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Exchange rate volatility and export performance:
A cointegrated VAR approach

Abstract:
During the last decades Norwegian exporters have – despite various forms of exchange rate targeting – faced a rather volatile exchange rate which may have influenced their behaviour. Recently, the shift to inflation targeting and a freely floating exchange rate has brought about an even more volatile exchange rate. We examine the causal link between export performance and exchange rate volatility across different monetary policy regimes within the cointegrated VAR framework using the implied conditional variance from a GARCH model as a measure of volatility. Although treating the volatility measure as either a stationary or a non-stationary variable in the VAR, we are not able to find any evidence suggesting that export performance has been significantly affected by exchange rate uncertainty. We find, however, that volatility changes proxied by blip dummies related to the monetary policy change from a fixed to a managed floating exchange rate and the Asian financial crises during the 1990s enter significantly in a dynamic model for export growth – in which the level of relative prices and world market demand together with the level of exports constitute a significant cointegration relationship. A forecasting exercise on the dynamic model rejects the hypothesis that increased exchange rate volatility in the wake of inflation targeting in the monetary policy has had a significant impact on export performance.

Keywords: Exports, exchange rate volatility, GARCH, CVAR, forecasting

JEL classification: C51, C52, F14, F17

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1. Introduction

Since the breakdown of the Bretton-Woods agreement and the transition to floating exchange rates the nature and magnitude of the relationship between exchange rate volatility and trade flows has been a subject of major concern to economists. A number of theoretical models exist showing that the impact of exchange rate volatility on trade may be postitive or negative depending on the assumptions made with respect to risk preferences, the availability of (forward) capital markets and the time horizon of trade transactions; see e.g. Ethier (1973), Hooper and Kohlhagen (1978), De Grauwe (1988), Franke (1991), Viaene and Vries (1992) and Sercu (1992) among others. At the empirical level, the evidence is no less inconclusive. Some studies such as Chowdhury (1993), Arize (1995), Arize *et al.* (2000) and de Vita and Abbot (2004) provide evidence that increased exchange rate volatility has an adverse effect on trade due to risk-averse traders. That is, higher exchange rate volatility leads to higher costs for risk-avers traders and thus to less volume of trade. Asseery and Peel (1991), Holly (1995) and Bredin *el al*. (2003) are among those who find that exchange rate volatility affects trade positively. When trade is considered as an option held by firms – like any other option – the option value of trade may rise with exchange rate volatility and hence also export supply. Others find no evidence to suggest that exchange rate volatility has any significant impact on trade; see e.g. Aristotelous (2001). Given today's well-developed financial markets, one may argue that traders (at least to some extent) should be able to reduce or hedge uncertainty associated with exchange rate volatility. The relationship between exchange rate volatility and trade may then be weak, if not completely absent.

McKenzie (1999) gives a thorough review of the literature and discusses several empirical 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 frequency and aggregation level to employ and which estimation method to apply in each specific study at hand. As pointed out by McKenzie (1999), each of these issues and how they are handled may be part of the explanations for the inconclusive findings in the literature. In this paper, we aim to provide further evidence on the impact of exchange rate volatility on exports while trying to take account of some of the questionable issues related to previous contributions. Specifically, we study exchange rate volatility and parts of Norwegian exports within a standard demand type model relying on a cointegrated VAR approach. 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.[1] Also, Norwegian exporters have faced rather volatile exchange rates during the last decades. It is thus likely that Norwegian exporters have behaved accordingly in some manner.

The cointegrated VAR approach is particularly beneficial in the present context as different characteristics of the time series involved, which is often neglected in existing studies, can be

[1] The volume of total Norwegian exports amounted to 738 billions (at fixed 2004-prices) in 2005, which made up around 40 per cent of total GDP; see Economic Survey (2007) available at http://www.ssb.no/english/subjects/08/05/10/es/.

treated by essentially the same method. We follow both Johansen (1995) and Rahbek and Mosconi (1999) and conduct cointegration rank inference by means of (*i*) a VAR model with all variables involved being non-stationary and (*ii*) a VAR model with the measure of exchange rate volatility being a stationary regressor, respectively. To proxy the measure of volatility, we make use of the conditional variance of the exchange rate from a generalized autoregressive conditional heteroskedasticity (GARCH) model. As Pagan (1984) shows, the use of ARCH based measures of volatility may create a generated regressor problem in that whilst the model produces consistent parameter estimates, they may not necessarily be efficient.[2] Nonetheless, the use of other measures such as the often utilised moving average standard deviation of the growth of the exchange rate may lead to a measurement error problem with inconsistent estimates of the impact of risk on firms' decision making, cf. Pagan and Ullah (1988). As a test of robustness, we consider GARCH based measures of volatility based on both the nominal and the real exchange rate. Unlike most related studies, which have used aggregated data, we model exports of machinery and equipment. Hence, we do not have to constrain the volatility estimates to be similar across sectors of the economy and may avoid pitfalls of data aggregation. Finally, we pay attention to special exchange rate events and monetary policy regime shifts during the selected sample period. Particularly, the monetary policy in Norway switched from a managed floating regime to inflation targeting and a freely floating exchange rate regime early in 2001. We test by means of out-of-sample forecasting whether this regime shift did have significant effects on the exporters' behaviour and thus on the parameters of the empirical model.

Our empirical findings suggest that a reduced rank VAR – in which exports, relative prices and world market demand represent the modelled variables – explains the data quite well. We were unable to identify any statistically significant cointegrating relationship among the selected variables when the information set also included a GARCH based measure of exchange rate volatility, treated as either a stationary or a non-stationary variable in the VAR. In this respect, the distinction between nominal and real exchange rate volatility appears to be unimportant. Rather, we find that volatility changes proxied by blip dummies connected to the monetary policy change from a fixed to a managed floating exchange rate and the Asian financial crises during the 1990s enter significantly in a dynamic model of export growth – in which the level of relative prices and world market demand together with the level of exports comprise a significant cointegration vector. We also demonstrate that the dynamic model performs well out-of-sample, a finding which rejects the hypothesis that increased exchange rate volatility following the introduction of inflation targeting has had a significant impact on export performance.

The rest of the paper is organized as follows: Section 2 outlines the economic background underlying the empirical analysis. Section 3 describes the data used, while Section 4 presents the volatility measures applied in our context. Section 5 reports results from the cointegrated VAR approach and Section 6 presents a parsimonious export model with particular emphasis on its economic content and forecasting ability. Section 7 concludes and points out directions for future research.

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[2] Pagan (1984) does not consider the generated regressor problem in the case of GARCH based measures of the kind used here. McKenzie (1999) points out, however, that the consistency property of estimated parameters in models with ARCH generated regressors extends to cases of more complicated conditional variance models.

