series *PI*.

3.2.2 Norwegian export prices

The export price series used in the empirical analysis is calculated as the total value of Norwegian exports of machinery and equipment divided by the volume of the corresponding exports. Figure 3.12 plots the natural logarithm of the export price index. With a sharp increase in the late 1980s, the index declines steadily throughout the remainder of the sample period.

Figure 3.12: Norwegian export prices (*pa*) denominated in NOK

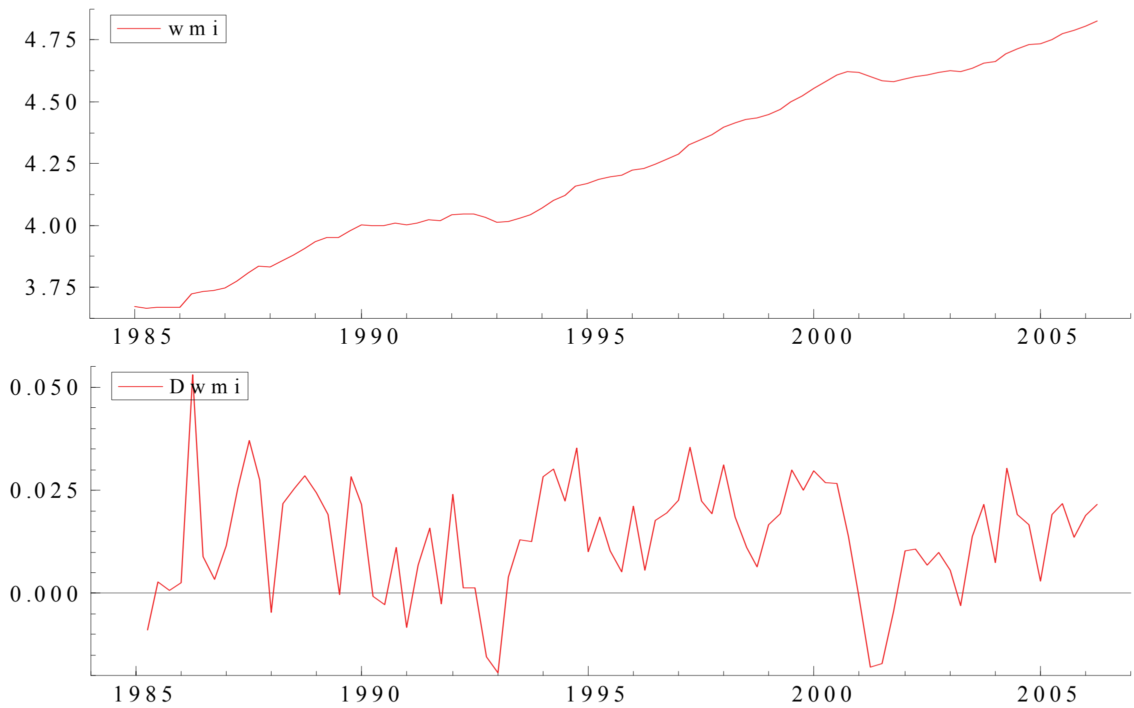


3.3 World market demand indicator

As a proxy for foreign demand conditions we use the world market demand indicator constructed by Statistics Norway. The world market indicator is meant to catch movements in international demand conditions for Norwegian exports. Using the GDP development in the five main trading partners discussed previously, and weighting with the corresponding weights, we obtain the measure of the world market conditions.

Figure 3.13 plots the natural logarithm of the world market indicator (*wmi*) and the quarterly growth (*Dwmi*). We see that the indicator grows steadily over the whole sample period, with to main drops in the years 1992 and 2001.

Figure 3.13: The indicator of world market demand (*wmi*) and the quarterly growth (*Dwmi)*



3.4 Exchange rate volatility

We have seen above that the exchange rates tend to make prices more volatile when converted to the Norwegian currency. Thus, for a Norwegian exporter exchange rates may affect the profits directly through their variations. If a contract promises to pay 1 million € in one year from now, the exporter knows how many euros this is, but the pay in Norwegian kroner depends on the exchange rate. If the NOK appreciates by 10 per cent, say, this will reduce the pay in NOK by 10 per cent. On the other hand, a depreciation of the NOK by the same amount will give an extra pay of 10 per cent in NOK. The former is what the exporter fears and considers as a risk in her decision-making. Having a fixed exchange rate removes this risk; hence being part of the Euro zone would completely remove the exchange rate risk when trading with other countries inside the Euro zone. Without taking part in a monetary union with fixed exchange rates, the exporter may want to hedge herself against such risk in the financial markets. In the simplest case the Norwegian exporter could make an agreement with a company within the Euro zone. Assume that a German company expects a payment of 8 million NOK in one year. Thus, agreeing on trading those 8 million NOK against the 1 million Euros ensures both companies to get a fixed pay in one year from now in their respective currencies. Hence, the exchange rate risk is completely removed by such an agreement. In this simple world, we can assume that the current exchange rate today is 8 NOK/Euro, and that there is a 50/50 chance that the exchange rate will either be 10 per cent lower or higher one year from now. In reality the deal between the two companies is nothing more than swapping the potential upside from a 10 per cent depreciation in their own currency (NOK for Norwegian exporter) against the potential downside of a 10 per cent appreciation. The example above illustrates a case with perfect hedge. Of course in the real world perfect hedge is hardly attainable for a variety of reasons. Even in highly developed financial markets, finding the perfect contract to remove such exchange rate risk is hard, both with respect to the volume and time horizon. Also, the hedging in itself tend to have costs, which may affect the trade directly in a negative way.

The above example rests on one important underlying assumption about the companies, namely that they are risk averse. A risk averse agent prefers the certain for the uncertain. Theoretically this is represented by a concave utility function. Imposing other properties on the agents' preferences would then lead to other conclusions. Considering trade as an option, we know from standard finance theory that increased volatility in the underlying security will increase the value of an option. From a theoretical point of view one would then expect positive impact of exchange rate volatility on trade flows. As noted in the introduction, existing literature has found evidence of both negative and positive effects from exchange rate volatility on trade. We proceed to consider some possibilities of how to measure exchange rate volatility in the present context. Specifically, to allow for comparability of performance of different volatility measures and as a robustness analysis, we here consider two alternative measures of volatility based on the nominal trade weighted exchange rate discussed above.[17] We present these two measures in turn.

In line with Koray and Lastrapes (1989) and Arize *et al.* (2000) among others, we first experimented with a measure of the moving average of the standard deviation of the nominal trade weighted exchange rate, which may be formulated as follows:

[17] We have also experimented with measures of exchange rate volatility based on a trade weighted real exchange rate. Although the distinction between real and nominal exchange rate volatility has generated a lot of debate in the literature, see McKenzie (1999), the empirical results obtained here suggest that this distinction dos not impact significantly on the results. Since the trade weighted nominal and real exchange rates in our case have moved closely together during the most part of the sample period, the distinction between nominal and real volatility makes no difference to the results obtained. Hence, we only report results of the impact of exchange rate volatility on Norwegian exports based on the nominal trade weighted exchange rate.

$$J_{t+m} = \left(\frac{1}{m} \sum_{i=1}^{m} (e_{t+i-1} - e_{t+1-2})^2 \right)^{1/2}, \quad (3.2)$$

where e is the natural logarithm of the exchange rate and m is the order of the moving average. To avoid an arbitrary choice of the order of the moving average, we tried different values of m ranging from 1 to 4. However, it turned out that the use of the volatility measure in (3.2) for different values of m produces similar results as the other measure of volatility considered here, namely a GARCH based volatility measure. In what follows, we therefore only report estimation results based on the latter volatility measure.[18] Below we present the theory behind the GARCH model. Then, we display the exchange rate volatility time series obtained from applying that model to our data set.

Again defining e_t as the natural logarithm of the nominal Norwegian exchange rate at time t, we assume that the exchange rate follows a random walk:

$$e_t = e_{t-1} + \varepsilon_t,\ e_t \sim N(0, \sigma_t^2). \quad (3.3)$$

The variance of the error term is here time dependent, and with the GARCH approach we can now describe how this variance changes over time. In the GARCH model developed by Bollerslev (1986) the variance is given by

$$\sigma_t^2 = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \ldots + \gamma_p \varepsilon_{t-p}^2 + \beta_1 \sigma_{t-1}^2 + \ldots + \beta_q \sigma_{t-q}^2, \quad (3.4)$$

where $\gamma_0 > 0$, $\gamma_j \geq 0, j = 1,\ldots,p$, $\beta_j \geq 0, j = 1,\ldots,q$.

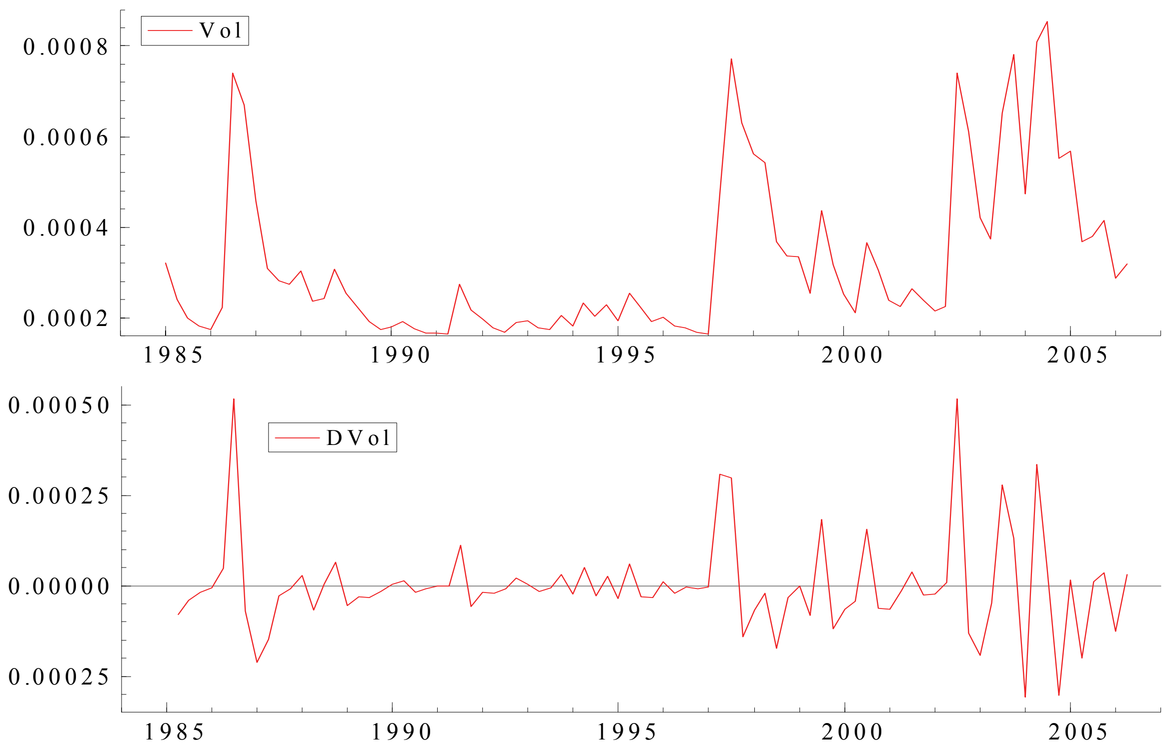
Bollerslev (1986) extended Engle's (1982) ARCH model by including the variance own past as explanatory variables in addition to the squared error terms. The simplest variant of the model above, a GARCH (1,1) turns out to be one most widely used:

$$\sigma_t^2 = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \beta_1 \sigma_{t-1}^2 \quad (3.5)$$

[18] Engle (1982) developed the Autoregressive Conditional Heteroskedastic (ARCH) model, which was later extended by Bollerslev (1986).

