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A MONETARY PERSPECTIVE ON INFLATION DYNAMICS IN NORWAY

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***Abstract**: This paper investigates a monetary perspective on Norway's inflation dynamics over the period of 1987- 2011. A set of Norway's quarterly data on its money market, price inflation and monetary policy is analyzed. A cointegration analysis reveals two long-run economic relationships, which can be interpreted from the viewpoints of money market equilibrium and an empirical monetary policy rule. Demonstrated in this research is that disequilibrium in the money market contributes to generating a long-run inflationary impetus. A preferred vector equilibrium correction model is judged as a data-congruent monetary representation of overall inflation dynamics, and thus can be utilized for applied purposes such as forecasting.*

***Keywords:** Inflation dynamics, Monetary policy rule, Cointegrated vector autoregressive analysis, Vector equilibrium correction model*

***JEL Classification:** C32. E41. E52*

1. INTRODUCTION

This paper aims to explore the monetary aspect of overall inflation dynamics in the Norwegian economy over the period of 1987-2011. A detailed multivariate time series analysis is conducted with a view to pursuing this empirical objective. As a result of a system-based cointegration analysis and the subsequent model reduction, this study achieves a data-congruent vector equilibrium correction model (VECM) for Norway's inflation under changing regimes of monetary policy; a historical review of Norway's monetary policy is provided in Section 2 of this paper. Based on our econometric investigations, we propose that the data of aggregate money contain information useful for explaining the dynamics of overall inflation. Also in this article, a Taylor-type policy rule is demonstrated as relevant towards grasping Norway's monetary policy, although there are some studies, such as Svensson (2003, 2006), which appear to disapprove such instrument rules for the Norwegian case. In this introductory section, a brief review of related literature is provided, combined with an explanation of motivations for the econometric research pursued in this paper.

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A number of the existing empirical studies on Norway's price inflation have been conducted along with those on its wage inflation, by following the tradition of the so-called "Scandinavian model" of inflation. See Aukrust (1977) for further details regarding this model. Thus, as shown by Eitrheim (1998) and Bårdsen *et al*. (2003), *inter alia*, various wage-price models have been developed for the Norwegian economy thus far. From the viewpoint of a detailed multivariate time series analysis, it is noteworthy that Bårdsen *et al*. (2003) estimate a well-formulated dynamic wage-price model by incorporating inflationary impetus from the labor market. They succeed in quantifying the transmission mechanisms that contribute to forecasting inflation. The empirical success of Bårdsen *et al*. (2003) is so remarkable that it may seem that there is little to be added to their study. However, by virtue of exploring the time series data covering the first decade of the 21st century, some additional evidence may still be found. In other words, we may be able to show that quantitative information on some other macroeconomic variables, such as monetary aggregates, contributes to the account of Norway's inflation dynamics. Bårdsen *et al*. (2003, p.432) seem to imply the possibility that monetary instruments may have been employed in pursuit of inflation targeting policy since the end of 1992, when the Norwegian krone started to float. This possibility, hence, offers motivation for our empirical investigations into the roles played by aggregate money in Norway's inflation over recent years.

Turning our attention to monetary policy rules in general, we recognize that Taylor (1993) is the seminal paper in this field, laying the foundations for a number of noteworthy studies about monetary policy rules; see, for example, Ball (1997), Clarida *et al.* (1998, 2000), Leitemo and Söderström (2001) and Christensen and Nielsen (2009), *inter alia*. Leitemo and Söderström (2001) are of particular relevance to our research, as they show that adding the exchange rate to an optimized Taylor rule gives only small improvements in terms of economic stability in most model configurations. This result may allow us to focus on the analysis of a small-scale econometric system excluding the exchange rate. Regarding the case of Norway, Olsen *et al*. (2003) argue that, with the exception of certain brief periods in the mid-1990s, monetary policy in Norway from 1993 onwards can be described as following close to a Taylor-type policy rule. They demonstrate that Norway's interest rate responds significantly to its inflation and output gap. This finding provides a research direction to seek an empirical Taylor-type rule function embedded in Norway's data.

Next, the issue of estimating empirical money demand functions needs clarification. Since grasping the underlying structure of the money market is essential for implementing effective monetary policy, the empirical exploration of money demand functions is one of the most important research themes in macroeconomic studies. A cointegrated vector autoregressive (CVAR) analysis, introduced and developed by Johansen (1988, 1996), allows us to explore wide-ranging research objectives in applied economics; see also Johansen and Juselius (1990), Hendry and Mizon (1993), Bårdsen *et al*. (2003), Burke and Hunter (2005) and Juselius (2006), *inter alia*. With regard to research on money demand in Norway, the following articles should be noted: Bårdsen

(1992), Eitrheim (1998), Bårdsen and Klovland (2000) and Eitrheim (2003). Bårdsen (1992) achieves a demand function for narrow money in dynamic error correction models. Eitrheim (1998) applies cointegration analysis to investigate long run relationships between money, prices, real output, interest rates and wages. Bårdsen and Klovland (2000) find a stable money demand function in Norway over the period of 1966-1994, although Norway went through a deregulation process of the financial system; they also argue for the importance of the information contained in the monetary aggregates. Lastly, Eitrheim (2003) analyzes a demand-for-money function and the role of money in the inflation process in Norway.

As reviewed above, we recognize that monetary policy rule and money demand functions tend to be regarded as separate research topics in macroeconomic literature, although both of these should be simultaneously related to various important elements of monetary economies under study. Brüggemann (2003) as well as Choo and Kurita (2011, 2012), however, appear to be seen as exceptional in the literature in that they achieve meaningful empirical outcomes in the joint modeling of monetary policy rule and money demand functions. Brüggemann (2003) adopts a structural or joint-modeling CVAR approach to these research themes for Germany, while Choo and Kurita (2011, 2012) employ the same methodological approach to the analysis of South Korea and New Zealand's data, respectively. See also Juselius and MacDonald (2004) for an application of this approach to other important research themes such as international parity conditions.

Following this line of research, this paper employs a joint-modeling CVAR approach to Norway's data associated with monetary policy rule and money demand functions, with a view to obtaining an empirical monetary model for inflation dynamics. It is true that the recent literature on monetary policy and inflation tends to assign less importance to monetary aggregates than before; see Nelson (2003) and Woodford (2008), *inter alia*. Indeed, Eitrheim (2003) indicates that Norway's monetary measures do not seem to play a significant role as predictors of its future inflation. However, as mentioned above, we aim to find empirical evidence supporting the view that aggregate money contains quantitative information in accounting for the dynamics of inflation data. This paper demonstrates that the supporting evidence is indeed revealed through a detailed analysis of Norway's time series data covering the period of 1987-2011.

Our research is organized into six sections. Section 2 presents a historical review of Norway's monetary policy, and Section 3 gives a brief review of a CVAR analysis. Section 4 then introduces a theoretical framework for the money market, inflation and a monetary policy rule. Section 5 conducts a formal CVAR analysis of quarterly time series data from Norway in order to seek our primary research objectives. Section 6 reduces the CVAR system to a VECM and proves its data-congruency. Concluding remarks are provided in Section 7. All the econometric analyses and graphics in this paper use *OxMetrics/PcGive* (Doornik and Hendry, 2007).

2. HISTORICAL REVIEW OF NORWAY'S MONETARY POLICY

From the early 1970s, the structure of the Norwegian economy transformed to include a huge oil exporting industry. A value-added-tax system was introduced in 1970, and incomes policies were implemented in the late 1970s. Bank credit aggregates for industries played a critical role as intermediate targets of monetary policy in the 1970s and the early 1980s. Interest rates were usually kept lower than the market-clearing level.