2. The economic background

A demand model for exports is basically like any other demand model. Equilibrium price and quantity are determined by the interaction of supply and demand. Usually the assumption of infinitely elastic export supply and so exogenous own export price is made in previous empirical studies; see Arize (1995), Sukar and Hassan (2001), Bredin *et al*. (2003) and de Vita and Abbott (2004) among others. In this paper, we pursue these studies and draw on Boug *et al*. (2006) who provide evidence that Norwegian exporters of machinery and equipment follow much more closely the prices of competitors than domestic costs in setting their export prices. Accordingly, we do not consider the export supply model and specify the following export demand model to be adequate in the present context:

$$(1) \qquad X_t^e = f(P_t / P_t^*, Y_t^*, V_t),$$

where X_t^e denotes the volume of exports in period *t*, P_t / P_t^* represents the relative price (export competitiveness) between own export price in period *t* (denominated in domestic currency) and the price of foreign substitute goods in period *t* (multiplied by the nominal exchange rate), respectively, Y_t^* denotes foreign demand and V_t is a measure of exchange rate volatility. Some previous studies have extended the scope of the underlying model similar to (1) by additional explanatory variables such as the distance between trading countries, transport costs, consumer tastes, foreign direct investments and third country exchange rate risk effects – see e.g. Thursby and Thursby (1987), Cushman (1986) and Égert and Morales-Zumaquero (2005) – in an attempt to enhance the empirical analysis. That third country effects may matter relates to the possibility of an increase in exchange rate risk in one currency, biasing traders' decisions in favour of another country in which the exchange rate exhibits a lower level of volatility. As pointed out by McKenzie (1999), however, the results obtained by these extended models do not differ substantially from those which have been established before them. Taking logs of the variables, assuming log-linearity, we write (1) as

$$(2) \qquad x_t^e = \beta_0 + \beta_1 (p - p^*)_t + \beta_2 y_t^* + \beta_3 v_t + \varepsilon_t,$$

where ε_t is a stochastic error term.[3] Standard trade theory predicts that increases in relative prices should depress exports, while foreign demand pressure should affect exports positively, hence $\beta_1 < 0$ and $\beta_2 > 0$. As noted in the introduction, the impact of exchange rate volatility on export performance is both theoretically and empirically less clear-cut. The focus of early work is on models with risk-averse firms and spot exchange rate changes representing the only source of risk for the economy; see e.g. Hooper and Kohlhagen (1978). Firms operating across borders are concerned with spot exchange rate fluctuations because currency values

[3] In what follows, lower case letters indicate natural logarithms of a variable, unless otherwise stated.

partly determine the price paid or received for goods and services. Unexpected variations in the spot exchange rate thus make prices and costs uncertain and hence also the unhedged profits (the variance of profits increases) because of the time lag between the purchase order and the payment date. Under such conditions, risk-averse firms prefer to reduce exports as they wish to reduce their risk exposure. In the presence of hedging opportunities in financial markets firms may reduce exchange rate risk and thereby the depressing impact of exchange rate volatility on exports. Nevertheless, the failure to provide a perfect hedge in many instances and the fact that hedging is not free of charge make the likely end result of exchange rate volatility a reduction in exports in the case of risk avers agents.[4] The focus of later work is on the effect of exchange rate volatility on *expected* profits; see e.g. Giovannini (1988), De Grauwe (1994) and Bredin *et al.* (2003). If profits are a convex function of the exchange rate, then increased exchange rate variability may lead to increased expected profits, and as such explain a positive relationship between exports and exchange rate volatility – especially if exporters are risk neutral.[5] The positive relationship between exchange rate variability and exports may still hold for risk-averse firms, provided that the increase in firms' utility from increased expected profits more than offsets the decline in utility from greater uncertainty of profits. Another approach explaining a positive relationship between exports and exchange rate variability views exports as an option to be exercised by firms; see e.g. Franke (1991). Like any other option, the option value of exports increases when the variability of the underlying asset increases. If increased exchange rate volatility causes expected profits to increase, then the value of the option to export has increased, which may lead firms to raise export supply.

Since no consensus about the effects of exchange rate volatility on trade exist in the literature, we do not have any a priori beliefs about the sign and magnitude of β_3 in (2). We notice that (2) is a *static* model of exports that may form a cointegration relationship among the level variables with β_1, β_2 and β_3 being the long run parameters of interest. Our modelling strategy thus involves investigating the empirical counterpart of (2) by means of VAR models and well-established multivariate cointegration techniques.

3. The data

The empirical quantification of (2) is conducted using quarterly, seasonally unadjusted data of exports of machinery and equipment from Norway to the main trading partners – treated as a single destination country – for the period 1985Q1 to 2005Q4, hence covering periods of both fixed, managed floating and freely floating exchange rate regimes. The chosen sector aggregate with relatively homogeneous products accounted for around 30 per cent of total Norwegian manufacturing exports in 2005, see Economic Survey (2007).

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[4] That hedging in forward markets fails to completely eliminate exchange rate risk is discussed in e.g. Arize *et al.* (2000). The difficulty to provide perfect hedge in financial markets is inter alia related to the fact that forward rates are a poor predictor of future spot rates; see Choudhry (1999) and the references cited therein.

[5] Bredin *et al.* (2003) provide a formal model and dicuss in detail the conditions under which increased exchange rate variability would lead to increased exports.

When Norway left the European exchange rate agreement at the end of the 1970s and established a currency basket, the Norwegian krone showed relatively high variability during the 1980s. Following a 12 percent devaluation of the krone in May 1986 a flexible interest rate policy was introduced with the explicit goal of a fixed exchange rate. However, the krone experienced significant revaluations and devaluations during the first decade of our sample period. After the turmoil following the speculative attacks against the krone by the end of 1992, Norway changed to a managed floating exchange rate regime in which the exchange rate was allowed to freely float within given target bands. The exchange rate was still sensitive to special events such as the Asian financial crises in the second half of the 1990s.[6] As already mentioned, Norway formally changed to freely floating exchange rates following the introduction of inflation targeting in late March 2001. Overall, Norway has experienced highly varying nominal and real exchange rates during the selected sample period crosswise different monetary policies, see Boug *et al*. (2006) for details. The change in monetary policy from exchange rate targeting to inflation targeting may in particular have caused the behaviour of exporters to shift in accordance with the Lucas critique. We pursue this hypothesis by performing an out-of-sample forecasting exercise to the estimated export model over the period 2001*Q*2 – 2005*Q*4.