In our analysis the model in (3.5) is applied, which produces the exchange rate volatility measure depicted in Figure 3.14. Except for a peak early in the period (due to the 1986 NOK devaluation) we clearly see the pattern of increased exchange rate volatility towards the end of the sample period. Bearing in mind the changes in monetary policy discussed earlier this may not come as a surprise. As may also become clear when considering the graphics is the different time series properties of the exchange rate volatility compared to the properties of the other time series presented and discussed above. Whereas the export volume, price indices and the world market demand indicator show signs of clearly trended behaviour, the pattern of the volatility appears more diffuse. These properties are in particular treated and commented upon in the next chapter, which is devoted to the unit root analysis.

Figure 3.14: The exchange rate volatility based on a GARCH(1,1) model



4. Unit Root Analysis

By graphical inspection of the time series presented in the preceding chapter, some features stand out. At first sight the series for export volume and the world demand indicator clearly show positively trended paths, while the price ratio seemingly falls throughout the sample period, at least after an initial increase until around 1990. The volatility measure may better be described as clusters of high and low value observations, without any distinguished direction of movements or apparent trend directions. One may, however, suspect a more dense concentration of high value observations after the mid 1990s following the periods of managed and freely floating exchange rate regimes in monetary policy in Norway. More formal analysis of these properties by means of unit root testing is important to conduct prior to econometric analysis of the data.

In this section we first review some of the key concepts of the theory concerning univariate time series analysis. Thereafter we conduct unit root testing using augmented Dickey-Fuller test, see Dickey and Fuller (1981) and Hendry and Doornik (2001), to more formally examine the properties of the time series.

4.1 Concepts of Order of Integration

A time series process a_t is said to be weakly stationary or covariance stationary if neither the mean nor any of the auto covariances are time dependent, e.g. Hamilton(1994), that is $E(a_t) = \mu_a$ and $E(a_t - \mu_a)(a_{t-j} - \mu_a) = \gamma_j$ for all *t* and *j*.

The random walk of (the natural logarithm of) the exchange rate given in equation (3.3) is an example of a nonstationary process. We repeat equation (3.3) here for the sake of convenience:

(4.1) $\quad e_t = e_{t-1} + \varepsilon_t, \; e_t \sim N(0, \sigma_t^2)$.

A random walk of this kind is nonstationary with a constant mean of zero, but time dependent auto covariances. Each of the innovations in the white noise process will have infinite lasting effect on e_t.[19] Differentiating e_t once yields a process that is stationary. Generally, a nonstationary process which is differentiated *n* times to become stationary is said to be integrated of order *n*, I(*n*).

Doing regression analysis with nonstationary variables may create spurious regressions, i.e. estimated regression coefficients may tend to be small in relation to their standard errors, see e.g. Murray (1994) for a discussion. A solution to this problem could be to perform a regression analysis based on changes in the variables, as suggested by Granger and Newbold (1974). If all the variables included in the analysis are integrated of order one, differentiating them once until they are I(0) ought to solve this issue. This procedure, however, may result in loss of important long run information or equilibrium-correction mechanisms[20] among the variables. For instance consumption and income are typically nonstationary processes on their own, but in the long run, consumption may be assumed to roughly amount to a constant share of income. Doing regression on the first differences alone would then ignore this long run relationship, hence resulting in a miss specified regression. Variables, which together exhibit these kinds of long run relations, are referred to as cointegrated variables, see e.g. Engle and Granger (1987). Hence, to perform an adequate econometric analysis of differentiated stationary variables, also the cointegration relationships (if any) between the variables in levels should be included. The theory of cointegration and its applications is discussed in more detail in chapter 5.

4.2 Testing for Unit Roots

Since the variables displayed graphically in chapter 3 apparently contain characteristics that violate the principles of stationarity, it is crucial to investigate these properties more thoroughly. One popular method of analysing time series properties is the Dickey-Fuller

[19] With initial value e_0 we can rewrite the random walk with infinite past as just a sum of e_0 and the infinite sequence of error terms: $e_t = e_0 + \sum_{i=1}^{\infty} \varepsilon_{t+1-i}$.

[20] In the literature, the concept of an *equilibrium correction mechanism* is also known as an *error correction mechanism*. In this thesis I consequently refer to this concept as *equilibrium correction*. The difference between the two lies only in the name, and for instance Lütkepohl (2005) utilizes the description *error correction*, page. 87.

Unit root test. To describe this method in its simplest form we consider the following AR(1) process:

(4.2) $$a_t = \beta_0 + \beta_1 a_{t-1} + \varepsilon_t ,$$

where $\varepsilon_t \sim (0, \sigma^2)$. Subtracting a_{t-1} on both sides yields:

(4.3) $$\Delta a_{t-1} = \beta_0 + (\beta_1 - 1)a_{t-1} + \varepsilon_t = \beta_0 + \gamma_0 a_{t-1} + \varepsilon_t .$$

With $\gamma_0 = 0$ (or correspondingly $\beta = 1$), a_t becomes a random walk with drift, which is nonstationary. The hypothesis that $\gamma_0 = 0$ is often referred to as the *unit-root hypothesis.*
It is possible to test this hypothesis, using the associated t-statistic from an OLS regression and comparing it with values from the Dickey-Fuller distribution.[21]

To improve the model diagnostics of the estimated equation in the case of autocorrelation or non-normality in the residuals, one may apply augmented Dickey-Fuller (ADF) tests[22] by adding *p* lagged differences to (4.3). Here one may also test for the possibility of trend stationarity by including a deterministic trend, t:

(4.4) $$\Delta a_{t-1} = \beta_0 + \gamma_0 a_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta a_{t-i} + At + \varepsilon_t$$

The standard procedure is then to formulate the H_0 hypothesis that the process actually contains a unit root ($\gamma_0 = 0$). The testing decision is still based on the t-statistic associated with the coefficient of the lagged variable (γ_0), but now with somewhat different critical values.[23] Unable to reject the formulated *unit-root hypothesis*, we conclude that the series contain a unit root, and we may go on to investigate the first differences to check weather differentiation is necessary to achieve stationarity. Also, non-rejection of the unit-root hypothesis for the differentiated variable would indicate that the true order of integration is two or larger, in short form at least I(2).

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[21] For critical values, see e.g. Hamilton(1994), Table B.6.
[22] See Hendry and Doornik (2001).
[23] These critical values are dependent on weather or not the trend is included in (4.3). For both the case with and without the trend these critical values are found in Hamilton (1994), "case" 2 and 4.

We now employ the augmented Dickey-Fuller (ADF) tests to determine the order of integration for the variables used in the empirical analysis. For lag determination in equation (4.4) we have both considered the lag reduction method[24] and the Akaike information criteria.[25] The lag reduction method consists of estimating equation (4.4) with a starting number of lags, *p*. Considering the significance of the coefficients for the lags estimated by means of ordinary F-tests then performs the lag selection. The lag order is set to the number of the highest significant lag. In a model with *p* lags, the highest potential lag selection is then *p*. In cases where the lag reduction method and the Akaike information criteria differ in lag selection, we have reported both statistics. As some of the series also look to be strongly trended we have investigated the possibility of a significant trend in equation (4.4). For all variables except the volatility measure, the trend is found significant and thus included in the tests. The statistics t_t and t_c are the t-statistics of the coefficient γ_0 in an estimated model with and without the trend term, respectively.

Table 4.1: Unit root tests

Variable	Lag length from AIC	Lag length from lag reduction	t_c statistic	t_t statistic
a	5	5		-2.301
Da	4	4	-3.267*	
pa-pk	5	5		-3.504*
D(pa-pk)	4	1	-3.434*/-10.96**	
wmi	1	1		-2.209
Dwmi	0	0	-5.925**	
Vol	0	2***	-3.559**/-2.709	
DVol	1	1	-8.802**	

Notes: The *Ho* hypothesis is that the series do contain a unit root (that is $\gamma_0 = 0$). ** indicates rejection of *Ho* at the 1 per cent level, * indicates rejection at the 5 per cent level and *** indicates that the lag length is significant at the 10 per cent level with the lag reduction method. t_c and t_t are the reported t- values of the γ_0 in an estimated model of (4.4) with and without the trend term, respectively.

Table 4.1 reports unit root tests of our selected variables. The tests show that the export volume and the demand indicator are clearly I(1). The case of the relative prices is, however,

[24] See e.g. Wolden Bache (2002).
[25] See Judge *et al.* (1985).

a borderline case. Strictly speaking, on a 5 per cent level the hypothesis that *pa-pk* contains a unit root is rejected. At the 1 per cent level, however, it is not. The reason can be found when investigating Figure 3.2. Before 1991 the price series exhibits an increasing pattern. Only from this point the series starts to decline steadily. We choose to treat *pa-pk* as an I(1) series in the continuing analysis.

The consideration of the unit root hypothesis for the exchange rate volatility measure is even more unambiguous. In Table (4.1) the Akaike information criteria and the lag reduction method yield two different lag suggestions. These two suggestions also lead to different conclusions concerning the order of integration of the volatility measure, that is either I(0) or I(1).[26] In the following empirical analysis both cases are considered in turn.

[26] As underlined by McKenzie (1999), doubt about the order of integration of exchange rate volatility measures is typically reported in existing studies.

5. Cointegration Analysis

In this section we present main theoretical concepts of cointegration and conduct several cointegration analyses using the multivariate method suggested by Johansen (1988, 1995) under different assumptions about the time series properties of the exchange rate volatility measure. First, we consider the case in which the export volume, the price ratio, the world market demand and the exchange rate volatility all are treated as endogenous I(1) variables in the VAR model underlying the cointegration analysis. Since it is hard to establish significant long run effects from the exchange rate volatility in this case, we question its endogeneity status and next consider the case in which the exchange rate volatility is treated as an exogenous I(1) variable. Again, finding no significant long run impact of the exchange rate volatility on exports, we then perform a cointegration analysis treating the exchange rate volatility as an exogenous I(0) variable. However, inclusion of the volatility variable in this manner does not provide significant information to the VAR model. As a final case, we therefore remove the volatility variable from the information set and conduct a cointegration analysis with the three remaining variables, namely the export volume, the price ratio and the world market demand entering endogenously in the VAR model. We find evidence supporting the hypothesis of one unique cointegrating vector among the selected variables, a cointegration relationship that is in line with the standard demand type model of exports.