Norway pursued extensive deregulation of financial markets from the mid-1980s. Selective and direct controls of bank credit were abolished, and capital controls were removed. Interest rates were gradually determined by market forces. Financial deregulation led to a rise in interest rates and a surge in bank lending, fueling price inflation. Then a stock market crash, a collapse of economic activity and failures of several major banks followed. A banking crisis took place in the late 1980s and peaked in the early 1990s. The banking crisis was accompanied by a severe recession. See Bårdsen (1992) and Bårdsen and Klovland (2000).

In 1986, targets for monetary policy changed from bank credit volume and interest rates to the exchange rate. After a 12% devaluation of the Norwegian krone in May 1986, the exchange-rate targeting was introduced along with a flexible interest rate policy. The authorities supported fixed exchange rates in trade-weighted terms or, as of October 1990, against the ECU or, as of January 1999, the euro (Bårdsen and Klovland, 2000, p.565). When a depreciating (appreciating) pressure arose, Norges Bank increased (decreased) the interest rate, resulting in a rise in the volatility of interest rates.

Gjedrem (2002) notes that the responsibility for interest rate decisions was delegated to Norges Bank in 1986. According to the OECD (2007), monetary policy in Norway was geared towards stabilizing the exchange rate, while fiscal policy was responsible for stabilizing the economy. Over the period of 1976-1986, ten devaluations were made, which led to price inflation and inflationary expectation. It is pointed by Bergo (2007) that the credibility of the fixed exchange rate regime was also impaired by the devaluations. During the disturbances in the European financial markets in 1992, the krone was depreciated and allowed to float under a managed float. Then, a free float followed the speculative attacks in 1997 and 1998.

In 2001, Norway ceased a regime of exchange-rate targeting and instead adopted inflation targeting. Its inflation targeting is flexible, in the sense that weights are given to both the inflation gap (the gap between inflation and the inflation target) and the output gap; see Svensson (2000). Norges Bank announced the key policy rate, the so-called sight deposit rate, as the instrument rate. The Consumer Price Index (CPI) is the basic measure of inflation and the operational target of monetary policy. Since 2001, Statistics Norway has also prepared another index CPI-ATE which is adjusted for tax changes and excludes energy prices. Since 2005, Norges Bank has published its own forecast for interest rate, along with forecasts for the output gap, inflation and other key variables. It presents a probability distribution ("fan chart") around the point forecast. When inflation targeting was adopted, a newly created fiscal rule was made effective as of 2002; the fiscal rule specified a gradual phasing in of oil revenue into the government budget. The objective was to stabilize future fiscal expenditures pressure, the prevailing exchange rate and exchange rate expectations. It was thus noted by the OECD (2007) that monetary policy in Norway has the tasks of stabilizing economic growth and anchoring price inflation, while fiscal policy takes care of the real exchange rate.

3. REVIEW OF ECONOMETRIC METHODS

This section briefly reviews likelihood-based econometric methods using a CVAR model, paying particular attention to a procedure for model reduction. Economic time series data tend to exhibit non-stationary trending behavior, and so they are often regarded as processes integrated of order 1, denoted as I(1) hereafter. In order to analyze time series data, likelihood-based CVAR models are introduced and developed by Johansen (1988, 1996), and they have since played critical roles in time series econometrics. See Juselius (2006), Choo and Kurita (2011, 2012), *inter alia*, for various empirical illustrations utilizing CVAR analysis. Let us note that the primary reference for this section is Johansen (1996).

First, we present a fully unrestricted VAR(k) model for a p-dimensional time series $X_{-k+1}, \cdots, X_T$. The VAR model is given as

$$\Delta X_t = (\Pi, \Pi_l)\begin{pmatrix} X_{t-1} \\ t \end{pmatrix} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \Phi D_t + \varepsilon_t, \ \textit{for} \ \ t = 1, \cdots, T \tag{1}$$

where D_t denotes an s-dimensional vector of deterministic terms apart from intercept and trend, such as impulse and seasonal dummy variables, and the innovations $\varepsilon_1, \ldots, \varepsilon_T$ have independent and identical normal $N(0, \Sigma)$ distributions conditional on the starting values $X_{-k+1}, \cdots, X_0$. The parameters in equation (1) vary freely, defined as $\Pi, \Gamma_i, \Sigma \in R^{p \times p}$ and $\Pi_l, \mu \in R^p$, with Σ being a positive definite matrix.

In order to carry out the likelihood-based CVAR analysis of $I(1)$ data using equation (1), it is necessary that three regularity conditions be fulfilled. The first condition is that the characteristic roots obey the equation $|B(z)| = 0$, where

$$B(z) = (1-z)I_p - \Pi z + \sum_{i=1}^{k-1} \Gamma_i (1-z) z^i,$$

and the roots satisfy $|z| > 1$ or $z = 1$. This condition ensures that the process is free from both explosive nature and seasonal cointegration. The second condition, which is often called a reduced rank condition, is given by

$$rank(\Pi, \Pi_l) \leq r \quad \text{or } (\Pi, \Pi_l) = \alpha(\beta', \gamma),$$

where $\alpha, \beta \in \mathbf{R}^{p\times r}$ are of full column rank for $r < p$ and $\gamma \in \mathbf{R}^{r}$; the notational conventions $\beta^{*\prime} = (\beta', \gamma)$ and $X^*_{t-1} = (X'_{t-1}, t)'$ are introduced for future reference. In the reduced rank condition above α, is referred to as adjustment vectors, β^* as cointegrating vectors, while r designates cointegrating rank. The implication of this condition is that at least $p - r$ common stochastic trends exist and cointegration arises when $r \geq 1$. Finally, the third condition is

$$rank(\alpha'_{\perp}\Gamma\beta_{\perp}) = p - r,$$

where $\Gamma = I_p - \Sigma_{i=1}^{k-1}\Gamma_i$, and $\alpha_{\perp}, \beta_{\perp} \in \mathbf{R}^{p\times(p-r)}$ are orthogonal complements such that $\alpha'\alpha_{\perp} = 0$ and $\beta'\beta_{\perp} = 0$ with $(\alpha, \alpha_{\perp})$ and $(\beta, \beta_{\perp})$ being of full rank, respectively. This condition prevents the process from being $I(2)$ or of higher order. Under the satisfaction of these three conditions given above, we find that the $I(1)$ CVAR model is defined as a sub-model of equation (1) as follows:

$$\Delta X_t = \alpha\beta^{*\prime}X^*_{t-1} + \sum_{i=1}^{k-1}\Gamma_i \Delta X_{t-i} + \mu + \Phi D_t + \varepsilon_t, \tag{2}$$

which leads us to the subsequent cointegration analysis and model reduction.

Let us note that the cointegrating rank r is unknown in practice, so that it needs to be estimated from the data using reduced rank regression. A log-likelihood ratio (log LR) test statistic, formulated by the null hypothesis of r rank or $H(r)$ against the alternative hypothesis $H(p)$, is designated as log $LR[H(r) \mid H(p)]$, and this test statistic is relied upon to determine r in this empirical study. The limiting quantiles of the log LR test statistic are provided by Johansen (1996, Ch. 15). For gamma- approximation methods to obtain the quantiles, see Nielsen (1997) and Doornik (1998). Selecting the cointegrating rank in the $I(1)$ CVAR model, one can then proceed to the inspection of various restrictions on maximum likelihood estimates for α and β^*. Cointegrating combinations, $\beta^{*\prime}X^*_{t-1}$, correspond to a set of stationary linear combinations of non-stationary variables and act as equilibrium correction mechanisms in equation (2). The combinations are, in general, conceived of as empirical representations of long-run economic relationships among the non-stationary variables. One thus finds it important, in empirical studies, to explore a set of interpretable restrictions that can be placed in the estimation of β^*.