Different approaches have been taken to approximate relative prices between domestic and foreign markets. Some use producer prices as proxies for both import prices and export prices due to data constraints. Others apply consumer prices, but then it is clear that one includes elements not directly of relevance for the traders. We employ the export price index of machinery and equipment denominated in Norwegian currency as a proxy for P_t in (2), see the Appendix for details. The construction of a proxy for P_t^* is based on the fact that Norwegian exporters face competition on the world market from exporters in other countries than the importing country. Then, the competing price facing Norwegian exporters can be thought of as a weighted sum of import price indices in the countries to which Norwegian exporters deliver goods and services. We therefore use import price indices of machinery and equipment denominated in foreign currencies, which are converted into the Norwegian currency and weighted together according to the following formula:

(3) $$P_t^* = P_{t-1}^* \cdot \left[1 + \sum_{i \in I} \left(\kappa_i \cdot \left[\frac{\left(PI^i \cdot E^i\right)_t}{\left(PI^i \cdot E^i\right)_{t-1}} - 1 \right] \right) \right],$$

where $I = \{$Sweden, the Euro area, the United Kingdom, the United States and Japan$\}$ includes five of Norway’s most important trading partners based on the OECD trade weights κ_i [7], the

[6] The financial crises, which broke out in Thailand in July 1997, spread itself quickly to several countries in the southeast of Asia during the following autumn and successively to the rest of the world, mainly through lower domestic demand in the troubled Asian economies. Consequently, international trading partners faced reduced export possibilities, which were further amplified by stronger competition in the wake of falling exchange rates in the Asian economies, with downward pressure on prices for trading partners and thus lower earnings, see Economic Survey (1998) available at http://www.ssb.no/english/subjects/08/05/10/es/9801/.

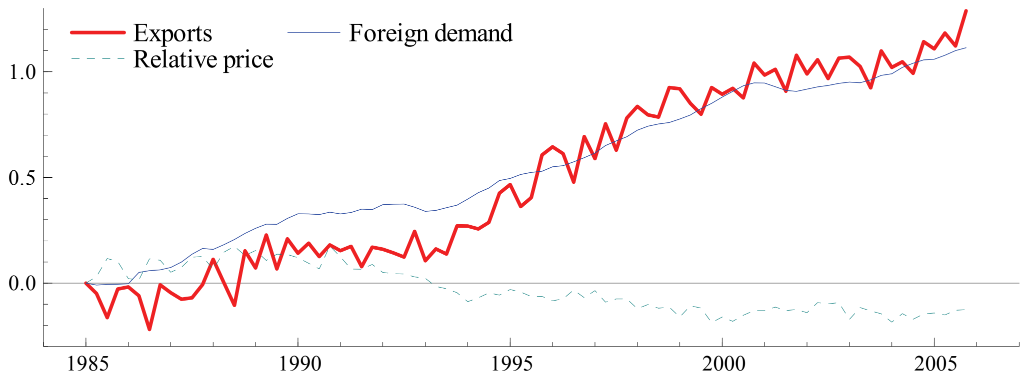
[7] Excluding the smaller trading partners these weights are 23.4 per cent for Sweden, 49.4 per cent for the Euro area, 14.1 per cent for the United Kingdom, 7.5 per cent for the United States and 5.6 per cent for Japan. These weights are kept time independent in (3) as they have been quite stable throughout the sample period. See the Appendix for further details about the underlying series in (3) and their sources.

initial value of P_t^* is normalised to unity ($P_0^* = 1$), PI_t^i denotes the import price index of machinery and equipment in country i at time t denominated in foreign currency and E_t^i represents the bilateral nominal exchange rate between the currencies of Norway and country i at time t. According to (3), the competing price facing Norwegian exporters of machinery and equipment is computed as a trade weighted sum of the growth rates of the Norwegian currency denominated import price indices of the main trading partners given the initial value of unity of P_t^*. Due to data constraints, producer prices are used in the case of the Euro area. Typically, both industrial production and gross domestic product are used as proxies for foreign demand in previous studies. We approximate Y_t^* in (2) as follows:

(4) $$Y_t^* = \sum_{i \in I} \kappa_i Y_t^i,$$

where Y_t^i denotes total imports in country i at time t. Our proxy for foreign demand thus is calculated by weighting together imports of the main trading partners with weights identical to those used in the construction of P_t^*. Figure 1 displays the log of the volume of exports (x_t^e), the log of the relative price $(p - p^*)_t$ and the log of foreign demand (y_t^*) over the sample period.[8] We observe that the three time series exhibit a clear trending behaviour, but with no apparent mean reverting property, suggesting that exports, relative prices and foreign demand are all non-stationary $I(1)$ series. Therefore, a reduced rank VAR is a candidate as an empirical model. However, we also need to construct and consider a measure of the exchange rate volatility to pursue the reduced rank VAR hypothesis, an empirical issue which we now turn to.

Figure 1. Machinery and equipment data (in logs)



[8] The time series are normalised to unity in 1985Q1.

4. The measure of exchange rate volatility

One frequently asked question in the literature on the impact of exchange rate volatility on trade is whether nominal or real exchange rate rate volatility enters the decision making of traders. Some argue that risk should regard nominal rather than real exchange rate risk as the latter depends in effect not only on the variance of the nominal exchange rate, but also on that of relative prices which constitute a different type of risk for traders. Volatility measures that partly reflect fluctuations in price levels thus do not distinguish between the risk associated with nominal exchange rate changes independent of price movements and the risk associated with all other factors which may affect domestic and foreign prices. Others argue that volatility based on the real exchange rate is the more relevant measure because the effects of uncertainty on a firm's profit that arise from fluctuations in the nominal exchange rate are likely to be offset in large part by movements in costs and prices, at least in the longer run.

Another often discussed question in the literature is which measure of volatility should be applied as a proxy for exchange rate risk. One of the most commonly used measures involves the moving average standard deviation of the growth of the exchange rate; see e.g. Fountas and Aristotelous (1999), Bredin *et al*. (2003) and de Vita and Abbott (2004). Such measures, however, have been questioned on the ground that they lack a parametric model for the time-varying variance of exchange rates. Moreover, as assessed by Pagan and Ullah (1988), they are likely to suffer from the measurement error problem and as such produce biased estimates of the impact of risk on the decision making of economic agents. An alternative volatility measure used with increasing frequency – which does not suffer from these shortcomings – is based on the autoregressive conditional heteroscedasticity (ARCH) model introduced by Engle (1982) and extended versions, see Kroner and Lastrapes (1993), Caporale and Doroodian (1994), Lee (1999), Sukar and Hassan (2001) and Choudhry (2005) among others. Indeed, ARCH based measures of volatility are likely to produce consistent estimates of parameters of interest in (2), but potentially inefficient ones due to the generated regressor problem, cf. Pagan (1984, theorem 12). So, for correct statistical inference, the standard errors of all the parameter estimates would need to be adjusted.

In line with the arguments above, and as a test of robustness, we choose to experiment with ARCH based measures of volatility based on both the nominal and the real exchange rate. The former is constructed as a trade weighted sum of the bilateral nominal exchange rates discussed in the previous section (with κ_i used as weights), whereas the latter is defined as the first right-hand side variable in (2). It turned out, however, that the distinction between nominal and real exchange rate volatility does not impact significantly on the cointegration results obtained.[9] Hence to save space, we only report results of the effects on exports of volatility which is expressed in terms of the nominal trade weighted exchange rate.

We assume here that the process of the nominal trade weighted exchange rate follows a random walk with drift and time-varying variance, which is modelled by the generalised

[9] These and the other test results not reported below are available upon request. Several previous studies also provide evidence to suggest that the distinction between nominal and real exchange rate volatility makes no difference to the results obtained; see e.g. Thursby and Thursby (1987) and Qian and Varangis (1994).