5.1 Concepts of Cointegration

Murray (1994) describes the concept of cointegration in a less formal manner by means of an example in which a drunk and her dog may form a cointegrating relationship. On her own, the drunk coming out of the bar late at night and then starting to wander off might be considered a nonstationary and unpredictable process. The same could be said about her dog, which she left before entering the bar. But together they might exercise a cointegration relationship, if we assume that there exists some sort of error correction mechanism between the two. Lets say, she regularly calls out to approximate how far away her dog is, and it responds with barking. Both the drunk and the dog do not want the other to drift to far away, forcing them to adjust their paths when this is about to happen. As such, the drunk and her dog may form a cointegrating relationship.

Drawing closely on the review in Lütkepohl and Krätzig (2004) and Lütkepohl (2005), we continue to present the main theoretical concepts underlying the cointegration analysis conducted here. Let y_t be a vector of *K* time series variables, $y_t = (y_{1t},...,y_{Kt})'$. Using these *K* variables in a VAR model of order *p* then yields (ignoring any deterministic terms for simplicitly[27]):

(5.1) $y_t = A_1 y_{t-1} + ... + A_p y_{t-p} + u_t$.

Here u_t is a $(K \times 1)$ vector of unobservable time invariant error terms and with positive definite covariance matrix ($E(u_t u'_t) = \Sigma_u$), hence $u_t \sim (0, \Sigma_u)$. And ($A_1,.., A_p$) are the $(K \times K)$ coefficient matrices for the VAR(*p*) process. For the purpose of cointegration analysis a VEqCM (vector equilibrium correction model) can be obtained by subtracting y_{t-1} from both sides of the VAR model in levels. Doing so and rearranging terms yields

(5.2) $\Delta y_t = \Pi y_{t-1} + \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{p-1} \Delta y_{t-p+1} + u_t$.

Working through the exact calculations would then identify the Π and Γ_i matrices as $\Pi = -(I_K - A_1 - ... - A_p)$ and $\Gamma_i = -(A_{i+1} + ... + A_p)$ for $i = 1,..., p-1$, respectively.
If all of the *K* variables in the system are at most I(1) variables, it follows by definition that the expression Πy_{t-1} must be I(0). The matrix Π is here a $(K \times K)$ matrix. Suppose now that the rank of Π is *r*, then Π can be written as a product of $(K \times r)$ matrices α and β such that $\Pi y_{t-1} = \alpha\beta' y_{t-1}$. Returning to the above discussion regarding cointegrating relations, the rank *r* of Π is just the number of linearly independent cointegration relations among the *K* variables in y_t. Let us look at an example using the variables from the export model (2.2). Assuming two cointegration relations among the four variables in the model, that is rank(Π)=2, we may express Πy_{t-1} as follows:

[27] See Lütkepohl and Krätzig (2004) for a detailed treatment of these and other cases.

$$\Pi y_{t-1} = \alpha\beta' y_{t-1} = \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \\ \alpha_{41} & \alpha_{42} \end{pmatrix} \cdot \begin{pmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} \end{pmatrix} \cdot \begin{pmatrix} a \\ pa - pk \\ wmi \\ Vol \end{pmatrix}_{t-1}$$

(5.3)

$$= \begin{pmatrix} \alpha_{11} eqcm_1 + \alpha_{12} eqcm_2 \\ \alpha_{21} eqcm_1 + \alpha_{22} eqcm_2 \\ \alpha_{31} eqcm_1 + \alpha_{32} eqcm_2 \\ \alpha_{41} eqcm_1 + \alpha_{42} eqcm_2 \end{pmatrix}_{t-1} .$$

The *eqcm* terms are here the two different cointegrating relations between the variables:

(5.4) $eqcm_i = \beta_{1i} a + \beta_{2i} (pa - pk) + \beta_{3i} wmi + \beta_{4i} Vol \ , \ i = 1,2 .$

Deviations from the equilibrium in the cointegration relations feed directly into the system and affect the variables modelled on first differences through the matrix of alpha coefficients. The magnitude of the different alpha coefficients indicates by how much such a deviation in each of the two long run relations affects each of the modelled variables. This is exactly the long run effect (equilibrium correction mechanism) discussed previously; potentially missing out if the variables were to be modelled on first differences solely. The effect of the equilibrium correction mechanism becomes clarified when considering the special case where the rank of the Π matrix is just equal to zero. In this case the equilibrium correction term vanishes, and we are left with a model in first differences only. The opposite special case exists when all variables are I(0) and $r = K$, which then would imply a stationary VAR model in levels.

Generally speaking, for any non-singular matrix Q of dimension $r \times r$, we have that

(5.5) $\Pi = \alpha\beta' = \alpha Q Q^{-1} \beta' = \alpha^* \beta^{*\prime}$ where $QQ^{-1} = I_{(r \times r)}$.

Hence, without any further identifying restrictions, the cointegration relations are not unique. We will return to the identification problem when dealing with the specific empirical models in our case. It follows from the discussion above that the determination of the cointegration rank r is important in order to specify the correct model. Johansen (1988, 1995) has developed a method to determine the cointegration rank. The likelihood ratio (LR) statistic for comparing a specific cointegration rank $r = r_0$ against a larger rank of cointegration $r = r_1$ can be set up as follows:

(5.6) $\lambda_{LR}(r_0, r_1) = 2(\ln l(r_1) - \ln l(r_0))$,

where $l(r_1)$ and $l(r_0)$ are the maximum values of the likelihood function for cointegration rank r_1 and r_0, respectively. Johansen (1988, 1995) derived the asymptotic distributions of these *LR* statistics, which also are referred to as the trace test statistics. These statistics are compared to specific critical values, which may be taken from e.g. Doornik (1998).[28] The order of the rank is then determined in a testing sequence, which may be described as follows:

(5.7) 1st step: $H_0 : rk(\Pi) = 0$ versus $H_1 : rk(\Pi) > 0$

If H_0 is not rejected, the rank is set to zero and a VAR model on first differences would be appropriate. However, if H_0 is rejected, one proceeds to the 2nd step of the testing sequence:

(5.8) 2nd step: $H_0 : rk(\Pi) \leq 1$ versus $H_1 : rk(\Pi) > 1$

We use the same procedure as in the 1st step and continue to the next step in case of rejection of H_0. If H_0 is not rejected, stop the testing sequence and use the obtained rank. The testing sequence ends when the rank under H_1 reaches the dimension (*K*) of the system. Below, we apply the cointegration tests described here.

5.2 Testing for Cointegration - the Case of a four Variable VAR

The starting point of the cointegration analysis and the tests that follows is an unrestricted VAR model of order five with export volume (*a*), the world market indicator (*wmi*), the price ratio ($pa - pk$) and the exchange rate volatility (*Vol*) from the GARCH(1,1) model specified

[28] Doornik and Hendry (2001, p. 175) point out that the sequence of trace tests leads to a consistent test procedure, but no such result is available for the alternative maximum eigenvalue test. Hence, current practice is to only consider the former.

as endogenous I(1) variables. Additionally, we include an unrestricted constant and seasonal dummies as deterministic components in the VAR model. Estimation based on this information set produces a miss specified model as the diagnostic tests suggest significant non-normality and autocorrelation in the equations for the export volume, the price ratio and the volatility. In order to improve specification properties, we include the following quarterly unrestricted impulse dummies to mop up outliers: 1983q3, 1989q2 and 1996q1. Doing so yields a well-specified model, with exception for the volatility equation. Further investigating of the residuals from the volatility equation reveals a pattern of large outliers towards the end of the estimation period. An option would then be just to mop up this noise with additional dummy variables until the specification tests are completely satisfied by statistical criteria. However, from an economic point of view, we find such a procedure somewhat inflexible, as it is exactly the volatile nature of the exchange rate uncertainty that we want to capture by including the volatility variable in the model. On this basis we continue with the three dummy variables indicated above.

Estimating a VAR(5) model requires a large amount of parameters to be estimated relative to the set of observations in the available sample period. Thus, considering the possibility of reducing the lag length to four (or less) will give us further degrees of freedom in the estimation. The F-statistic for the null hypothesis that the coefficients of the fifth lag of the variables in the VAR are zero is $F(16,150) = 1.2846$ (0.2139). The hypothesis can then not be rejected, which supports the reduction of the VAR(5) to a VAR(4). Further reductions did not succeed.[29] The diagnostics of the VAR(4) model are reported in Table 5.1.

Table 5.1: Diagnostics of the VAR(4) model

Equation	a	$pa - pk$	wmi	Vol
AR 1-5 $F(5,52)$	0.4543 [0.808]	1.3279 [0.2670]	0.88835 [0.4957]	0.96202 [0.4496]
Norm Chi^2(2)	4.6917 [0.096]	1.5566 [0.4592]	0.09304 [0.9545]	18.95 [0.0001]**
ARCH 1-4 $F(4,49)$	2.4999 [0.055]	0.40839 [0.8017]	1.0985 [0.3679]	0.0000 [1.0000]
Het $F(28,28)$	0.3441 [0.997]	0.31700 [0.9983]	0.52764 [0.9518]	0.58408 [0.9195]

Notes: AR 1-5 is the Harvey (1981) test for fifth-order residual autocorrelation, Norm is the normality test described in Doornik and Hansen (1994), ARCH 1-4 is the Engle (1982) test for 4th order autoregressive conditional heteroskedasticity in the residuals and Het is the White (1980) test for residual heteroskedasticity. *P*-values in brackets. ** indicates rejection of the hypothesis that the model residuals are normally distributed at the 1 per cent level.

[29] The Akaike information criterion (Hendry and Doornik (2001)) was here also considered. This criterion suggested a lag order of 3. Estimating a VAR(3) did, however, produce a miss specified system (problems with autocorrelation), and we have in the following chosen to use the VAR(4) model.

As in the case of the VAR(5) model, we see that the hypothesis of normally distributed residuals in the equation for exchange rate volatility is clearly rejected. We also notice that the test statistics associated with the test for normality and autoregressive conditional heteroskedasticity are close to the critical values in the case of the export volume equation. For our VAR model to be considered as a valid starting point of the cointegration analysis, it should also be reasonably constant within the sample period. Figure 5.1 plots recursively estimated one-step ahead residuals. Not surprisingly, we observe some instabilities in the volatility equation. Especially we remark how the uncertainty increases (by considering the 2 standard deviation confidence bands) from the 2nd quarter of 1997 and onwards. For the other variables the estimated residuals are more stable, albeit we notice the borderline case in the export equation in the 2nd quarter of 1997. Nevertheless, we conclude that the VAR(4) model is a reasonably valid model underlying the proceeding cointegration analysis.