Next, a brief review is made regarding a partial CVAR model and model reduction procedure. Let us break down equation (2) into $X_t = (Y'_t, Z'_t)'$ for $Y_t \in R^m$ and , and $Z_t \in R^{p-m}$, and $r \leq m < p$. The parameters and innovations are also expressible in a conformable fashion as follows:

$$\alpha = \begin{pmatrix} \alpha_y \\ \alpha_z \end{pmatrix}, \quad \Gamma_i = \begin{pmatrix} \Gamma_{y,i} \\ \Gamma_{z,i} \end{pmatrix}, \quad \mu = \begin{pmatrix} \mu_y \\ \mu_z \end{pmatrix}, \quad \Phi = \begin{pmatrix} \Phi_y \\ \Phi_z \end{pmatrix}, \ \varepsilon_t = \begin{pmatrix} \varepsilon_{y,t} \\ \varepsilon_{z,t} \end{pmatrix},$$

$$\Sigma = \begin{pmatrix} \Sigma_{yy} & \Sigma_{yz} \\ \Sigma_{zy} & \Sigma_{zz} \end{pmatrix},$$

and let us define $\omega = \Sigma_{yz}\Sigma_{zz}^{-1}$. Suppose that a parametric condition $\alpha_z = 0$ is fulfilled. Equation (2) is then expressed as the combination of a partial CVAR model for Y_t conditional on Z_t and a marginal system for Z_t as follows:

$$\Delta Y_t = \omega \Delta Z_t + \alpha_y \beta^{*'} X_{t-1}^* + \sum_{i=1}^{k-1} \Gamma_{y,i}^* \Delta X_{t-i} + \mu_y^* + \Phi_y^* D_t + \varepsilon_{y,t}^*, \tag{3}$$

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_{z,i} \Delta X_{t-i} + \mu_z + \Phi_z D_t + \varepsilon_{z,t}, \tag{4}$$

where $\Gamma_{y,i}^* = \Gamma_{y,i} - \omega\Gamma_{z,i}$, $\mu_y^* = \mu_y - \omega\mu_z$, $\Phi_y^* = \Phi_y - \omega\Phi_z$ and $\varepsilon_{y,t}^* = \varepsilon_{y,t} - \omega\varepsilon_{z,t}$. If the condition $\alpha_z = 0$ is satisfied, Z_t is then weakly exogenous with respect to the parameters appearing in equation (3); this implies that, without estimating the marginal model or equation (4), the parameters of the partial model or equation (3) can be estimated with no loss of information. The parametric condition $\alpha_z = 0$ enables us to concentrate on the analysis of equation (3) alone, as long as all the parameters of interest are nested in the parameters of equation (3). For details of weak exogeneity in CVAR models, see Johansen (1992) and Urbain (1992).

One sees, therefore, that it is important to investigate whether $\alpha_z = 0$ is valid or not in empirical CVAR modeling. If this condition is empirically fulfilled, one can then move from the joint model to the partial model, which may be further reduced to a parsimonious VECM by employing a general-to-specific modeling approach. For details regarding this approach, see Hendry (1995), *inter alia*. The VECM can correspond to the underlying data generating mechanism and thus be utilized for the purposes of economic policy analysis and forecasting.

4. MODEL FOR MONEY DEMAND AND MONETARY POLICY RULE

This section, following Choo and Kurita (2011, 2012) and referring to the existing research in monetary economics, introduces a theoretical framework for money demand and monetary policy rule functions. This framework lays the foundations for long-run econometric analyses carried out in the following sections.

4.1. Money Market Equilibrium

With price homogeneity imposed, demand for real broad money may be given as follows:

$$m_t - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 i_t^s - \gamma_3 i_t^l - \gamma_4 \pi_t, \tag{5}$$

where m_t, p_t, and y_t are the logarithms of real money balances, price level and real output, respectively; i_t^s is a short-term interest rate, which measures the own rate of money; i_t^l is a long-term interest rate such as yields on government bonds; π_t is inflation rate, which is the opportunity cost of money in terms of physical assets; and, $\gamma_i > 0$ for $i = 1, \cdots, 4$. Thus, Equation (5) is perceived as a specification of the underlying equilibrium condition for the money market.

We may take a unit value for the income elasticity, that is, $\gamma_1 = 1$. It is also useful to test if the two interest rates, i_t^s and i_t^l, hold equal coefficients with opposite signs. Equation (5) is thus rearranged as follows:

$$m_t - p_t = \gamma_0 + y_t + \gamma_2 (i_t^s - i_t^l) - \gamma_4 \pi_t.$$

Furthermore, we are interested in testing $\gamma_2 = \gamma_4$ so that we can arrive at

$$m_t - p_t = \gamma_0 + y_t - \gamma_2 (i_t^l - i_t^s + \pi_t). \tag{6}$$

Or, if an empirical study indicates that the interest rates do not play any significant role, it follows that $\gamma_2 = 0$ holds, leading to

$$m_t - p_t = \gamma_0 + y_t - \gamma_4 \pi_t.$$

Therefore,

$$-(m_t - p_t - y_t) = -\gamma_0 + \gamma_4 \pi_t. \tag{7}$$

In short, the equilibrium condition (5) can be modified to express a positive relationship between the velocity of money and inflation as well as a negative relationship between money and inflation. Empirical evidence suggests the velocity of money and inflation tend to be pro-cyclical if other variables are treated as given (see, for instance, Gylfason and Herbertsson, 2001). A negative relationship between money and inflation is examined in an empirical study of Norway by Eitrheim (2003). In the empirical study conducted in the next section, the money market equation, (6) or (7), is treated as a candidate relationship that may coincide with one of the long-run cointegrating linkages underlying the data.

4.2. Monetary Policy Rule based on Interest Rates

Let us move on to monetary policy. Taylor (1993) suggests a simple policy rule in which the short-term policy rate is responsive to deviations of output and inflation from their respective policy targets:

$$i_t^s = \pi_t + \eta_0 + \eta_1 (y_t - y_t^*) + \eta_2 (\pi_t - \pi_t^*), \tag{8}$$

where y_t^* represents the natural logarithm of potential real output; π_t and π_t^* are actual inflation and target inflation, respectively; and η_0 denotes a target real short-term interest rate, and $\eta_i > 0$ for $i = 0, 1, 2$. In an open economy like Norway, the exchange rate could play a role in the monetary policy. Svensson (2000) and Taylor (2001) consider the existence of the real exchange rate in the implementation of monetary policy. However, a complicated time lag structure of the exchange rate's influence on inflation, among others, may be difficult to model in the cointegration framework, and thus the exchange rate is not explicitly considered in this paper. This view on the exchange rate is partially justified in the empirical analysis performed at the end of Section 5.3.