ARCH model of Bollerslev (1986). In this model, the conditional variance depends not only on lagged disturbances, but also on its own lagged values. Our point of departure is the following GARCH(p,q) model in order to obtain the exchange rate volatility:

$$
\begin{aligned}
&y_t = \mu + u_t, \\
(5) \qquad &u_t = \varepsilon_t h_t^{1/2}, \varepsilon_t \mid I_{t-1} \sim N[0,1], \\
&h_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i u_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j}, \quad t = 1,...,T,
\end{aligned}
$$

where y_t is equal to the log of the difference of the trade weighted nominal exchange rate and μ is the mean of y_t. Assuming $\varepsilon_t \sim N[0,1]$ for all t gives $u_t \sim N[0,h_t]$ so that y_t conditional on the past (I_{t-1}) is normal, but heteroscedastic. Estimating (5) by maximum likelihood we obtain estimates of the parameters μ, α_0, α_i (i = 1,…,q) and β_j (j = 1,…,p), and hence also the conditional variance (h_t). Noticeably, the conditional variance must be nonnegative, which requires that $\alpha_0 \geq 0$, $\alpha_i \geq 0$ (i = 1,…,q) and $\beta_j \geq 0$ (j = 1,…,p). The values of p and q may be selected on the basis of likelihood ratio tests. Table 1 shows the results from estimating the simple GARCH(1,1) model in our case.

Table 1. The estimated GARCH(1,1) model

$$h_t = \underset{(0.0000827)}{0.0000823} + \underset{(0.157)}{0.269} u_{t-1}^2 + \underset{(0.336)}{0.477} h_{t-1}$$

$$L = 224.8 \quad \hat{\alpha}_1 + \hat{\beta}_1 = 0.745$$

$$ARCH: F(2,76) = 0.679 \quad [0.510]$$

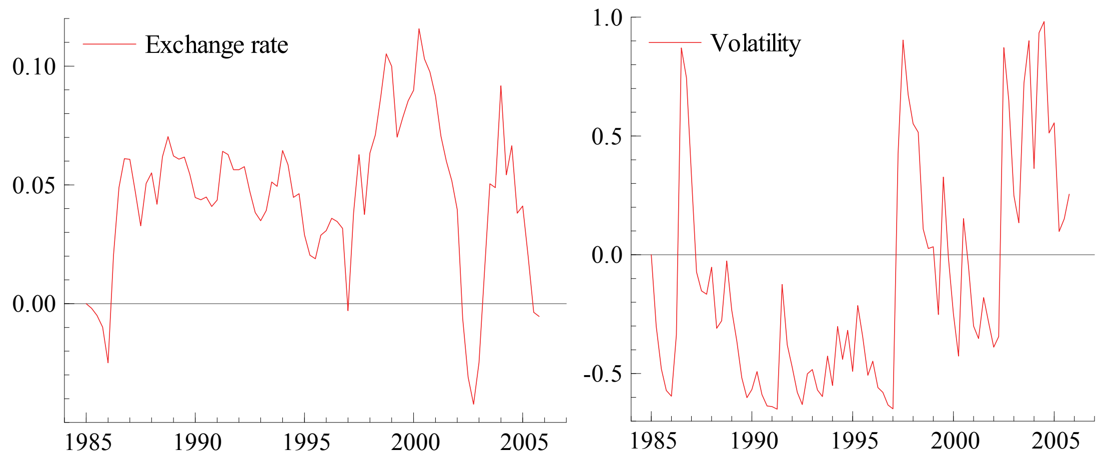
$$PORTM: \chi^2(12) = 11.32 \quad [0.502]$$

Sample period: $1985Q1 - 2005Q4$

Notes: L represents the maximum likelihood value of the model, *ARCH* is the Engle (1982) test for until second order autoregressive conditional heteroscedasticity in the residuals and *PORTM* is the Portmanteau test for serial correlation. Square brackets […] and parentheses (…) contain p-values and robust standard errors, respectively, see Doornik and Hendry (2001b, p. 14).

Owing to parameter restrictions we use one sided t-tests for the GARCH parameters. The estimates of α_1 and β_1 are significant – albeit the latter is a borderline case – at conventional levels. The diagnostic tests indicate that the residuals are well behaved. Absence of serial correlation in the residuals implies no need to fit a higher order GARCH model to the data. Moreover, results from likelihood ratio tests suggest that the composite hypothesis $p = q = 1$ is not rejected. Figure 2 shows the log of the trade weighted nominal exchange rate over the sample period together with the log of the exchange rate volatility ($v_t = \hat{h}_t = \hat{\alpha}_0 + \hat{\alpha}_1 \hat{u}_{t-1}^2 + \hat{\beta}_1 \hat{h}_{t-1}$) generated from the estimation of (5).[10] Not surprisingly – given the changes in monetary policy discussed earlier – we see that the volatility increases somewhat over time and that no clear tendency of mean revertion is apparent in the series, at least not from the early 1990s. Interestingly, the increase in volatility through 1997 and the next few years coincides well with the Asian financial crises. Also, the devaluation of the krone in the second quarter of 1986 is visible in the volatility measure. Taken together, it is likely that the volatility measure is a non-stationary $I(1)$ series. Needless to say, the time series properties of the volatility measure are not as clear-cut as exports, foreign demand and relative prices.[11] We thus consider different VAR models in our context assuming either stationarity or non-stationarity in the measure of volatility.

Figure 2. Trade weigted nominal exchange rate and volatility (in logs)



5. The cointegrated VAR

As emphasised by McKenzie (1999), only a few studies in this field have made explicit account for the stationarity status or otherwise of the data. Even fewer have considered the possibility of cointegration relationships among variables involved. Some recent studies, however, do pay attention to these econometric issues; see e.g. Sukar and Hassan (2001), Bredin *et al*. (2003), de Vita and Abbott (2004) and Choudhry (2005). The starting point of

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[10] The time series are normalised to unity in 1985Q1.