Figure 5.1: Recursive residuals of the VAR(4) model

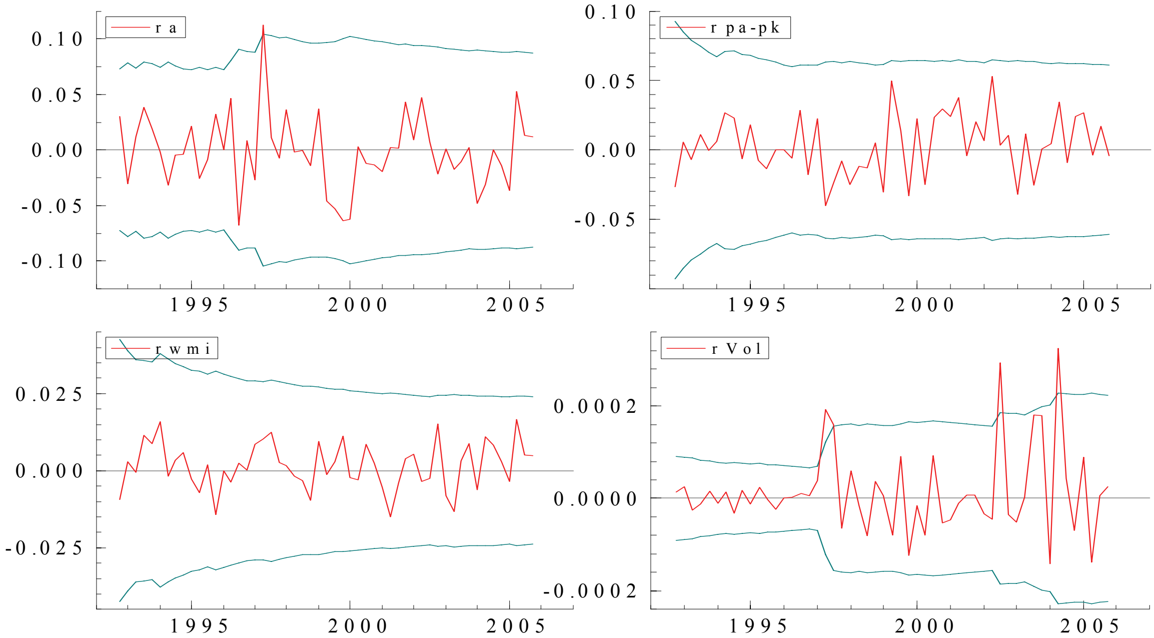


Table 5.2 reports results from applying the cointegration testing procedure suggested by Johansen (1988, 1995). We see that the rejection of the null hypothesis terminates at rank equal or less than two on a 5 per cent significance level. Hence, we may impose a rank of two, meaning that there exist two significant cointegration relationships between the selected variables.

Table 5.2: Cointegration analysis assuming a VAR(4) model

Rank under null hypothesis	Trace test statistic	P-value
$r = 0$	58.662	[0.003]**
$r \leq 1$	34.453	[0.013]*
$r \leq 2$	13.410	[0.100]
$r \leq 3$	0.10774	[0.743]

Notes: The *p*-values, which are reported in PcGive, are based on the approximations to the asymptotic distributions derived by Doornik (1998). It should be noted that the inclusion of impulse dummies in the VAR affects the asymptotic distribution of the reduced rank test statistics and therefore the critical values are only indicative. However, the bias induced by such deterministic components is supposed to be minor as only three impulse dummies are included in our case. The asterisk * and ** denote rejection of the null hypothesis at the 5 per cent and 1 per cent significance levels, respectively.

However, as previously noted, imposing a rank of two it is necessary to consider different identifying restrictions in order to establish a unique system of cointegrating vectors. Without further restrictions only the product of α and β' will be estimated consistently, c.f. equation (5.5). One example of such identifying restrictions is to set the first part of the beta matrix equal to the identity matrix, I_r. In our system with four variables and the order of rank equal to two implies the following β matrix:

$$(5.9) \qquad \beta' = \begin{pmatrix} 1 & 0 & \beta_{31} & \beta_{41} \\ 0 & 1 & \beta_{32} & \beta_{42} \end{pmatrix}.$$

Having some economic reasoning behind the ordering is here essential, otherwise randomness will distinguish the results obtained. In our work we have tested several hypothesis of identifying restrictions (including the one with the identity matrix using different orderings) extracted from economic theory and intuition. Assuming the first of the two cointegration vectors to be an export equation dictated by the model and its underlying identifying restrictions described in chapter 2 seems reasonable. Identifying the second cointegration vector may be dictated in the same fashion and from the fact that the Norwegian economy is a small open economy. Hence, we may claim that the market power of Norwegian exporters in foreign markets is to small to have any significant impact on the pricing behaviour. Under such circumstances, competition would adjust the domestic export prices in the direction of the competitive world market price. Imposing such a hypothesis would in our beta matrix correspond to setting β_{22} equal to 1 and the rest of the beta coefficients in the second column equal to zero:

$$(5.10) \quad \beta' = \begin{pmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} \\ 0 & 1 & 0 & 0 \end{pmatrix}.$$

However, such a hypothesis is clearly rejected by a likelihood ratio test and may after all indicate that Norwegian exporters have some market power in the foreign markets. Evidence in Boug *et al.* (2006) support the conclusion that export prices of machinery and equipment do not enter into a long run relationship with world market prices as specified in the tested hypothesis above.[30] Since no identifying and sensible second cointegrating vector with the exchange rate volatility treated as an endogenous variable can be established, we next consider the possibility of treating the volatility as an exogenous I(1) variable. We may justify this property of the volatility on the ground of the Norwegian economy being a small open economy. For example, the exchange rate fluctuations following the 1997 economic Asia crisis can be considered as quite exogenous from a Norwegian viewpoint. In the above-discussed framework of four endogenous variables in the VAR model such an assumption can be tested imposing restrictions of weak exogeneity for the volatility variable of the system. That is, the loading coefficients α_{41} and α_{42} in equation (5.3) are set to zero. If this is the case, neither of the two cointegration relationships affects the exchange rate volatility. Table 5.3 reports test results of the weak exogeneity status of the volatility variable with respect to the two cointegration vectors.

Table 5.3: Tests for weak exogeneity of the exchange rate volatility

Null hypothesis	Test statistic	LR statistic	P-value
$\alpha_{41} = 0, \alpha_{42} = 0$	χ^2 (1)	7.4186	[0.0245]*

Notes: The weak exogeneity test, which are asymptotically distributed as χ^2(.) under the null with the number of restrictions imposed as the degrees of freedom in parentheses [see Johansen (1995)], is calculated under the assumption that r =2. The LR statistic is the likelihood ratio statistic which compares the restricted and the unrestricted model. A lower likelihood ratio (or correspondingly higher p-value) means moving in the direction of not rejecting the imposed hypothesis.

[30] Boug *et al.* (2006) find evidence of a significant long run relationship between export prices, import prices and unit labour costs in the machinery and equipment industry. Including unit labour costs as an additional explanatory variable in our context is, however, beyond the scope of this thesis and left as a topic for future research.

Setting the loading coefficients α_{41} and α_{42} jointly to zero produces somewhat ambiguous results with respect to the weak exogeneity status of the volatility variable. With a 5 per cent significance level the hypothesis is rejected, but at a significance level of 1 per cent it is not. Nevertheless, we next investigate the cointegration properties among the four variables assuming volatility to be an exogenous variable in the VAR model.

5.3 Testing for Cointegration - Volatility as an Exogenous Nonstationary Variable

Testing for cointegration in the case of the exchange rate volatility being an exogenous I(1) variable is not straight forward. Harbo *et al.* (1998)[31] analyse the likelihood ratio test for cointegrating rank for partial systems in the VEqCM framework containing exogenous I(1) variables. Under the assumption of weak exogeneity for the cointegrating parameters, tables of critical values are provided in the mentioned study. However, the implemented testing sequence in PcGive to determine the cointegration rank is based on critical values which only cover cases of endogenous variables in the system (with a constant and/or a trend term included, either restricted or unrestricted), see Doornik (1998). Hence, the statistics and p-values reported when an exogenous variable is included in the VAR model, can only be used as guidelines in determining the cointegration rank. To simplify matters, we use the reported critical values from PcGive.

To incorporate the feature of an exogenous I(1) variable in the cointegrating relationship we need to rewrite the general VEqCM in (5.2). Following Harbo *et al.* (1998), letting $p = 4$ (as in our export model) and assuming weak exogeneity of a vector of *z* variables, the adjusted VEqCM in our case may be represented as

(5.11) $$\Delta y_t = \alpha_1 \beta' x_{t-1} + \Gamma_{11}\Delta x_{t-1} + \Gamma_{12}\Delta x_{t-2} + \Gamma_{13}\Delta x_{t-3} + u_t .$$

Here y_t is redefined as the $(m \times 1)$ vector of *m* endogenous variables in the system, z_t the $(n \times 1)$ vector of *n* exogenous variables and $x_t = (y_t, z_t)$ the $((m+n) \times 1)$ vector of all the variables in the system. The exchange rate volatility would in our example constitute the

[31] See also Pesaran *et al.* (2000).

single variable in the z_t vector. The α_1 is the remaining upper part of the alpha matrix in (5.3) not set equal to zero due to the assumed endogeneity status of the export volume, the price ratio and the world market demand. Finally, Γ_{1i} (i= *1,2,3)* are the coefficient matrices of the lagged differenced variables.

Specifying this model in PcGive with the exchange rate volatility as an I(1) exogenous variable is done in the following manner: As before, we formulate a fourth order VAR model with the three endogenous variables together with a constant term, seasonal dummies and two impulse dummies (1989q2 and 1996q1) entering unrestrictedly.[32] In addition, we include the volatility in level as a *restricted* variable in the VEqCM, whereas its first difference with zero lag up to three lags are included as *unrestricted* variables. Table 5.4 reports diagnostics of this model. Compared to the case with volatility treated as an endogenous variable this system performs better.

Table 5.4: Diagnostics of the VAR(4) model

Equation	a	$pa-pk$	wmi
AR 1-5 test $F(5,52)$	0.57381 [0.7197]	0.74293 [0.5949]	2.2709 [0.0609]
Norm: Chi^2(2)	4.8318 [0.0893]	0.88232 [0.6433]	2.3492 [0.3089]
ARCH 1-4 $F(4,49)$	1.7099 [0.1629]	0.45897 [0.7654]	0.78649 [0.5395]
Het $F(25,31)$	0.48816 [0.9652]	0.36618 [0.9940]	0.86250 [0.6446]

Notes: See Table 5.1. Volatility is treated as an exogenous I(1) variable.

Also, the recursively estimated one step ahead residuals displayed in Figure 5.2 suggest that the system is reasonably stable. Albeit a borderline case, one possible instability may be the case for the export volume equation in 1997, which we remember was also detected in Figure 5.1 above.

[32] Here the dummy for 1986q3 is left out, as this initially was included to mop up a large outlier in the volatility equation due to exchange rate fluctuations in the wake of the Norwegian devaluation in 1986.