Equation (8) may be regarded as a policy rule to pursue the objective of inflation targeting. In other words, the Taylor rule can be viewed as an operating procedure for the monetary authority to achieve its medium-term inflation target. The rule contains the output gap or $y_t - y_t^*$, which is viewed as a measure of inflation pressure underlying the macro economy in real terms. If we could assume that the potential output shows a log linear deterministic trend, then we find $y_t^* = \theta_1 + \theta_2 t$ with $\theta_i > 0$ for $i = 1, 2$. Furthermore, supposing that the inflation target could be assumed to be a fixed target in such a way as $\pi_t^* = \theta_3$ for $\theta_3 > 0$, then equation (8) leads to

$$i_t^s = \nu_0 + \nu_1 y_t + (1 + \nu_2)\pi_t + \nu_3 t, \tag{9}$$

where $\nu_i = \eta_i$ for $i = 1,2$, $\nu_0 = \eta_0 - \nu_1\theta_1 - \nu_2\theta_3$ and $\nu_3 = -\nu_1\theta_2$. Laurent (1988) as well as Bernanke and Blinder (1992) argue that the spread between the Federal funds rate and a long-term government bond rate is an empirical indicator of monetary policy stances; their argument is basically backed up by the credit view of monetary policy. Furthermore, Bernanke and Blinder (1992) point out that the Federal funds rate, not the long-term bond rate, dominates the behavior of the interest rate spread over time. The same feature also seems to be found in the corresponding interest rate spread in Norway; the correlation coefficient between $(i_t^s - i_t^l)$ and i_t^s is 0.309, while that between $(i_t^s - i_t^l)$ and i_t^l is just -0.097, as shown in Appendix B. Mehra (2001) notices the long-term inflationary expectations as reflected in the long-term bond rate in the monetary policy of the Fed. Christensen and Nielsen (2009) show that an adjusted Taylor-type policy rule, in which the bond rate is used instead of inflation, is judged to be congruent with the data in the U.S. Let us suppose that the long-term bond rate moves in line with inflation and also note that the rate of inflation already exists in equation (9). It is then justifiable to add to equation (9) a real long-term rate, $i_t^l - \pi_t$, which yields

$$\begin{aligned} i_t^s &= \nu_0 + \nu_1 y_t + (1 + \nu_2)\pi_t + \nu_3 t + \nu_4\left[\left(i_t^l - \pi_t\right) - \bar{r}^l\right] \\ &= \nu_0^* + \nu_1 y_t + (1 + \nu_2 - \nu_4)\pi_t + \nu_4 i_t^l + \nu_3 t, \end{aligned} \tag{10}$$

where $\bar{r}^l$ denotes the mean of the real long-term rate, which is assumed to be approximately time-invariant, and $\nu_0^* = \nu_0 - \nu_4 \bar{r}^l$. Suppose that $v_4 = 1$ holds, that is, there is a one-to-one relationship between the long-term rate and the short-term rate. Equation (10) then leads to

$$i_t^s = \nu_0^* + \nu_1 y_t + \nu_2 \pi_t + i_t^l + \nu_3 t.$$

Due to $v_3 = -v_1\theta_2$, this is expressed as

$$i_t^s = \nu_0^* + \nu_1(y_t - \theta_2 t) + \nu_2 \pi_t + i_t^l, \tag{11}$$

Note that the last terms $v_2\pi_t + i_t^l$ indicate influences stemming from both inflation and expected inflation (hidden in the long-term bond rate) on the short-term interest rate. Rearranging the equation above, the following is obtained:

$$i_t^s - i_t^l = v_0^* + v_1(y_t - \theta_2 t) + v_2\pi_t. \tag{12}$$

Thus, equation (12) shows a monetary policy target expressed in terms of the interest rate spread, $i_t^s - i_t^l$. In empirical analysis performed in the following section, we regard the specification (12) as a possible long-run economic relationship that may be revealed from the Norway data.

5. MULTIVARIATE COINTEGRATION ANALYSIS

This section conducts a formal CVAR analysis of quarterly time series data from Norway, with a view to casting light on the monetary aspect of its inflation dynamics. The canonical model presented in the previous section allows us to introduce the following variables to be studied:

$$X_t = (\pi_t \quad i_t^s - i_t^l \quad m_t - p_t \quad y_t)',$$

which results in a four-dimensional unrestricted VAR model as in Choo and Kurita (2012); see equation (1) as well as Appendices A and B for details of the data and their overview, respectively. Unable to find any significant results with the inflation as measured by the CPI, we thus use the inflation as measured by the GDP deflator. For the short-term policy rate, the discount rate at Norges Bank is used rather than the sight deposit rate because the latter is not always available for the whole sample period. Let us point out that the measure of real broad money, $m_t - p_t$, is contained in the variable set above. This is because our goal is to achieve an econometric model that casts light on the underlying dynamic linkage between monetary aggregate and inflation. This section is composed of three sub-sections. Section 5.1 obtains an unrestricted VAR system and determines its cointegrating rank. Using the estimated VAR model, Section 5.2 examines the weak exogeneity and Section 5.3 identifies the adjustment and the long-run economic structure.

5.1. Choosing the Cointegrating Rank

This sub-section estimates an unrestricted VAR model to examine its cointegrating rank. The sample period available for estimation ranges from the first quarter in 1987 to the fourth quarter in 2011, henceforth denoted by 1987.1 - 2011.4. The number of observations amounts to 100. The initial estimation of the model is carried out by choosing the lag length of the model for 3. Due to observed seasonality, a set of centered seasonal dummy variables is included in the model. It is then found that the lagged regressors at length 3 are insignificant and can be removed from the model. The selection of lag length 2 or $k = 2$ is thus appropriate from the viewpoint of statistical significance, although some evidence is found indicating residual autocorrelation in the equation for π_t. It is hence judged that the model's short-run dynamics should be adjusted in such a manner as proposed by Kurita and Nielsen (2009) with the selection of $k = 2$ for the model's lag length. Namely, a (three-quarter) lagged second-order difference term for inflation, $\Delta^2\pi_{t-3}$, is added to the VAR(2) model unrestrictedly, resulting in the autocorrelation problem being addressed while the standard limit theory for the cointegrating rank test remains valid. Furthermore, the data seem to be much influenced by outliers in 1998.3, 2008.4 and 2009.1. As described in Section 2, the first outlier is attributable to a policy response to the Asian currency crisis occurring in 1998, whereas the second and third outliers are caused by global economic depression triggered by the US financial crisis starting in 2008. Therefore, the following three dummy variables, holding either 0 or 1, are included unrestrictedly in the model: $D_{1,t} = 1(1998.3)$, $D_{2,t} = 1(2008.4)$, $D_{3,t} = 1(2009.1)$ and 0 otherwise.

Table 1 documents a battery of diagnostic tests for the residuals of the VAR model. Most of the test results are presented in the form F_j $(k, T - l)$, which designates an approximate F test against the alternative hypothesis j: kth-order autocorrelation (F_{ar}: see Godfrey, 1978, Nielsen, 2006), kth-order autoregressive conditional heteroskedasticity or ARCH (F_{arch}: see Engle, 1982), and heteroskedasticity (F_{het}: see White, 1980). The table also provides a chi-square test for normality (χ^2_{nd} : see Doornik and Hansen, 2008). These mis-specification test statistics are all insignificant at the 5% level and are thus in favor of the view that the VAR model represents the data well. It is, therefore, justifiable for us to employ this VAR model for the purpose of conducting subsequent cointegration analysis and model reduction.