[11] A battery of augmented Dickey-Fuller tests suggest that the time series for exports, foreign demand and relative prices are all $I(1)$, whereas the volatility series is a borderline case when it comes to being integrated of order zero or one. We remark that the order of integration of the volatility series in principle is possible to deduce from the GARCH model. From (5) we have that $u_t^2 = \varepsilon_t^2 h_t$. Inserting this in the expression for h_t in (5) assuming a GARCH(1,1), known α_0, α_1 and β_1 and neglectable estimation uncertainty with respect to these parameters (due to enough available data) gives $h_t = \alpha_0 + (\beta_1 + \alpha_1 \varepsilon_{t-1}^2) h_{t-1}$. Assume now that this expression holds for $t = 1,\ldots,T$ and that $h_0 = 0$, then it follows that $h_1 = \alpha_0$, $h_2 = \alpha_0 + (\beta_1 + \alpha_1 \varepsilon_1^2)\alpha_0$, $h_3 = \alpha_0 + (\beta_1 + \alpha_1 \varepsilon_2^2)[\alpha_0 + (\beta_1 + \alpha_1 \varepsilon_1^2)\alpha_0]$ and so on. Since the distribution of ε is specified it follows what the order of integration of the volatility series would be. We thank Terje Skjerpen for pointing out this to us.

the Johansen (1995, p. 167) trace test for cointegration rank, adapted in this paper, is an unrestricted p-dimensional VAR of order k having the form

$$(6) \qquad x_t = \sum_{i=1}^{k} \Pi_i x_{t-i} + \mu + \delta t + \varepsilon_t, t = k+1,...,T,$$

where x_t is a ($p \times 1$) vector of modelled variables at time t, μ represents a ($p \times 1$) vector of intercepts, δ is a ($p \times 1$) coefficient vector of a linear deterministic trend t, $\Pi_1,...,\Pi_k$ is ($p \times p$) coefficient matrices of lagged level variables and $\varepsilon_{k+1},...,\varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) ($p \times p$) covariance matrix Ω. The initial observations of $x_1,...,x_k$ are assumed to be fixed.

The question now is how (6) can be reparameterised to a cointegrated VAR (henceforth CVAR) in which the cointegration hypothesis can be formulated as a redused rank restriction on the impact matrix $\Pi = -(I - \Pi_1 - ... - \Pi_k)$. The way the CVAR is formulated in our context depends on the assumptions made about the time series properties of the exchange rate volatility series. First, we shall consider the case when the volatility series is non-stationary, hence $x_t = [x_t^e, (p-p^*)_t, y_t^*, v_t]'$ is the relevant vector of modelled non-stationary variables. Once $x_t \sim I(1)$, then the first difference $\Delta x_t \sim I(0)$ implying either $\Pi = 0$ or Π has reduced rank such that $\Pi = \alpha\beta'$, where α and β are $4 \times r$ matrices and $0 < r < 4$. Here r denotes the order of the rank of Π. Thus, under the $I(1)$ hypothesis, assuming for notational simplicity that $k = 2$, the CVAR becomes

$$(7) \qquad \Delta x_t = \alpha\beta' x_{t-1} + \Gamma_1 \Delta x_{t-1} + \mu + \delta t + \varepsilon_t,$$

where $\beta' x_{t-1}$ is an $r \times 1$ vector of stationary cointegration relations among exports, relative prices, foreign demand and exchange rate volatility, and $\Gamma_1 = -\Pi_2$ is the (4×4) coefficient matrix of the lagged differenced variables. Next, we shall consider the case when the volatility series is treated as a stationary explanatory variable in the CVAR. Following Rahbek and Mosconi (1999), we formulate the relevant model as (again assuming $k = 2$)

$$(8) \qquad \Delta x_t = \alpha\beta' x_{t-1} + \Gamma_1 \Delta x_{t-1} + \sum_{i=0}^{2} \psi_i z_{t-i} + \mu + \delta t + \varepsilon_t,$$

where $x_t = [x_t^e, (p-p^*)_t, y_t^*]'$ now is the vector of modelled non-stationary variables and $z_t = v_t$ is the supposedly stationary regressor. As pointed out by Rahbek and Mosconi (1999), the inclusion of stationary explanatory variables as extra regressors will lead to nuisance parameters in the asymptotic distribution of the trace statistic for cointegration rank. Hence, the critical values of the trace test reported by PcGive, which are based on the assumptions underlying (7), can only be used as approximations in determining the cointegration rank. As an additional approach, we shall follow the proposal put forward by Rahbek and Mosconi (1999) and analyse an extended model of (8) given by

$$(9) \qquad \Delta x_t = \alpha\beta^{*}{}'\begin{pmatrix} x_{t-1} \\ \sum_{i=1}^{t} z_i \end{pmatrix} + \Gamma_1 \Delta x_{t-1} + \sum_{i=0}^{2} \psi_i z_{t-i} + \mu + \delta t + \varepsilon_t ,$$

where $\sum_{i=1}^{t} z_i$ is the cumulated level of the volatility series restricted to lie in the cointegration space, thereby leading to nuisance-free rank determination. Since by definition the cumulated regressor enters the cointegrating relations, the rank determination may be altered by (9) compared to (8). However, after the rank is determined, we may test the presence of the cumulated series by means of the hypothesis $\beta^{*}{}' = (\beta', 0)$ using standard χ^2 inference. As argued by Rahbek and Mosconi (1999), the associated likelihood ratio test can then be regarded as a misspecification test of the original model in (8).

Having established the different CVAR in our context, we now wish to evaluate their empirical counterparts by fitting (6) to the data with an unrestricted constant to reflect the trending behaviour in the level of the series. The linear deterministic trend is restricted to lie in the cointegrating space, thereby restricting the system to at most having a linear deterministic trend in levels of the series. In addition, centred seasonal dummies are included in the VAR unrestrictedly. Assuming $x_t = [x_t^e, (p-p^*)_t, y_t^*, v_t]' \sim I(1)$, a battery of diagnostic tests suggests that $k = 3$ is the appropriate choice of lag length to arrive at a model that produces residuals with statistically acceptable properties. A likelihood ratio test of model reduction [see Doornik and Hendry (2001a, p. 51)] from a VAR with the linear trend to a VAR without the linear trend, yields $\chi^2(4) = 6.288$ with a p-value of 0.179. So the linear trend is *in*significant at conventional tests levels in the VAR. Similarly, when $x_t = [x_t^e, (p-p^*)_t, y_t^*]' \sim I(1)$ and v_t (and its lags) instead is included as an unrestricted stationary regressor in the model, $k = 3$ is necessary to exclude any problems with autocorrelated residuals. The linear trend turns out *in*significant also in this case.[12] Table 2 reports the trace statistic for determination of the cointegration rank for the models given by (7), (8) and (9), respectively.

We clearly observe that the rank should be set to zero in the case of model (7), indicating nonexistence of any cointegration relationships between exports, relative prices, world market demand and exchange rate uncertainty. Apparently, the rank is equal to zero also in the case of model (8), whereas the rank determination when specifying the CVAR in line with (9) is

[12] The choice of lag length and the *in*significance of the linear trend are unaltered when the cumulated volatility series is included restrictedly in addition to the volatility series itself (and its lags) entering unrestrictedly in the VAR.

somewhat more ambigious.[13] Strictly speaking, the trace test statistic indicates that the hypothesis of no cointegration cannot be rejected at the 10 per cent significance level. Nevertheless, testing the hypothesis $\beta^{*}{}' = (\beta', 0)$ assuming $r = 1$ gives $\chi^2(1) = 0.048$ with a *p*-value of 0.827. Hence, the hypothesis is clearly not rejected and the volatility series does not enter cumulated in a cointegrating relationship. In this sense the model given by (8) passes the misspecification test. Following Harbo *et al.* (1998), we also considered a similar CVAR when the volatility series was assumed weakly *exogenous* for the cointegrating parameters, but still possibly entering the cointegrating space, to make inference on the rank order. Again, no formal support to any significant cointegrating vector was obtained.