Figure 5.2: Recursive residuals of the VAR(4). Volatility as an exogenous I(1) variable

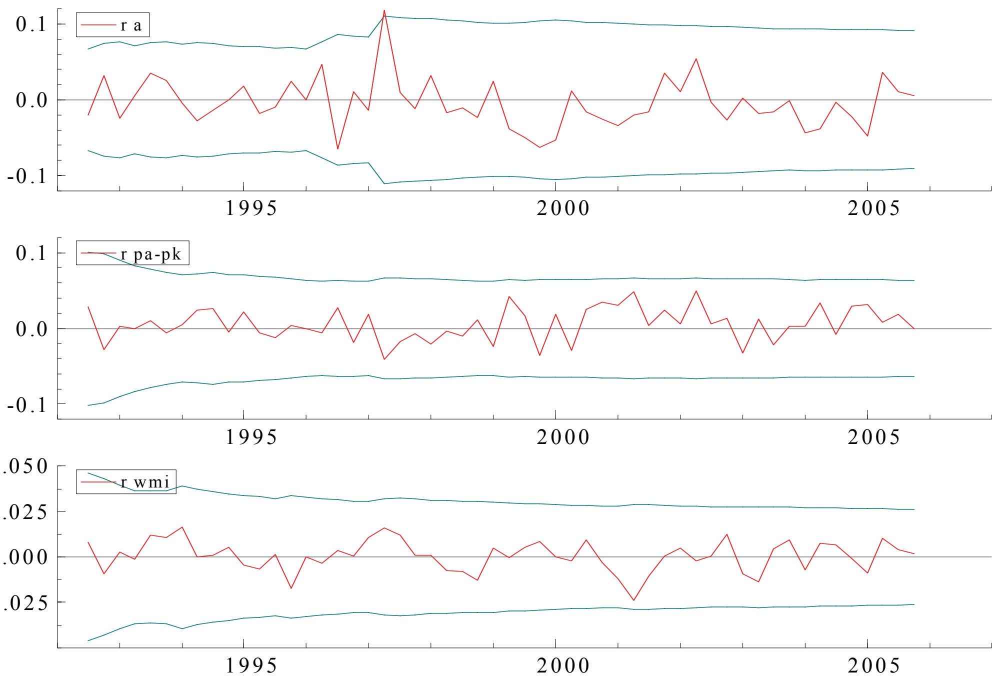


Table 5.5 reports results from a cointegration analysis assuming a VAR(4) model with volatility as an exogenous I(1) variable. We find evidence of at least one cointegrating relation among the variables in this case. Bearing in mind that this test now is to be considered with some caution, we impose the rank of one to investigate the case with one cointegrating relationship. In equation (5.12) the unrestricted estimate of the cointegration relationship is reported with standard errors in parentheses.

Table 5.5: Cointegration analysis assuming a VAR(4) model

Rank under null hypothesis	Trace test statistic	P-value
$r = 0$	30.730	[0.039] *
$r \leq 1$	10.751	[0.231]
$r \leq 2$	0.48902	[0.484]

Notes: See Table 5.2. Volatility is treated as an exogenous I(1) variable.

$$(5.12) \quad \hat{\alpha}\hat{\beta}' \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \\ Vol \end{pmatrix}_{t-1} = \begin{pmatrix} -0.2759 \\ [0.0837] \\ 0.0794 \\ [0.0574] \\ 0.0451 \\ [0.0246] \end{pmatrix} \cdot \begin{pmatrix} 1 & 0.4760 & -1.2814 & 25.590 \\ [0] & [0.3877] & [0.1335] & [138.56] \end{pmatrix} \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \\ Vol \end{pmatrix}_{t-1}$$

Considering the estimated cointegration relationship we notice the very large standard deviation of the volatility coefficient in the beta vector. Indeed, imposing the hypothesis that the volatility coefficient in the beta vector is zero is clearly not rejected. The likelihood ratio statistic with one restriction imposed yields Chi^2(1) = 0.037958 and the p-value of 0.8455, which is far from rejection of the hypothesis.[33] This is quite strong evidence supporting an exclusion of the exchange rate volatility as an exogenous I(1) variable from the cointegration vector. Consequently, we next consider the case in which the exchange rate volatility is an exogenous I(0) variable to the system.

5.4 Testing for Cointegration - Volatility as a Stationary Variable

Rahbek and Mosconi (1999) address cointegration rank inference with stationary regressors in VAR models. Inclusion of the exchange rate volatility as a stationary variable in the VAR will as in the case of volatility included as an exogenous I(1) variable disturb the usual critical values when determining the cointegration rank. However, as we shall see, the volatility variable as a stationary variable in the VAR turns out insignificant. We may then interpret the cointegration analysis in this case as if it was a cointegration analysis of the three endogenous variables, namely the export volume, the relative prices and the world market demand indicator. Hence, the critical values reported in PcGive in connection with the cointegration analysis may then be considered as asymptotically valid, and not just as guidelines.

Following Rahbek and Mosconi (1999), treating volatility as I(0), we again respecify the general VEqCM in (5.2) to become

$$(5.13) \quad \Delta y_t = \alpha_1 \beta' y_{t-1} + \sum_{i=1}^{3} \Gamma_i \Delta y_{t-i} + \sum_{i=0}^{4} z_{t-i} + u_t \,.$$

[33] This result is valid independent of which volatility lag to be included in the long run solution. The same test procedure does in all cases of lags up to four clearly not reject the hypothesis that the coefficient is zero.

Keeping the notation from the preceding section, y_t is here the vector of endogenous variables and z_t the (stationary) exogenous variable(s). In our export model the volatility variable would here again enter into the z_t vector as the single exogenous variable. Specifying equation (5.13) as a VAR model in PcGive is done by setting up a system of the three endogenous variables together with volatility in levels up to the 4^{th} lag, a constant term, seasonal dummies and impulse dummies as unrestricted terms in the model. In the final VEqCM the volatility terms appear as short-term regressors in levels, due to the assumed stationary nature of the variable. Estimating a VAR(4) model in this case, we get a well-specified system of equations. But the unrestricted volatility terms are not entering significantly into the model as clearly indicated by $F(15,152) = 0.50537$ and its associated p-value of 0.9349.

5.5 Testing for Cointegration - the Case of a Three Variable VAR

With the test results presented above, we may remove the volatility variable from the VAR model completely. We are then left with a system of three endogenous variables, which we now confront with cointegration analysis. As before, we initially estimated a fifth order VAR, which was reduced to a fourth order VAR on the basis of the same testing procedure as discussed previously. This leaves us with a well specified VAR(4) system with the diagnostics reported in Table 5.6.

Table 5.6: Diagnostics of the VAR(4) model

Equation	a	$pa - pk$	wmi
AR 1-5 test $F(5,57)$	0.47129 [0.7961]	0.72780 [0.6055]	1.0510 [0.3970]
Norm Chi^2(2)	3.5811 [0.1669]	0.89556 [0.6390]	1.5894 [0.4517]
ARCH 1-4 $F(4,54)$	1.6433 [0.1768]	0.51458 [0.7253]	0.71173 [0.5875]
Het $F(24,37)$	0.52961 [0.9479]	0.39148 [0.9910]	1.1859 [0.3138]

Notes: See Table 5.1.

Specifically, it is worth noticing the p-values for normality and conditional heteroskedasticity in the export equation. Compared to the diagnostics of the system including the volatility measure (Table 5.1), the p-values are now considerably improved.

Figure 5.3 plots recursively estimated one step ahead residuals. We conclude that the system is fairly stable. Noticeably, the possible instability in 1997, which was revealed in the Figures 5.1 and 5.2, is more or less absent in Figure 5.3.

Figure 5.3: Recursive residuals of the VAR(4) model

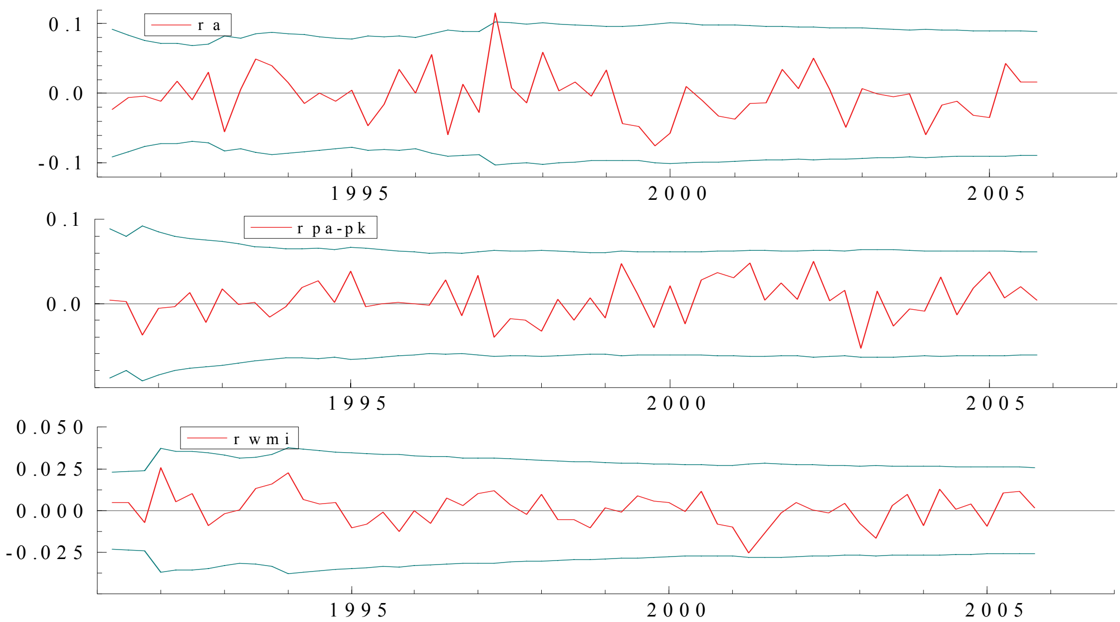


Table 5.7 reports results from the cointegration analysis in the case of a three variable VAR model without the exchange rate volatility.

Table 5.7: Cointegration analysis assuming a VAR(4) model

Rank under null hypothesis	Trace test statistic	Prob-value
$r = 0$	29.92	[0.048]*
$r \leq 1$	9.19	[0.354]
$r \leq 2$	0.35	[0.556]

Notes: See Table 5.2.

The test fails to reject the hypothesis of a rank of unity or less, leading to the conclusion that there exists one significant cointegration relationship among the variables in the system. Our original system of four endogenous variables, containing two cointegrating vectors, is now reduced to a three variable system with one cointegrating vector as specified in equation (5.14), α and β are here (3×1) vectors:

$$(5.14)\quad \alpha\beta' \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \end{pmatrix}_{t-1} = \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{pmatrix} \cdot \begin{pmatrix} \beta_{11} & \beta_{21} & \beta_{31} \end{pmatrix} \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \end{pmatrix}_{t-1}$$

Estimating a reduced rank VAR model yields the following unrestricted estimates of the alfa and beta matrices, where β_{11} is normalised to one and standard errors are reported in square brackets:

$$(5.15)\quad \hat{\alpha}\hat{\beta}' \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \end{pmatrix}_{t-1} = \begin{pmatrix} -0.29300 \\ [0.081783] \\ 0.081647 \\ [0.055792] \\ 0.043691 \\ [0.024204] \end{pmatrix} \cdot \begin{pmatrix} 1 & 0.46542 & -1.2770 \\ [0] & [0.34321] & [0.12447] \end{pmatrix} \cdot \begin{pmatrix} a \\ pa-pk \\ wmi \end{pmatrix}_{t-1}$$

First, we notice that the estimates in (5.15) are almost identical to those in (5.12), hence adding force to the validity of the assumption that the exchange rate volatility could be excluded from the VAR model. At first glance, these estimates seem quite plausible and in line with the underlying theoretical model in chapter 2. The price ratio has the expected negative effect on the export volume, whereas the world demand indicator clearly has a positive impact on exports as predicted by theory.[34] Interestingly, the magnitude of the long run coefficients are practically speaking about identical to those in the existing export model of machinery and equipment in MODAG, see Boug *et al.* (2002). We also notice that the loading coefficient associated with the export equation is highly significant, meaning that the cointegrating vector enters the export equation in the VEqCM. The significance of the other loading coefficients is more uncertain, which is related to the issue whether some of the other variables are weakly exogenous to the model. To conclude the cointegration analysis,

[34] We should remark that the long run coefficient of the price ratio is somewhat dependent on the inclusion of the dummy variable for the 2nd quarter of 1989 in the VAR model. It is initially included to reduce problems of autocorrelation and non-normality in the residuals. Leaving it out, however, makes the long run effect of relative prices on export volume somewhat uncertain. As can be seen from the standard errors in (5.15), the significance of the long run price effect may be questionable. Considering the estimate of the price coefficient and its belonging standard error, we may compute the t-value to be *t=0.4652/0.3432=1.356*. Using critical values from a t-distribution this coefficient clearly would be deemed not significantly different from zero in a two-sided test. As it may be more relevant to consider a one-sided test in line with the economic model in chapter 2, calculating the p-value with 80 observations and a critical value of 1.356 yields p=0.0895. Hence, at a 10 per cent level this coefficient is statistically significant and not just significant from an economic point of view.

we address this issue further below together with hypothesis testing about the beta parameters.