Table 1
Diagnostic tests

	πt	$i_t^s - i_t^l$	$m_t - p_t$	y_t
Autocorr. [$F_{ar}(5,78)$]	0.56[0.73]	0.86[0.51]	0.81[0.55]	1.63[0.16]
ARCH [$F_{arch}(4,75)$]	0.48[0.75]	1.90[0.12]	1.91[0.12]	1.06[0.38]
Hetero.[$F_{het}(18,64)$]	0.91[0.57]	0.90[0.58]	1.45[0.14]	0.60[0.88]
Normality [χ^2_{nd} (2)]	1.70[0.43]	3.17[0.21]	0.01[0.99]	4.94[0.08]

Note: Figures in the square brackets are p-values.

The issue of selecting appropriate cointegrating rank is addressed using the estimated VAR model above. The first panel of Table 2 presents a set of log *LR* test statistics for the choice of cointegrating rank. The null hypotheses of $r = 0$ and $r \leq 1$ are both rejected at the 5% significance level, whereas the hypothesis of $r \leq 2$ is not rejected even at the 10% level. In order to verify the selection of $r = 2$, the second panel of Table 2 reports the modulus (denoted by *mod*) of the six largest eigenvalues obtained from a companion matrix for the CVAR model, which is estimated with the restriction of $r = 2$. All the eigenvalues, except the imposed two unit roots, are much smaller than 1, suggesting that there is neither explosive nor $I(2)$ root involved in the model. These pieces of evidence allow us to reach the conclusion that the cointegrating rank is set at 2, which leads to further analysis using the CVAR model with $r = 2$.

Table 2
Cointegrating rank tests

	$r = 0$		$r \leq 1$	$r \leq 2$		$r \leq 3$
log LR{H(r) | H(p)}	78.81[0.00]**		44.65[0.03]*	17.93[0.36]		1.10[0.99]
mod ($r = 2$)	1	1	0.66	0.55	0.55	0.51

Notes: Figures in the square brackets are p-values.
** and * denote significance at the 1% and 5% levels, respectively.

5.2. Testing Weak Exogeneity

The selection of the cointegrating rank enables us to carry out hypothesis testing for the estimates of α and β^*, relying on χ^2-based asymptotic inferences. As reviewed in Section 3, we find it possible to explore weak exogeneity by testing for zero restrictions on all adjustment coefficients for the variable in question. If two variables, $m_t - p_t$ and y_t, are judged to be weakly exogenous with respect to the parameters of a partial system for π_t and $i_t^s - i_t^l$, it is then justifiable to focus on the analysis of this partial system alone for the purpose of making statistical inferences with no loss of information. The modeling approach relying upon a partial system, which is consistent with the canonical model developed in Section 4, helps us to lessen the modeling efforts necessitated in the econometric exploration.

Table 3
Weak exogeneity tests

πt	$i_t^s - i_t^l$	$m_t - p_t$	y_t
6.50 [0.04]*	8.80[0.01]*	1.25[0.53]	0.42[0.81]

Notes: Figures in the square brackets are p-values according to $\chi^2(2)$.
* denotes significance at the 5% level.

Table 3 documents a battery of log *LR* test statistics for weak exogeneity. As shown in the table, the null hypothesis of weak exogeneity is not rejected for $m_t - p_t$ and y_t, with both of their p-values being much greater than the level of 0.05. Judging from these test results, one is able to conclude that both $m_t - p_t$ and y_t are conceived of as weakly exogenous for the parameters of the partial model for π_t and $i_t^s - i_t^l$. The analysis of the bivariate partial model, instead of that of the full model, is therefore justified, paving the way for further model reduction in pursuit of a parsimonious expression of the data. We address the issue of identifying equilibrium correction mechanisms in the next sub-section to give a basis for the exploration of such a parsimonious model.

Furthermore, according to the results of Table 3, one can point out that $\alpha_\perp$ is revealed to be

$$\alpha_\perp' = \begin{pmatrix} 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}.$$

It is known that $\alpha_\perp' \Sigma_{i=1}^t \varepsilon_i$ acts as the common stochastic trends for the CVAR model (see Johansen, 1996, p. 41); the trends give rise to the non-stationary behavior of X_t. One may thus view $\alpha_\perp{}' X_t$ as a set of linear combinations of those variables whose innovations amount to the overall driving force embedded in the CVAR system. See also Burke and Hunter (2011) for economic interpretations of the common trends in the context of long-run price targeting. In our case, $\alpha_\perp' X_t$ is given as

$$\alpha_\perp' X_t = (m_t - p_t \quad y_t)'.$$

It is, therefore, justifiable to conceive that two real variables, $m_t - p_t$ and y_t, work as the overall pushing force; in contrast, two nominal variables, π_t and $i_t^s - i_t^l$, purely adjust to the cointegrating combinations, thereby acting as the overall pulling force. See Hoover *et al*. (2008) for such interpretations as pushing and pulling forces in CVAR analysis. For this reason, one may argue that the set of real shocks drives the CVAR system in the long-run, while the nominal variables react to long-run disequilibrium errors, pulling back the system on the right track.

5.3. Identifying the Structure of Equilibrium Correction

This sub-section explores equilibrium correction mechanisms embedded in the time series data of Norway. The empirical results obtained so far agree with Choo and Kurita (2012) and thus encourage following the same route in order to reveal the underlying long-run structure. Under the zero restrictions imposed on the adjustment vectors for $m_{t-1} - p_{t-1}$ and y_{t-1} in accordance with the results in Table 3, the two cointegrating vectors are normalized with respect to π_{t-1} and $i_{t-1}^s - i_{t-1}^l$, respectively. A series of econometric investigations results in revealing cointegrating and adjustment vectors that are subject to some meaningful economic interpretations. Let us recall that the definition of X_{t-1}^* is given as

$$X_{t-1}^{*} = \left(\pi_{t-1}, i_{t-1}^{s} - i_{t-1}^{l}, m_{t-1} - p_{t-1}, y_{t-1}, t\right)',$$

and reduced rank regression reveals the following structure:

$$\hat{\alpha}\,\hat{\beta}^{*\prime} = \begin{bmatrix} \underset{(0.048)}{-0.269} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(0.042)}{-0.207} \\ \underset{(-)}{0} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(-)}{0} \end{bmatrix} \begin{bmatrix} \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.117)}{0.291} & \underset{(-)}{-0.291} & \underset{(0.00028)}{-0.000631} \\ \underset{(-)}{0} & \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.050)}{-0.134} & \underset{(0.00035)}{0.000827} \end{bmatrix}', \tag{13}$$

where the figure in the parenthesis under each coefficient is the standard error. The log *LR* test statistic for the restrictions in equation (13) is 5.21, and its *p*-value according to $\chi^2(8)$ is 0.73. Thus, the null hypothesis of the overall restrictions on the cointegrating and adjustment vectors is not rejected even at the 10% level.

Let us interpret the restricted cointegrating combinations in equation (13) as long-run equilibrium relationships backed up by economic theory. For this purpose, let $v_{i,t-1}$ for $i = 1, 2$ denote a stationary error term, which corresponds to a deviation from an estimated long-run equilibrium relationship. The first cointegrating combination, as reported in equation (13), is then expressible as

$$\pi_{t-1} = 0.291(p_{t-1} - m_{t-1} + y_{t-1}) + 0.000631t + v_{1,t-1}, \tag{14}$$

which can be seen as an empirical representation of equation (7) in the theoretical model in Section 4. The equilibrium relationship (14), coupled with its adjustment structure in equation (13), may be viewed as evidence indicating that disequilibrium in the money market gives rise to inflationary pressures in the overall economy. From equation (14), one can infer that the level of equilibrium inflation is determined by the velocity $p_{t-1} - m_{t-1} + y_{t-1}$ as well as the linear trend t, which may approximate the underlying trending behavior of the velocity itself. Let us also point out that these empirical findings match those of the first cointegrating relationship discussed in Choo and Kurita (2012) with respect to the New Zealand economy. Overall, it seems that the monetary aspect of long-run inflation dynamics is well represented by the equilibrium relationship (14) as well as its adjustment structure.