Table 2. The Johansen trace test for cointegration rank determination

	Model (7)		Model (8)		Model (9)	
	λ_i	λ_{trace}	λ_i	λ_{trace}	λ_i	λ_{trace}
$r = 0$	0.197	35.89 [0.407]	0.139	18.33 [0.552]	0.140	27.10 [0.101]
$r \leq 1$	0.141	18.11 [0.568]	0.071	6.14 [0.682]	0.115	14.86 [0.061]
$r \leq 2$	0.068	5.81 [0.720]	0.002	0.16 [0.685]	0.059	4.95 [0.026]*
$r \leq 3$	0.001	0.08 [0.779]				

Notes: See Rahbek and Mosconi (1999) for details about the CVAR models (8) and (9). The underlying VAR models of the cointegration analysis have $k = 3$ with an unrestricted constant, no linear trend and no other dummies than unrestricted centred seasonals included. r denotes the cointegration rank, λ_i are the eigenvalues from the reduced rank regression and λ_{trace} denotes the value of the trace statistic, see Johansen (1995). The *p*-values in square brackets, which are reported in PcGive, are based on the approximations to the asymptotic distributions derived by Doornik (1998). The asterisk * denotes rejection of the null hypothesis at the 5 per cent significance level.

Based on the cointegration results above, we leave out the GARCH based exchange rate volatility series entirely from the information set. Instead, we shall continue the cointegration analysis by enlarging the underlying VAR with various dummy variables possibly capturing effects of exchange rate uncertainty on exports following the special exchange rate events and the various monetary policies described previously. In other words, we shall re-estimate the VAR in (6) with exports, relative prices and world market demand and the hypothetical permanent effects from the devaluation episode in 1986, the switch from a fixed to a managed floating exchange rate regime in 1993, the Asian financial crises from mid 1997 and the move to inflation targeting and a freely floating exchange rate regime in 2001 by means of permanent blip dummies. Formally, we write the enlarged VAR as

(10) $$x_t = \sum_{i=1}^{k} \Pi_i x_{t-i} + \mu + \delta t + \Phi_1 D862_t + \Phi_2 D931_t + \Phi_3 D972_t + \Phi_4 D012_t + \varepsilon_t,$$

[13] The rank tests are virtually unchanged with different lag length of the volatility series in (8) and (9).

where $x_t = [x_t^e, (p-p^*)_t, y_t^*]' \sim I(1)$, $D862_t$ equals unity for t = 1986Q2, zero otherwise, $D931_t$ equals unity for t = 1993Q1, zero otherwise, $D972_t$ equals unity for t = 1997Q2, zero otherwise and $D012_t$ equals unity for t = 2001Q2, zero otherwise. These four dummies enter the VAR equations unrestrictedly. Diagnostic tests reveal that k = 2 is now the appropriate lag length to produce a model with no serious misspecification. Again, the linear trend becomes *in*significant in the model and is thus dropped in the successive rank determination of the CVAR representation of (10). Likewise, the estimated effects of $D862_t$ and $D012_t$ are far from being significant and can thus be removed from the model without much influence on the successive test results.[14] The estimates of $D931_t$ and $D972_t$ are, however, strongly significant in the export equation. Interestingly, these dummies correspond quite well with the dummies detected in Bjørnland and Hungnes (2006).[15] Table 3 reports the trace statistic for cointegration rank determination of the CVAR in line with (10). Based on the 10 per cent critical value, we conclude that the rank is equal to unity, which suggests that one cointegration vector exists between exports, relative prices and world market demand. We remark that the unrestricted vector is virtually identical to the comparable part of the unrestricted vector (assuming r =1) when the GARCH based exchange rate volatility series belongs to the information set, hence adding force to the assumption that the volatility series may be excluded from the VAR without loss of relevant information. When the unrestricted cointegration vector is equipped with long run weak exogeneity restrictions on relative prices and world market demand, that specification is clearly not rejected as indicated by $\chi^2(2)$ = 1.243 (with a p-value of 0.537). Besides, these restrictions do not change the estimated long run coefficients substantially.

Table 3. The Johansen trace test for cointegration rank determination of (10)

	λ_i	λ_{trace}
$r = 0$	0.170	27.32 [0.096]
$r \leq 1$	0.135	12.07 [0.155]
$r \leq 2$	0.002	0.17 [0.683]

Notes: The underlying VAR of the cointegration analysis has k = 2 with no trend and a constant, centred seasonals, $D892_t$, $D931_t$ and $D972_t$ included unrestrictedly. r denotes the cointegration rank, λ_i are the eigenvalues from the reduced rank regression and λ_{trace} denotes the value of the trace statistic, see Johansen (1995). The p-values in square brackets, which are reported in PcGive, are based on the approximations to the asymptotic distributions derived by Doornik (1998). It should be noted that the unrestricted permanent blip dummies included in the VAR, which unlike unrestricted shift dummies (…,0,0,0,1,1,1,…) do not cumulate to (broken linear) trends in the levels of the data, are not likely to affect the asymptotic distribution of the reduced rank test statistics, see Juselius (2006, p. 139).

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[14] The issue of possible effects on exports of the move to inflation targeting in 2001 is pursued further in the next section.

[15] We should mention that an additional blip dummy $D892_t$ – albeit with no particular economic rationale – enters the VAR unrestrictedly to mop up extraordinary large residuals in 1989Q2. Noticeably, including $D892_t$, $D931_t$ and $D972_t$ does not change the estimates of Π_1, Π_2 and Π dramatically compared to the no-dummy VAR. We also notice that the blip dummies in (10) appear less significant when modelled as transitory rather than permanent variables, see Juselius (2006, p. 106).

The estimated cointegration vector is given in (11), with standard errors in parentheses just below the coefficients. We interpret the vector as a long run export relation as the estimated coefficients for relative prices and world market demand are both statistically significant and economically reasonable with expected signs. If the Norwegian export price is high relative to the competing price, then the equilibrium volume of exports must be low, consistent with low competitive strength in the foreign markets. Analogously, if world market demand is high, then the equilibrium volume of exports must be high.

(11)
$$x^e = \underset{(0.39)}{-1.18}(p - p^*) + \underset{(0.12)}{0.98}y^*$$

Furthermore, the loading coefficient linked to (11) is strongly significant ($t = -3.64$), meaning that the cointegrating vector enters significantly in the export relation in the CVAR. And, the equilibrium correction term crosses its mean value several times over the sample period, which is a further strong indication of the existence of cointegration. Since relative prices and world market demand seem to be weakly exogenous with respect to the cointegrating vector, the empirical analysis can proceed within a single equilibrium correction model (EqCM) for the growth in exports, which we now turn to.