5.6 Testing cointegration restrictions

With the single cointegration vector identified, investigating possible restrictions on α and β is possible by means of the testing procedure reviewed in Johansen (1995). Table 5.8 reports several hypothesis about the alfa and beta matrices in (5.15) assuming one significant cointegrating vector.

Using likelihood ratio tests, we first test if any of the beta coefficients are not significantly different from zero. As we have already seen, the significance of the relative price coefficient is somewhat of a borderline case. The test in Table 5.8 confirms this by not rejecting the null hypothesis that β_{21} actually is equal to zero. However, a one sided test indicates that it is indeed significantly different from zero. We chose to use the estimate of the price ratio parameter based on this result and economic reasoning. The hypotheses that the two other variable coefficients are zero are clearly rejected by the data.

Next we test for weak exogeneity. The results here are in line with prior expectations. The hypothesis that the export volume is weakly exogenous is clearly rejected at the 1 per cent level. For the two remaining variables, weak exogeneity cannot be rejected. Also, the hypothesis of homogeneity of degree one between exports and the world demand indicator is not rejected. Lastly, we test the joint hypothesis of homogeneity between exports and world demand and weak exogeneity status of world demand and relative prices. We see that the joint hypothesis is not rejected with a p-value above 20 per cent.

Table 5.8: Testing restriction hypotheses[35]

Null hypothesis	Test statistic	*LR* statistic	p-value
$\beta_{11} = 1, \beta_{21} = 0$	Chi^2(1)	1.0512	[0.3052]
$\beta_{11} = 1, \beta_{31} = 0$	Chi^2(1)	10.763	[0.0010]**
$\beta_{11} = 0, \beta_{21} = 1$	Chi^2(1)	11.741	[0.0006]**
$\alpha_{11} = 0, \beta_{11} = 1$	Chi^2(1)	8.8075	[0.0030]**
$\alpha_{21} = 0, \beta_{11} = 1$	Chi^2(1)	2.4858	[0.1149]
$\alpha_{31} = 0, \beta_{11} = 1$	Chi^2(1)	2.6093	[0.1062]
$\beta_{11} = 1, \beta_{31} = -1$	Chi^2(1)	3.0185	[0.0823]
$\beta_{11} = -\beta_{31} = 1, \alpha_{21} = \alpha_{31} = 0$	Chi^2(3)	4.6290	[0.2011]

Notes: The tests, which are asymptotically distributed as $\chi^2(.)$ under the null [see Johansen (1995)], are calculated under the assumption that r =1.The test statistics are Chi^2 distributed with the number of restrictions imposed as the degrees of freedom in parentheses. The *LR* statistic is just the likelihood ratio statistic comparing the restricted and the unrestricted model. A higher likelihood ratio (or correspondingly lower p-value) means moving in direction of rejection of the imposed hypothesis of restrictions. The asterisk ** denotes rejection of the null hypothesis at the 1 per cent significance level.

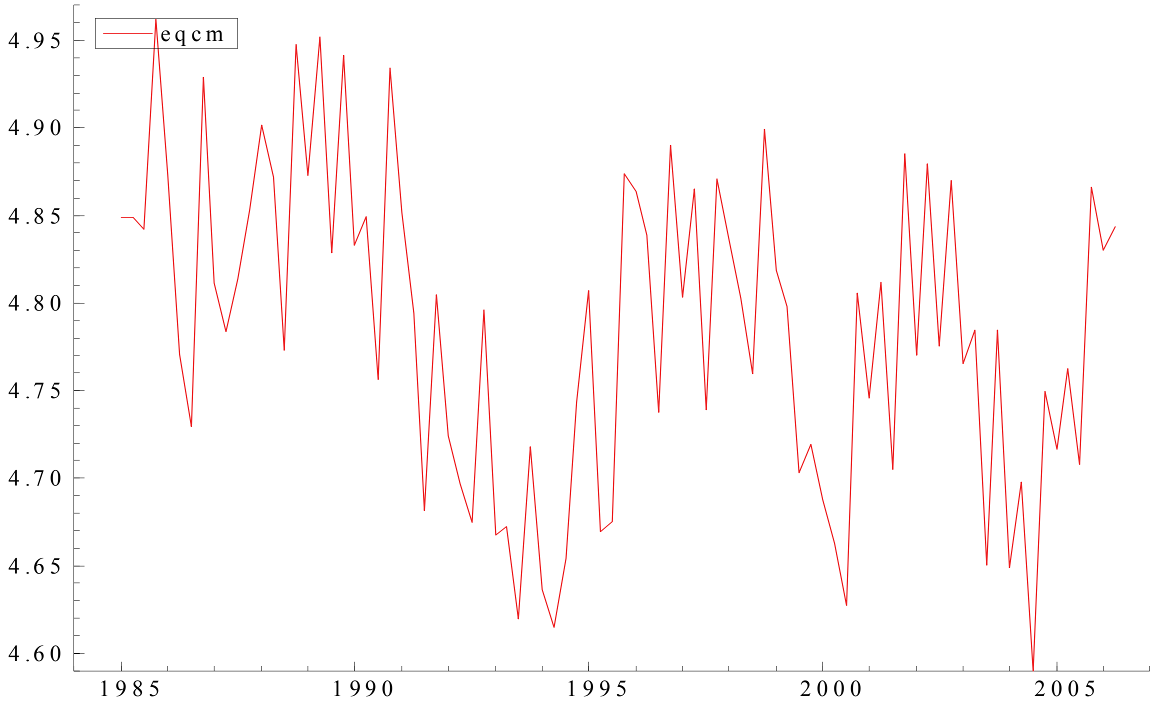
Imposing the joint hypothesis, results in the following restricted cointegration relationship with standard errors in parenthesis:

$$(5.16) \qquad eqcm_t = a_t + \underset{[0.1499]}{1.2457}\ (pa - pk)_t - wmi_t$$

We notice that the coefficient of the price ratio enters significantly into the restricted cointegrating vector. Figure 5.4 displays the $eqcm_t$ variable. We observe that the variable is fairly stationary during the sample period. Since relative prices and the world market demand may be considered as weakly exogenous to the cointegrating vector, our original VEqCM in (5.2) reduces to a single equilibrium correction model (EqCM) for the growth of exports, which we now turn to.

[35] In the same fashion we have also performed multivariate testing of order of integration of the variables. Setting one of the coefficients in β equal to 1 and the remaining to zero means that the variable left in the restricted cointegration vector would be a stationary process under the null hypothesis. Imposing such hypotheses for the three variables of the system in turn are clearly rejected by the LR tests.

Figure 5.4: The equilibrium correction mechanism



6. The Equilibrium Correction Model

Using the same information set as in the cointegration analysis and relying on a general-to specific modelling strategy, see Hendry and Doornik (2001), we establish a parsimonious EqCM model for exports of machinery and equipment. First, we derive at a specific model which includes short run exchange rate volatility effects. Since none of these effects enter significantly into the model, we next estimate a specific model in which volatility plays no role at all. Finally, we conduct careful analyses of the parsimonious export model with respect to its in-sample as well as its out-of-sample economic properties. Herein, we pay particular attention to the possibility of structural breaks in the estimated export model in 1997 and 2001 due to the Asian crisis and the monetary policy regime shift in Norway, respectively. We proceed from the following general model, which is consistent with the reduced rank VAR model established above:

$$
\begin{aligned}
\Delta a_t &= const. + \sum_{i=1}^{3} \Delta a_{t-i} + \sum_{i=0}^{3} \Delta(pa - pk)_{t-i} + \sum_{i=0}^{3} \Delta wmi_{t-i} + \sum_{i=0}^{4} Vol_{t-i} \\
&\quad + [a + 1.2457(pa - pk) - wmi]_{t-1} + dummies + u_t,
\end{aligned}
\qquad (6.1)
$$

where *dummies* includes the same set of dummy variables as in the VAR model. We notice that the general export model allows for the possibility of short run effects from the exchange rate volatility. A general-to-specific approach applied to (6.1), removing insignificant variables sequentially, produces the specific model presented in Table 6.1.

We notice that the coefficients of the volatility variable are clearly not significant using a two-sided test. It is also interesting to see the estimated symmetric effects these coefficients apparently bring about. If volatility does have some effects from the first lag, it is neutralised by the effects from the second lag. Hence, if volatility plays a role at all, its effects are clearly small and not significant from a statistical point of view.

Table 6.1: OLS estimates of an equilibrium correction model for exports

Δa_t	=	1.10760 (0.3061)	-0.366534 Δa_{t-1} (0.08423)	-0.735 $\Delta(pa-pk)_t$ (0.1385)	+1.01942 Δwmi_{t-2} (0.3640)
		-0.211 $eqcm_{t-1}$ (0.06434)	-46.078 Vol_{t-1} (32.70)	+46.789 Vol_{t-2} (31.60)	+0.101 D_{1t} (0.04479)
		+0.0916583 D_{2t} (0.04338)	-0.0849 S_{0t} (0.02352)	-0.115704 S_{1t} (0.01532)	-0.168 S_{2t} (0.0153)
		sigma=0.0418	R^2 =0.85289	T=80	

Notes: Standard errors in parenthesis. Sigma is the equations standard error and R^2 is the squared multiple correlation coefficient, see Doornik (2001). T is the number of observations used in the estimation. D1 is the impulse dummy for the observation 1989q2 and D2 is the impulse dummy for the observation 1996q1. S_i ($i=0,1,2$) are seasonal dummy variables.