Let us move on to consider the economic interpretation of the second cointegrating combination in equation (13). This is expressed as

$$i_{t-1}^{s} = i_{t-1}^{l} + 0.134(y_{t-1} - 0.00617t) + v_{2,t-1}, \tag{15}$$

where the term in the bracket on the right hand side is viewed as an empirical measure of the output gap. Equation (15) is consistent with equation (12) obtained from a conceivable Taylor-type monetary policy rule in Section 4. Again, let us note that (15) matches the second cointegrating relationship in Choo and Kurita (2012) in terms of its long-run structure and adjustment mechanism. This equilibrium relationship suggests that the output gap plays a critical role in the underlying monetary policy rule. Macroeconomic literature, in general, seems to support this view. Svensson (1997), for instance, shows that the speed of adjustment in inflation forecasts towards its target level depends upon the weight attached to output stabilization. It is also demonstrated, by Fazzari *et al.* (2010), that a policy response to output is more effective in stabilizing business cycles and fluctuations than that to inflation. In addition, notable is the possibility that information on *expected* inflation can be contained in the long-term interest rate in equation (15), as pointed out in Section 4, although the rate of actual inflation itself is judged insignificant and therefore removed from the equation.

The identified equilibrium relationships, (14) and (15), act as a set of important factors for the dynamics of a parsimonious VECM pursued in the next section. One may also argue that the pushing force revealed in the previous sub-section plays critical roles in these equilibrium combinations. Moreover, it should be again stressed that the overall structure revealed here is fairly similar to that found in Choo and Kurita (2012) for New Zealand, the first economy that officially adopted a policy rule for inflation targeting. Thus, the revealed structure may be seen as a set of common features of economies that obey the principle of monetary policy rules, although further empirical explorations of various countries and regions are required to reach a decisive conclusion on this hypothetical view.

Before moving on to the estimation of a parsimonious VECM, one may find it informative to check possible long-run influences of an exchange rate variable on the CVAR system above. For this purpose, the log of the Norwegian krone - US dollar exchange rate, denoted as e_t, is newly introduced in the CVAR model, which leads to the investigation of long-run structure in a manner similar to that employed in equation (13). That is, X_t is temporarily re-defined as

$$X_t = \left(\pi_t, i_t^s - i_t^l, m_t - p_t, y_t, e_t\right)',$$

and reduced rank regression is then performed to see whether or not e_t can be eliminated from both of the cointegrating relationships. For this extended system the following structure is revealed:

$$\hat{\alpha}\,\widehat{\beta}^{*\prime} = \begin{bmatrix} \underset{(0.05)}{-0.261} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(0.042)}{-0.203} \\ \underset{(-)}{0} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(-)}{0} \\ \underset{(0.114)}{0.357} & \underset{(-)}{0} \end{bmatrix} \begin{bmatrix} \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.114)}{0.319} & \underset{(-)}{-0.319} & \underset{(-)}{0} & \underset{(0.00027)}{-0.000679} \\ \underset{(-)}{0} & \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.051)}{-0.141} & \underset{(-)}{0} & \underset{(0.00036)}{0.000886} \end{bmatrix},$$

in which the log *LR* test statistic for all the restrictions is 13.8 with its *p*-value according to $\chi^2(11)$ being 0.24. Note that a set of zero restrictions can be imposed on the fifth column of $\hat{\beta}^{*\prime}$, thus e_t can be excluded from the cointegrating space, allowing us to infer that the krone-dollar exchange rate may play a limited role in the long-run structure of the system. It is true, according to the estimates for adjustment parameters above, that e_t is not weakly exogenous for cointegrating parameters in this extended CVAR system. The finding that e_t is long-run excluded, however, may justify focusing on a further investigation of the preceding small CVAR system without reference to e_t. A further investigation of the roles of various exchange rate variables is beyond the scope of this paper, but we note that such investigation should be viewed as one of the most important objectives in future research extensions. The next section explores a reduced VEC representation incorporating equation (13) by concentrating on the small CVAR system.

6. A REDUCED VECTOR EQUILIBRIUM CORRECTION MODEL

This section seeks a parsimonious VECM for $\Delta\pi_t$ and $\Delta(i_t^s - i_t^l)$, which are viewed as the overall pulling force for the empirical CVAR system explored thus far. Mapping the data to the $I(0)$ space by using the restricted cointegrating relationships (denoted by $ecm_{1,t-1}$ and $ecm_{2,t-1}$, respectively) and differencing all the variables, we build a bivariate VECM given the two weakly exogenous variables. With a view to inspecting net monetary influences on inflation acceleration, or $\Delta\pi_t$, we adopt the growth of *nominal* money, Δm_t and Δm_{t-1}, as money-related regressors in the equation for $\Delta\pi_t$, instead of the growth of real money. Next, insignificant contemporaneous and lagged regressors are removed from the VECM step by step. Lastly, as a result of model reduction, we reach the following parsimonious VECM:

$$\begin{aligned}
\Delta\pi_t = &-\underset{(0.062)}{0.335}ecm_{1,t-1} + \underset{(0.111)}{0.240}\Delta m_{t-1} + \underset{(0.081)}{0.327}\Delta\pi_{t-1} + \underset{(0.059)}{0.323}\Delta^2\pi_{t-3} + \underset{(0.015)}{0.074} \\
&- \underset{(0.025)}{0.098}D_{3,t} + \hat{\varepsilon}_{1,t}, \\
\Delta(i_t^s - i_t^l) = &-\underset{(0.042)}{0.188}ecm_{2,t-1} + \underset{(0.074)}{0.320}\Delta(i_{t-1}^s - i_{t-1}^l) - \underset{(0.014)}{0.042}\Delta(m_{t-1} - p_{t-1}) \\
&-\underset{(0.006)}{0.025} + \underset{(0.004)}{0.028}D_{1,t} - \underset{(0.004)}{0.016}D_{2,t} + \hat{\varepsilon}_{2,t},
\end{aligned} \tag{16}$$

where

$$ecm_{1,t-1} = \pi_{t-1} + 0.291(m_{t-1} - p_{t-1} - y_{t-1}) - 0.000631t,$$
$$ecm_{2,t-1} = i_{t-1}^s - i_{t-1}^l - 0.134y_{t-1} + 0.000827t,$$

Autocorr. $[F_{ar}(20,166)] = 0.74\ [0.78]$,

Hetero. $[F_{het}(69,204)] = 0.92\ [0.65]$,

Normality$[\chi^2_{nd}(4)] = 4.23\ [0.38]$.

Note that the figure in the parenthesis under each coefficient represents a standard error. A battery of system-based diagnostic tests is presented below equation (16), coupled with their *p*-values. It is evident, as seen from the *p*-values, that the null hypotheses are not rejected at the conventional significance level. The fitted values of the VECM, together with corresponding actual values, are plotted in Figs. 1 (a) and (c). The data tracking looks satisfactory. Moreover, the quantile-quantile (QQ) plots of scaled residuals based on standard normal distribution are displayed in Figs. 1 (b) and (d); in line with the system-based diagnostic tests above, no strong evidence is found for the rejection of the normality assumption with regard to the residuals. In addition, Figs. 2 (a) and (c) report scaled residuals, while Figs. 2 (b) and (d) present recursive break-point Chow tests (see Chow, 1960). Again, it seems there is no strong evidence suggesting any model mis-specification issues. We can therefore arrive at the conclusion that the parsimonious VECM (16) is a data-congruent representation.