6. A parsimonious export model

Equipped with the long run equilibrium relationship, we proceed to estimate a parsimonious export model relying on the general-to-specific approach. Our point of departure is a general EqCM model of exports, consistent with the reduced rank VAR model in the previous section, written as

(12)
$$\begin{aligned}\Delta x_t^e &= \varphi_{11}\Delta x_{t-1}^e + \sum_{i=0}^{1}\varphi_{2i}\Delta(p - p^*)_{t-i} + \sum_{i=0}^{1}\varphi_{3i}\Delta y_{t-i}^* \\ &\quad + \lambda[x^e + 1.18(p - p^*) - 0.98y^*]_{t-1} + \varphi_0 D_t + \varepsilon_t,\end{aligned}$$

where D_t is a vector containing all the deterministic components (the constant, the centred seasonals and the blip dummies). The general model thus contains contemporaneous and one lag of the difference of each of the explanatory variables involved in addition to one lag of the difference of the explained variable. Also, the EqCM term – which is based on (11) – is included, lagged one period. Simplifications from the general to the specific model is performed using PcGets, see Hendry and Krolzig (2001). Briefly speaking, PcGets first tests the general model for misspecification to ensure data coherence. If data coherence is satisfied, then the general model is simplified by excluding statistically insignificant variables. Since PcGets controls for any invalid reduction by means of diagnostic tests, the specific model choice will not loose any significant information about the relationship from the available data set. As a result, the specific model parsimoniously encompasses the general model and is not dominated by any other model. PcGets picks the following specific model in our case, with estimated standard errors just below the parameter estimates:[16]

$$\Delta x_t^e = const. - \underset{(0.078)}{0.293}\Delta x_{t-1}^e - \underset{(0.134)}{0.611}\Delta(p-p^*)_t - \underset{(0.064)}{0.266}[x^e + 1.18(p-p^*) - 0.98y^*]_{t-1}$$

(13)

$$- \underset{(0.022)}{0.082}S1_t - \underset{(0.014)}{0.118}S2_t - \underset{(0.014)}{0.169}S3_t + \underset{(0.043)}{0.131}D892_t - \underset{(0.042)}{0.135}D931_t + \underset{(0.042)}{0.114}D972_t,$$

where $S1_t$, $S2_t$ and $S3_t$ denote centred seasonal dummies. The model shows that all picked variables enter significantly. Particularly, the equilibrium correction term enters the model with a *t*-value of –4.2, which supports the conclusion from the cointegration analysis. The equilibrium correction term has the expected sign, such that the volume of exports eventually adjusts towards its equilibrium level. The significant negative effect on exports from the blip dummy in 1993 may – as interpreted above – reflect increased exchange rate uncertainty following the transition to a new monetary policy regime with a managed floating exchange rate. On the other hand, the significant positive effect on exports from the blip dummy in 1997 may seem puzzeling. A severe appreciation pressure against the Norwegian krone at about the time when the Asian financial crises broke out should in principle have reduced and not increased exports due to reduced price competitiveness. Also, lower domestic demand in the depressed Asian countries should have given rise to less export possibilities for the trading partners. However, Norwegian exporters at that time sold most of their products to countries in Western Europe and not to Asian countries with depreciated currencies. Neither the European countries exported much to the Asian regions.[17] In this sense, Norwegian produsers were not much, if at all, negatively affected by the Asian financial crises. Rather, it may have been the case that Norwegian exporters *expected* the appreciation pressure to be less profound in the wake of the Asian crises. A positive effect on exports from the blip dummy in 1997 may then be understood from the fact that Norges Bank historically has reduced the interest rate in periods of appreciation pressure. Certainly, the trade weighted nominal exchange rate depreciated quite substantially through 1997 (see Figure 2) along with the reduction in the interest rate ahead of time[18], thereby stimulated export possibilities for Norwegian producers.

The model implies further that the short run price elasticity is smaller than its long run counterpart, which is consistent with no overshooting in the volume of exports to shocks in relative prices. A one per cent increase in relative prices will cause the volume of exports to decrease by 0.6 per cent contemporaneously, and then adjust gradually to its new equilibrium level. There is also a significant negative short run autoregressive effect in the model,

[16] Although the sign and magnitude of the estimated coefficiens are hardly affected by the presence of $D892_t$, it is included in the specific model to remove borderline autocorrelation in the residuals at conventional levels.

[17] The share of total exports to the troubled Asian countries was only between 5 and 10 per cent for the European countries; see Economic Survey (1998).

[18] See http://www.norgesbank.no/Pages/Article____55476.aspx.

represented by Δx^e_{t-1}. Finally, Table 4 reports that the model shows no sign of misspecification. This model property is further reinforced by recursive break point Chow statistics and recursively estimated coefficients, which exhibit constancy.

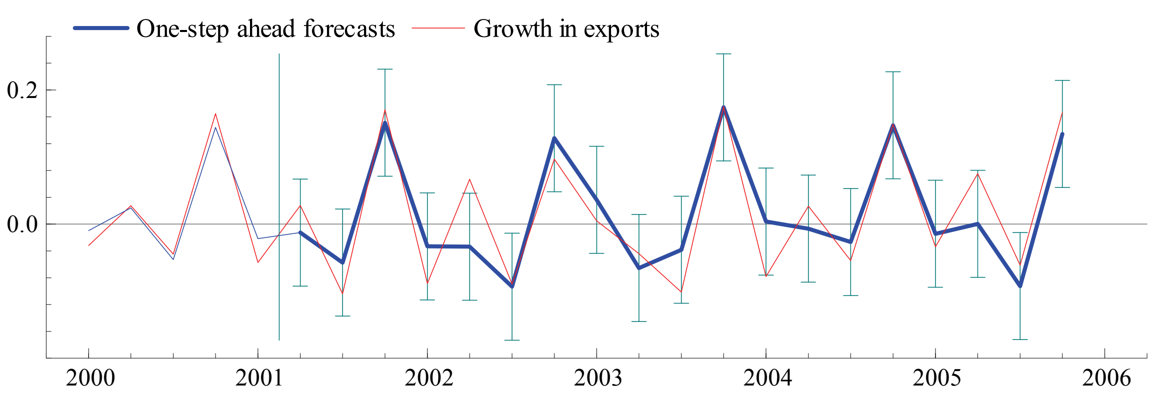
Table 4. Diagnostic tests for model (12)

AR_{1-5}	$F(5,67)$ = 1.015	[0.416]
$ARCH_{1-4}$	$F(4,64)$ = 1.006	[0.411]
$NORM$	$\chi^2(2)$ = 2.442	[0.295]
HET	$F(12,59)$ = 0.629	[0.809]
$RESET$	$F(1,71)$ = 0.012	[0.912]

Notes: AR_{1-5} is Harvey's (1981) test for until 5th order residual autocorrelation; $ARCH_{1-4}$ is the Engle (1982) test for until 4th order autoregressive conditional heteroskedasticity in the residuals; *NORM* is the normality test outlined in Doornik and Hansen (1994), *HET* is a test for residual heteroskedasticity due to White (1980) and *RESET* is the Ramsey (1969) test for functional form misspecification. Figures in square brackets are *p*-values.