Accordingly, we move on to estimate a final equilibrium correction model without the volatility variable in the information set. Following the same approach as above, we derive at the following parsimonious EqCM model for exports of machinery and equipment:

Table 6.2: OLS estimates of an equilibrium correction model for exports

Δa_t	=	1.114 (0.3042)	-0.362 Δa_{t-1} (0.085)	-0.702 $\Delta(pa-pk)_t$ (0.1448)	+0.921 Δwmi_{t-2} (0.3531)
		-0.213 $eqcm_{t-1}$ (0.064)	+0.10 D_{1t} (0.045)	+0.097 D_{2t} (0.043)	-0.083 S_{0t} (0.024)
		-0.108 S_{1t} (0.015)	-0.167 S_{2t} (0.015)		
		sigma=0.042	R^2 =0.847	T=80	

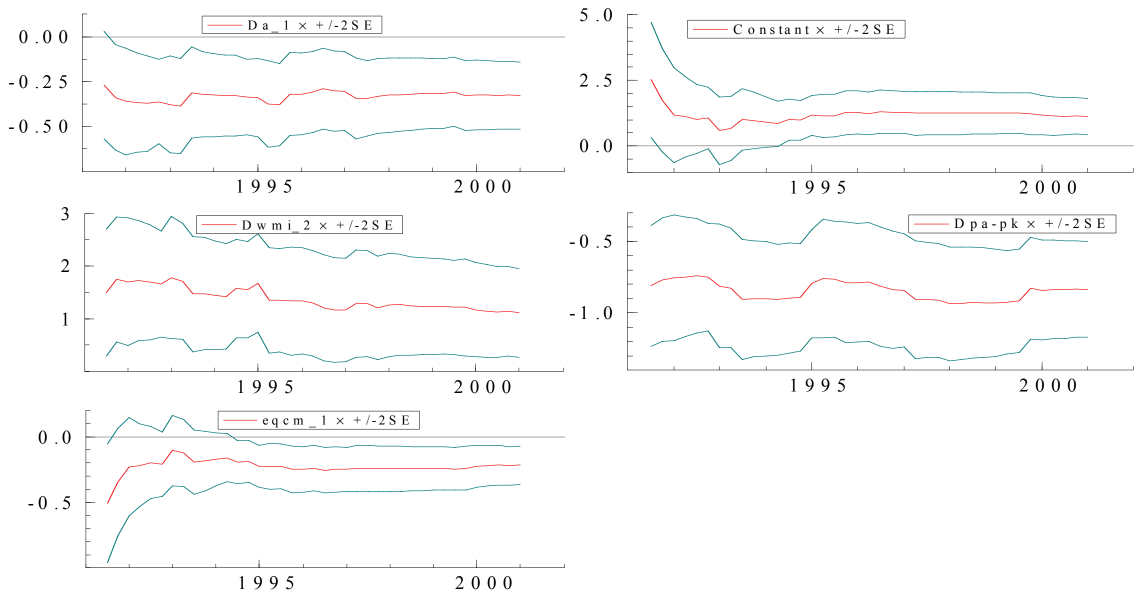
Model diagnostics:

Test	Value	p-value
AR 1-5:	$F(5,65)$ = 1.1937	[0.3221]
ARCH 1-4:	$F(4,62)$ = 0.89806	[0.4707]
Norm. test:	Chi^2(2) = 4.3329	[0.1146]
Het:	$F(13,56)$ = 0.53947	[0.8895]
RESET test:	$F(1,69)$ =0.0022657	[0.9622]

Notes: Standard errors in parenthesis. Sigma is the equations standard error and R^2 is the squared multiple correlation coefficient, see Doornik (2001). T is the number of observations used in the estimation. D1 is the impulse dummy for the observation 1989q2 and D2 is the impulse dummy for the observation 1996q1. S_i ($i=0,1,2$) are seasonal dummy variables. AR 1-5 is the Harvey (1981) test for fifth-order residual autocorrelation, Norm is the normality test described in Doornik and Hansen (1994), ARCH 1-4 is the Engle (1982) test for 4th order autoregressive conditional heteroskedasticity in the residuals, Het is the White (1980) test for residual heteroskedasticity and RESET is the Ramsey (1969) test for functional form misspecification

Below the EqCM model we report several diagnostic tests. The model seems well specified as none of the tests are significant at conventional levels. Also, all variables are significantly estimated. Particularly, the *eqcm* variable enters the model with a t-value of −3.328, supporting the conclusions from the cointegration analysis. Moreover, the estimated short run dynamic effects are in line with prior expectations and smaller in magnitude than their long run counterparts. Hence, no overshooting in the export volume with respect to shocks in the explanatory variables is present in the model. Studying the model closer, we observe a potential problem of simultaneity bias due to the contemporaneous effects of the price ratio (*pa−pk*). Applying instrumental variables may, however, solve this issue. One such model is estimated in the Appendix B. The estimation results are similar to that of the model in Table 6.2. Especially, the important coefficients including that of the equilibrium correction term do not change substantially. We may therefore conclude that the model in Table 6.2 is consitently estimated by OLS. Finally, as seen from the recursive estimates of the coefficients reported in Figure 6.1, all parameters are reasonably stable and significant, especially towards the end of the estimation period.

Figure 6.1: Recursive estimates of the coefficients of the parsimonious equilibrium correction model



We have seen that the parsimonious EqCM model is well specified and exhibits constancy in-sample. To conclude this section, we study out-of-sample forecasting performance of the estimated export model in order to shed light on its robustness with respect to the Asian financial crisis in 1997 and the monetary policy regime change in Norway in 2001. As noted in the introduction, these events may have influenced the Norwegian exporters' behaviour in a significant way. For instance, the introduction of inflation targeting and freely floating exchange rates in late March 2001, has without doubt brought about increased exchange rate volatility and risk faced by Norwegian exporters. So, if exchange rate volatility do play a role in the decision process of the exporters, we should expect instabilities in the estimated model as, for example, indicated by poor out-of-sample forecasting ability. Here we use simple one step ahead forecasts, which in general may be formulated as follows, see e.g. Doornik and Hendry (2001):

(6.2) $y_t = \tau_1 y_{t-1} + \tau_2 y_{t-2} + \tau_3 z_t + \varepsilon_t$.

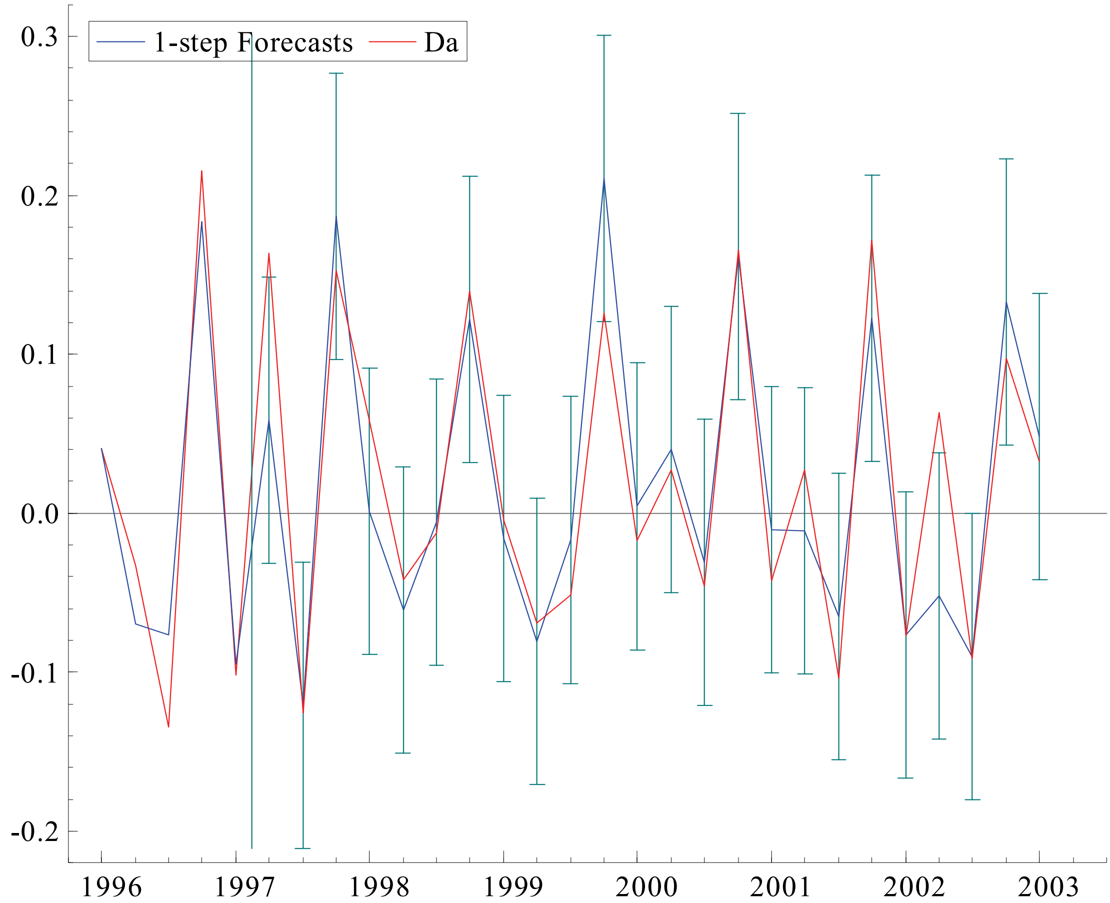
A static one period ahead forecast in period T (T+1) is then given as

(6.3) $\hat{y}_{T+1} = \hat{\tau}_1 \, y_T + \hat{\tau}_2 \, y_{T-1} + \hat{\tau}_3 \, z_{T+1}$.

To assess the forecasting performance of the export model in the event of the financial crises in Asia, we reestimate the model based on observations until 1997 second quarter and employ thirty five quarters (1997q2 – 2005q4) of out-of-sample observations. Figure 6.2 depicts actual values of Δa_t together with one-step ahead forecasts, adding bands of 95 per cent confidence intervals to each forecast in the forecasting period. We see that the actual value of the growth in exports stays outside its corresponding confidence interval in the first quarter of the forecasting period (albeit a borderline case), suggesting that the financial crisis in Asia indeed had some impact on the exporting behaviour. However, in the consecutive quarter the actual value of Δa_t is clearly on track with its forecasting value. Thereafter, the estimated model continues to forecast the observed values with more or less accuracy[36] and may indicate and understate that transitory exchange rate volatility at least in the long run has no substantial impact on export volume.