Let us proceed to take a close look at the VECM coefficients above. As expected, the two ECM terms, $ecm_{1,t-1}$ and $ecm_{2,t-1}$, act as highly significant factors in the equations for $\Delta\pi_t$ and $\Delta\left(i_t^s - i_t^l\right)$, respectively. These statistical findings ensure the existence of strong equilibrium correction mechanisms in the ECM. Furthermore, it is noteworthy that a lagged difference of nominal money, Δm_{t-1}, plays a significant role in the equation for inflation acceleration, coupled with $ecm_{1,t-1}$, *i.e.* equilibrium correction working in

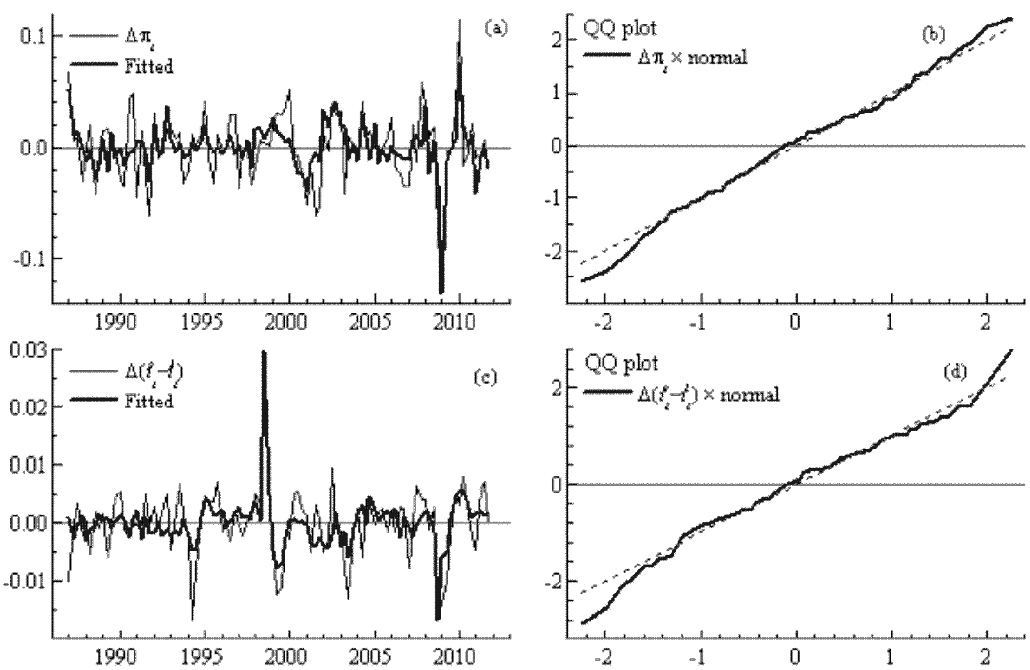

Figure 1: Fitted values and residual QQ plots

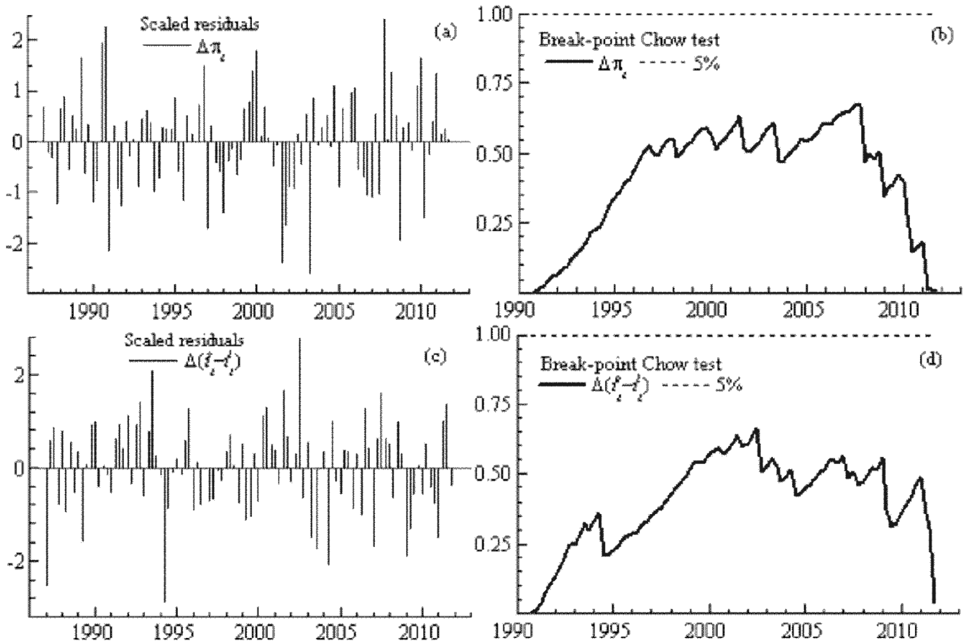

Figure 2: Scaled residuals and recursive Chow tests

the money market. In order to double-check its significance, a log *LR* test for the exclusion of Δm_{t-1} from this equation is carried out; the null hypothesis of its exclusion is indeed rejected at the 5% level with its test statistic being 4.68 [0.03]*, where the figure in the bracket is a *p*-value according to $\chi^2(1)$. According to the inflation acceleration equation, the coefficient for Δm_{t-1} is positive, indicating that nominal money growth contributes to the acceleration of inflation. We can thus infer from these outcomes that aggregate money in both level and difference contains useful information about the dynamics of inflation acceleration. Furthermore, a lagged real money difference, $\Delta(m_{t-1} - p_{t-1})$, is judged to be a highly significant factor in the $\Delta\left(i_t^s - i_t^l\right)$ equation, accompanied with a negative coefficient. It seems plausible that a monetary expansion in real terms brings about a short-run downward pressure on the short-term interest rate relative to the long-term rate. With the passage of time, such a downward impact may be absorbed in the long-term rate by way of expected decreases in the short–term rate, as indicated in the expectation theory of the term structure. Overall, we judge that several important roles played by aggregate money in the economy are manifested in the VECM above.

In order to demonstrate the practical usefulness of the VECM, equation (16) is re-estimated using a truncated sample period, 1986.2 - 2009.2, so that the sequences of 1-step forecasts of $\Delta\pi_t$, $\Delta\left(i_t^s - i_t^l\right)$, $ecm_{1,t}$ and $ecm_{2,t}$ for 2009.3 - 2011.4 are derived from it.

Note that equation (16) is not significantly influenced by any contemporaneous regressors; therefore, the VECM can be used as a device generating these 1-step ahead forecast sequences. The sequences of forecasts, which are displayed in Fig. 3, track the actual data fairly well, indicating that the model can be relied upon as a macroeconomic forecasting device.

This section demonstrates that the CVAR model is successfully reduced to a parsimonious VECM accounting for the behavior of $\Delta\pi_t$ and $\Delta\left(i_t^s - i_t^l\right)$. The VECM is seen as a data-congruent representation and sheds light on macroeconomic dynamics in Norway. One can also view the VECM as a useful empirical reference for the purpose of grasping the monetary aspect of its inflation dynamics.