We have established that the parsimonious model is well specified in-sample. A natural question to ask would then be whether the model is able to predict exports out-of-sample to shed light on its robustness with respect to the monetary policy regime change in late March 2001. As noted in the introduction, the switch from exchange rate targeting to inflation targeting – which has without doubt brought about increased exchange rate volatility (see Figure 2 above) – may in particular have influenced the exporters' behaviour in a significant way. So, if exchange rate volatility did play a role for the export performance in the wake of the regime shift, we should expect instabilities in the estimated model as, for example, indicated by poor out-of-sample forecasting ability. Here we shall use simple one-step ahead forecasts by reestimating (13) based on observations until 2001Q2 and retaining nineteen quarters (2001Q2 – 2005Q4) for out-of-sample forecasts, see e.g. Hendry and Doornik (2001, p. 62) for details. Figure 3 depicts actual values of Δx^e_t together with one-step ahead forecasts, adding bands of 95 per cent confidence intervals to each forecast in the forecasting period.

Figure 3. Actual values of Δx^e_t and one-step ahead forecasts with 95 per cent bands



The model forecasts only misses significantly the observed values once (albeit a borderline case), namely in the second quarter of 2002. The point in time of the forecasting failure does not, however, coincide with the time of the formal change in the monetary policy regime. Also, a Chow test statistic of parameter constancy between the sample and the forecasting periods, cf. Hendry and Doornik (2001, p. 241), is far from being significant; with $F[19, 53] = 1.059$ and the corresponding p-value of 0.417. We therefore conclude that the out-of-sample forecasting ability of the parsimonious model is satisfactory despite a major regime shift in monetary policy. The fact that increased exchange rate volatility following the new monetary policy regime does not seem to have influenced the export performance significantly may reflect that agents' expectation formation already had been changed gradually since the beginning of 1999 at the time when the present central bank governor was appointed. Several Norwegian economists have argued that the regime change indeed took place in a gradual manner from early 1999 and that agents at the time of the formal change in the monetary policy already had learned and experienced (at least to some extent) how Norges Bank was likely to behave under inflation targeting. Accordingly, one may argue that export performance should have been affected, if at all, in 1999Q1 rather than in 2001Q2. To shed some light on these arguments, we reestimated (13) based on observations until but not including 1999Q1 and redid the forecasting exercise for the period 1999Q1 – 2005Q4. It turned out, however, that the reestimated model is virtually unchanged from (13) with respect to parameter estimates, diagnostics as well as forecasting abilities.

7. Conclusions

In this paper we have investigated the impact of exchange rate uncertainty on export performance within different CVAR models using sector specific Norwegian data for machinery and equipment. The underlying theoretical model was a demand type model with relative prices, world market demand and exchange rate volatility (or uncertainty) as explanatory variables. As a measure of volatility we used the estimated conditional variance of both the nominal and real exchange rate generated from a GARCH model.

Our empirical findings indicate that the causal link operating from exchange rate uncertainty to export performance is at best weak if present at all. This finding agrees with a number of studies on the topic even though predecessors have mostly analysed aggregated trade flows and typically neglected time series properties with respect to the volatility measure. We were unable to establish any statistically significant cointegration vector when the GARCH based volatility measures were included in the information set, treated as either stationary or non-stationary in the VAR. Instead, we found that volatility changes proxied by blip dummies in relation to the monetary policy change from a fixed to a managed floating exchange rate and the Asian financial crises during the 1990s enter significantly in a dynamic model for export growth – in which the level of relative prices and world market demand together with the level of exports form a significant cointegration relationship. In the same model we also found that a blip dummy for the recent shift in monetary policy from exchange rate targeting to inflation targeting is not significant. A forecasting exercise revealed further that the export model performs well out-of-sample, a finding which is at odds with the hypothesis that increased exchange rate volatility following the latter change in the monetary policy had a significant influence on export performance.

Our empirical analysis is based on the assumption of a linear relationship between the variables of interest. Possible nonlinear nature of causal links between exports and exchange rate uncertainty may very well be the case. We leave this issue for future work. A further extension to the present analysis would be to consider a broader selection of industries in the economy and the trade relations between Norwegian exporters and each of the trading partners on a bilateral basis. Such an avenue of future research would reveal whether any causal link between export performance and exchange rate uncertainty differs across sectors and markets of destination.

Data appendix

X^e — Volume of exports of machinery and equipment, measured as an index (1985Q1=1). Source: Quarterly national accounts, Statistics Norway.

P — Export price index (1985Q1=1) of machinery and equipment, expressed in the Norwegian currency. Source: Quarterly national accounts, Statistics Norway.

P^* — Trade weighted competitive price index (1985Q1=1) of machinery and equipment, expressed in the Norwegian currency, calculated on the basis of import price indices denominated in foreign currencies, bilateral nominal exchange rates and OECD trade weights of the main trading partners Sweden, the Euro area, the United Kingdom, the United States and Japan, see formula (3) in the main text. Sources: EcoWin (Reuters) database and individual country's own statistical online services:
http://www.scb.se, http://www.boj.or.jp, http://www.statistics.gov.uk, http://www.bls.gov, http://www.europa.eu.int/comm/eurostat, http://www.unstats.un.org/unsd/default.htm, http://www.norges-bank.no.

Y^* — Trade weighted imports of the main trading partners, measured as an index (1985Q1=1) and based on the OECD trade weights used in the construction of P^*, see formula (4) in the main text. Source: Statistics Norway.

V — GARCH based measure of exchange rate volatility based on the trade weighted nominal exchange rate calculated by means of the OECD weights and the bilateral nominal exchange rates of the main trading partners; see Section 4 in the main text. Sources: Statistics Norway and Norges Bank.

$D862$ — Blip dummy used to account for hypothetical effects on exports from the devaluation episode in May 1986. It equals unity in the second quarter of 1986, zero otherwise.

$D892$ — Blip dummy used to account for an outlier in the VAR. It equals unity in the second quarter of 1989, zero otherwise.

$D931$ — Blip dummy used to capture effects on exports of the transition from the fixed to the floating exchange rate regime following the speculative attacks against the Norwegian krone in December 1992 and January 1993. It equals unity in the first quarter of 1993, zero otherwise.

$D972$ — Blip dummy used to capture effects on exports of the Asian financial crisies which broke out in July 1997. It equals unity in the second quarter of 1997, zero otherwise.

$D012$ — Blip dummy used to capture hypothetical effects on exports of the transition from exchange rate targeting to inflation targeting at the end of the first quarter of 2001. It equals unity in the second quarter of 2001, zero otherwise.

Si	Centred seasonal dummies, where quarter i = 1, 2, 3, $Si = 0.75$ in quarter i and −0.25 in quarters $i+1$, $i+2$ and $i+3$.

Further details of data definitions and sources can be found in Fagereng (2007, ch. 3). The data are available upon request.

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