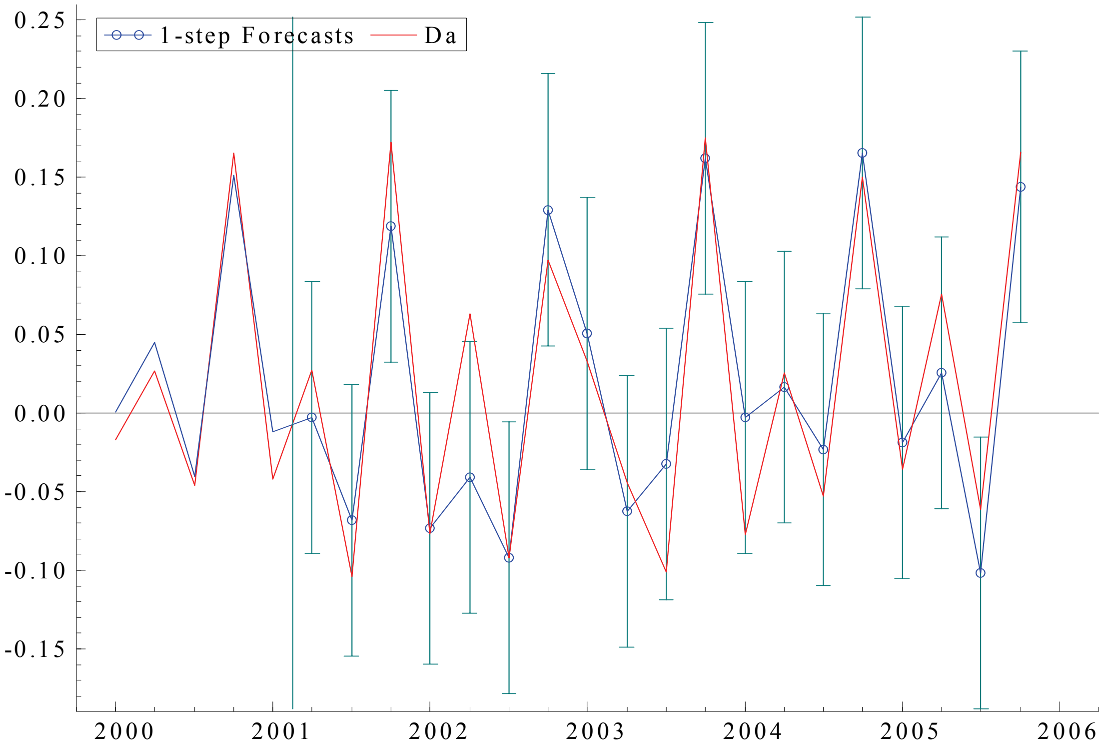
[36] The same pattern can also be seen in the beginning of 2002.

Figure 6.2: One step ahead forecasts for the growth in exports



To assess the forecasting performance of the export model in the event of the monetary policy regime shift in Norway, we reestimate the model based on observations until 2001 second quarter and use nineteen quarters (2001q2 – 2005q4) of out-of-sample observations. Figure 6.3 depicts actual values of Δa_t together with one-step ahead forecasts, adding bands of 95 per cent confidence intervals to each forecast in the forecasting period. As seen, the model forecasts only misses significantly the observed values once, namely in the second quarter of 2002. Since the period of the forecasting failure does not coincide with the second quarter of 2001, it has nothing to do with the monetary policy regime shift. Thus, we may conclude that the out-of-sample forecasting ability of the parsimonious export model is satisfactory despite a major regime change in monetary policy.

Figure 6.3: One step ahead forecasts for the growth in exports



7. Conclusions

This thesis has sought to investigate possible effects of exchange rate volatility on Norwegian exports by means of a demand type export model with relative prices, world market demand and exchange rate volatility as explanatory variables using disaggregated quarterly data and multivariate cointegration techniques. In the wake of the breakdown of the Bretton-Woods agreement in the early 1970s, it has been claimed that the transition to floating exchange rate regimes in many of the worlds major trading countries has induced negative effects on international trade flows. The argument according to standard microeconomic theory is that increased exchange rate volatility lead risk averse exporters to reduce their trade volumes. This hypothesis is, however, not unanimously supported by empirical studies. Researchers have reported findings in either direction, if any effects are proved significant at all.

As corresponding evidence from analyses performed on Norwegian data are rather rare, this thesis has aimed at contributing to the literature by investigating possible links between exchange rate volatility and exports using data from one major export industry in Norway, namely machinery and equipment. The fact that the Norwegian economy since the early 1990s has been a trade dependent small open economy with managed and freely floating exchange rates makes this study highly relevant. Also, the conducted analyses of special events in foreign financial markets and monetary policy regime shifts in Norway and their potential impacts on Norwegian exporting behaviour are major contributions to the literature.

In the empirical analysis, we have paid special attention to the construction of and evaluation of the exchange rate volatility measure applied. The time series of exchange rate volatility used in related studies are typically regarded as nonstationary variables. Thus, applying Johansen's (1988, 1995) multivariate method for determining cointegrating relations between selected variables is a popular tool. However, as commented upon by McKenzie (1999) among others, the nature of the time series properties of the exchange rate volatility is often hard to establish and often ignored or neglected in previous studies. As the measure of exchange rate volatility in this thesis displays dubious nature with respect to time series properties, we have conducted cointegration analyses under different assumptions about the stationarity status of the exchange rate volatility. Our starting point has been the case in which the volatility measure is assumed to be an *endogenous* I(1) variable. We were, however, in this case unable to identify economically meaningful cointegrating relationships among the selected variables using the procedure suggested by Johansen (1988, 1995). Hence, we next treated the case in which the exchange rate volatility is assumed to be an *exogenous* I(1) variable, still possibly entering the cointegrating relationships in the spirit of Harbro *et al.* (1998). Again, we did not succeed in finding significant long run effects from the exchange rate volatility on the export volume. As a final case, we considered the possibility of volatility being an *exogenous* I(0) variable, thus only allowing for potential short run effects in the fashion of Rahbek and Mosconi (1999). This approach did also, however, lead to the rejection of the hypothesis that volatility has any significant effects on the export volume. Leaving out the volatility variable from the information set completely, we established a unique cointegrating relationship between exports, relative prices and world market demand consistent with the underlying theoretical model.

Finally, we estimated a well-specified and stable parsimonious EqCM model for Norwegian exports of machinery and equipment without any effects from the exchange rate volatility relying on a general-to-specific modelling strategy. We also examined the out-of-sample forecasting ability of the estimated export model in order to shed light on its robustness with respect to the events of the financial crisis in Asia in 1997 and the monetary policy regime shift in Norway in 2001. The parsimonious model seems to handle the regime shift well, whereas the point in time of the financial crisis in Asia makes the model produce a forecast that significantly differs from the corresponding actual value of the growth in exports. In the following periods, however, the model produces forecasts that stay clearly within their corresponding confidence intervals, suggesting that exchange rate volatility plays a minor role in the decision making of Norwegian exporters, at least in the long run. Based on these forecasting exercises, we have in the present thesis concluded that the parsimonious export model performs well out-of-sample despite events of a monetary policy regime shifts and a foreign financial crisis.

A potential caveat in our empirical analysis is the use of quarterly data. Most related studies make use of monthly observations. We have strived to apply proxies as close to the underlying theoretical variables in the economic model as possible, which have led us to use quarterly data. A future project would then naturally involve the consideration of finding and applying more frequently reported time series. Another aspect for future research may be to consider a broader selection of industries in the economy. The findings in this thesis are only valid for the investigated sector of machinery and equipment. As pointed out by McKenzie (1999), differences among industries in the effects from exchange rate volatility on trade are most likely to occur.

Another potential caveat is related to the identification of the two cointegration relationships detected in the empirical analysis in the case of the exchange rate volatility being an endogenous nonstationary variable. As commented upon, Boug *et al.* (2006) find evidence of a long run relationship between export prices, import prices and unit labour costs in the machinery and equipment industry. Hence, by imposing the unity restriction on the price ratio as an identifying restriction and leaving out unit labour costs, we may miss out relevant information in the identification of the cointegration relationships. Extending the analysis to also include unit labour costs in order to model both export volume and prices simultaneously may provide a solution to the identification problem. In this thesis, we have left this issue open for future investigation.

Yet, another research area may involve studying the trade relations between Norway and the different Euro countries more closely. The debate around the possible Norwegian EU membership (also leading to an adoption of the euro) has so far not extensively treated the issue of exchange rate volatility effects on export performance, and potential gains/losses from avoiding these through a common currency with the Norwegian main trading partners in the Euro zone.

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Appendix A - Variable Definitions and Sources

Time series	Description
A	Export volume index (2000*Q*1=1), machinery and equipment. Source: Quarterly National Accounts (Statistics Norway).
PA	Export price index expressed in the Norwegian currency (2000*Q*1=1), machinery and equipment. Source: Quarterly National Accounts (Statistics Norway).
PK	Competitive world market price index* (2000*Q*1=1), machinery and equipment. See chapter 3 for construction of the index.
WMI	World market demand index for Norwegian goods (2000*Q*1=1). Source: Statistics Norway.
VOL	Measure of trade weighted Norwegian exchange rate volatility. See chapter 3 for construction of the volatility measure.
E	Nominal bilateral exchange rate. Source the Norwegian Central Bank: http://www.norges-bank.no

* In the construction of *PK* data from several sources are collected.
In addition to the EcoWin (Reuters) database, series have been found at the individual country's own statistical online services:

- Statistics Sweden: http://www.scb.se
- Bank of Japan: http://www.boj.or.jp
- UK National Statistics: http://www.statistics.gov.uk
- Bureau of Labor Statistics: http://www.bls.gov

Other online resources:

- Eurostat: http://europa.eu.int/comm/eurostat
- UN: http://unstats.un.org/unsd/default.htm

Appendix B - Instrumental variable (IV) estimation of the EqCM

Δa_t	=	0.196	-0.295 Δa_{t-1}	-1.166 $\Delta(pa-pk)_t$	+0.942 Δwmi_{t-2}
		[0.330]	[0.100]	[0.336]	[0.378]
		-0.230 $eqcm_{t-1}$	+0.0814 D_{1t}	+0.095 D_{2t}	-0.107 S_{0t}
		[0.0696]	[0.050]	[0.046]	[0.030]
		-0.106 S_{1t}	-0.161 S_{2t}		
		[0.016]	[0.016]		
		sigma=0.045		T=80	

Notes: Standard errors in brackets. Sigma is the equations standard error and T is the number of observations used in the estimation. D1 is the impulse dummy for the observation 1989q2 and D2 is the impulse dummy for the observation 1996q1. S_i ($i=0,1,2$) are seasonal dummy variables.

Here instrumental variables have been in the place of the contemporaneous growth of the price ratio due to the possible simultaneity problem. Additional instruments for the modelling of the price ratio growth are the three first lags of the variable. The instrumental variables pass the Sargan (1958) test checking the requirement of an instrumental variable to be independent of the models error terms (the null hypothesis is that the instrument and the error terms are independent, and the p-value indicates that this hypothesis is not rejected).

Model diagnostics

Test	Value	p-value
AR 1-5:	$F(5,65)$ = 1.1671	[0.3349]
ARCH 1-4:	$F(4,62)$ = 0.72644	[0.5773]
Norm. test:	Chi^2(2) = 5.4273	[0.0663]
Het:	$F(13,56)$ = 0.56472	[0.8713]
RESET test:	$F(31,38)$ = 0.60576	[0.9226]
Sargan spec. test	Chi^2(2) = 3.6291	[0.1629]

Notes: AR 1-5 is the Harvey (1981) test for fifth-order residual autocorrelation, Norm is the normality test described in Doornik and Hansen (1994), ARCH 1-4 is the Engle (1982) test for 4th order autoregressive conditional heteroskedasticity in the residuals, Het is the White (1980) test for residual heteroskedasticity and RESET is the Ramsey (1969) test for functional form misspecification The Sargan spec. test is Sargan's (1958) specification test for the independence of the instrumental variable and the error terms of the model.