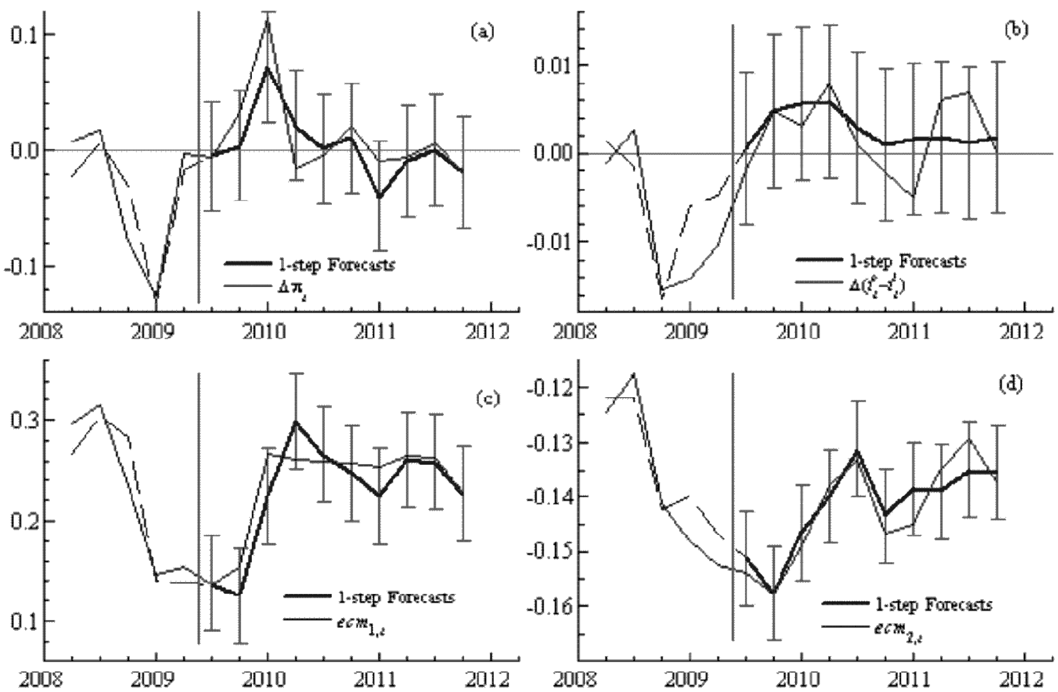

Figure 3: **1-step forecasts generated from the VECM**

7. CONCLUDING REMARKS

Utilizing a detailed multivariate time series analysis, this paper explores a monetary perspective on Norway's inflation dynamics over the period of 1987-2011. A set of Norway's quarterly data on its money market, price inflation and monetary policy are thoroughly analyzed for this empirical purpose. A CVAR analysis reveals two long-run economic relationships, which are seen as interpretable from the standpoints of money market equilibrium and an empirical monetary policy rule. It is also shown that disequilibrium in the money market brings about a long-run inflationary impetus in the economy. Finally, a CVAR system is reduced to a parsimonious VECM, which is viewed as a data-congruent monetary model of Norway's overall inflation dynamics. Also, the preferred model is demonstrated as reliable for the purpose of applied economic analysis such as forecasting. The overall empirical evidence supports the view that aggregate money contains quantitative information useful for explaining the dynamics of inflation data. One can also recognize VECM as a useful empirical reference with a view to grasping the monetary aspect of Norway's inflation dynamics.

Furthermore, it is argued that the distinguishing feature of the estimated model is fairly similar to that found in Choo and Kurita (2012) in application to the case of New Zealand, the first economy that has officially employed an inflation targeting policy rule since the early 1990s. We may, therefore, conjecture that the revealed structure can be perceived as a type of common features of economies that tend to rely on monetary policy rules rather than discretionary policies. As this hypothetical view is, of course, an open issue, further detailed analyses of various countries and regions are necessary to conclude that this view is indeed justified.

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Appendix A: data definitions, sources and notes

A.1. Data definitions

π_t = the percentage change in the GDP deflator index (2005 = 100) over the previous four quarters, *i.e.* $\Delta^4 p_t$,

$m_t - p_t$ = the log of the end-of-period broad money M2 – the log of the GDP deflator index,

y_t = the log of the real GDP,

$i_t^s - i_t^l$ = the discount rate at Norges Bank – the yield on five-year government bond, and

e_t = the log of the Norwegian krone-US dollar exchange rate (period average),

A.2. Source and notes

The data of π_t, $m_t - p_t$, $y_t, i_t^s - i_t^l$ and e_t are obtained from *International Financial Statistics* (International Monetary Fund). Each component in the interest rate spread is defined as $i_t^s = \log(1 + I_t^s/100)$ and $i_t^l = \log(1 + I_t^l/100)$, where I_t^s and I_t^l denote the corresponding original series (in percent) available in the data source.

Appendix B

Table A
Correlation matrix (1987Q1 - 2011Q4)

	$i_t^s - i_t^l$	i_t^s	i_t^l
$i_t^s - i_t^l$	1	0.309	-0.097
i_t^s	*	1	0.916
i_t^l	*	*	1

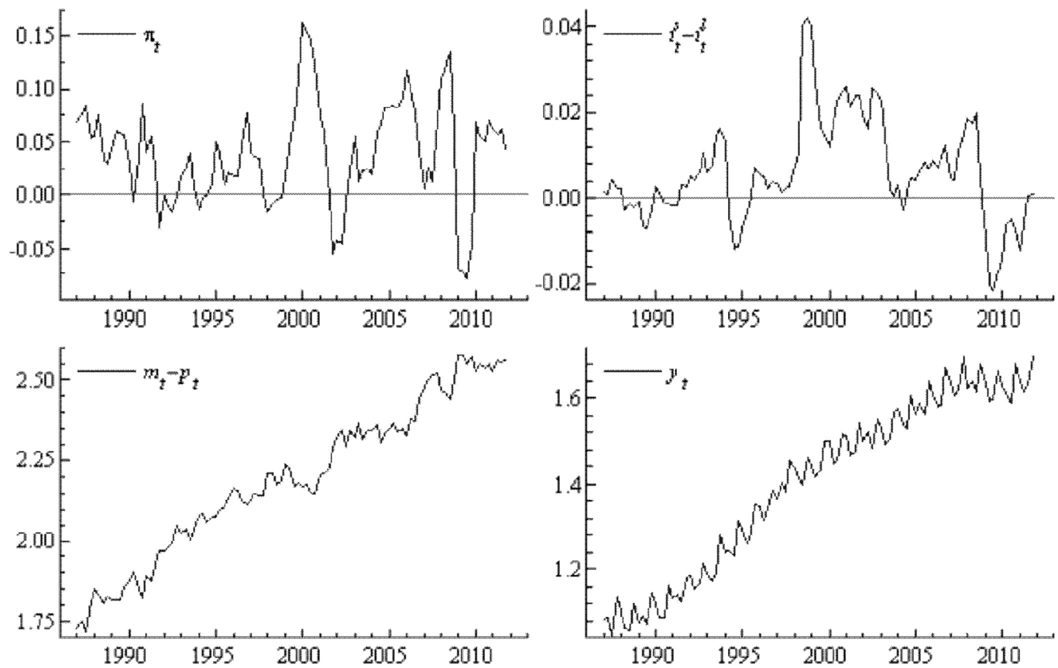

Figure A: Overview of the data

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