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A Monetary Perspective on Inflation Dynamics in Norway

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Abstract

The objective of this paper is to investigate a monetary perspective on Norway's inflation dynamics over the period of 1987- 2011 by relying on a detailed multivariate time series analysis. 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, Money market, 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

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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, p.13), 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.

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 Autkrust (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 utilized in such a way as the employment of the 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 Soderstrom (2001), and Christensen and Nielsen (2009), *inter alia*. Leitemo and Soderstrom (2001) is 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 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 Johansen and Juselius (1990), Hoffman and Rasche (1991), Hendry and Mizon (1993), and Bårdsen *et al*. (2005), *inter alia*, for various empirical illustrations using a CVAR analysis. 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 are often treated as separate research topics in macroeconomic literature, although both of these should be simultaneously related to various important elements of monetary economies under study. However, Brűggemann (2003) as well as Choo and Kurita (2011, 2012) may be considered as exceptional in the literature; Brűggemann (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's and New Zealand's data, respectively. These articles achieve meaningful empirical outcomes in the joint modeling of monetary policy rule and money demand functions. 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 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 large oil exporting sector. A value-added-tax system was introduced in 1970 and incomes policies were implemented in the late 1970s. Bank credit aggregates for industrial sectors played a key role as intermediate targets of monetary policy in the 1970s and the early 1980s. Nominal interest rates were usually kept lower than the market-clearing level, and raised when the monetary authorities needed restraining credit demand. Norway pursued extensive deregulation of financial markets from 1984 onwards. Direct and selective controls of credit volume were abolished, and then capital controls were completely removed in 1990. Interest rates were gradually determined by market forces. Financial deregulation led to a surge in bank lending, as well as in interest rates, fueling inflation. Then a stock market crash, a collapse of economic activity and failures of several major banks followed. A banking crisis developed after 1987 and peaked in 1991 and 1992. The banking crisis coincided with a severe

recession. See Bårdsen and Klovland (2000).

In 1986, targets for monetary policy changed from credit volume and interest rates to the exchange rate. Following a 12% devaluation of the 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 (as of January 1999, the euro). Norges Bank increased (decreased) the interest rate when there was depreciating (appreciating) pressure. 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 had a main responsibility for stabilizing the economy. Over the period of 1976-1986, ten devaluations were made, which led to inflationary expectation, actual inflation and instability. It is pointed by Bergo (2007) that the credibility of the fixed exchange rate regime was also impaired. During the disturbances in the European financial markets in late 1992, the krone was depreciated and allowed to float under a managed float, with a more-or-less free float following the speculative attacks in 1997 and 1998. Thus, there began a *de facto* shift toward a flexible exchange rate policy.

In March 2001, Norway abandoned a regime of exchange-rate targeting and instead adopted inflation targeting. The bank is explicit about being a *flexible* inflation targeter, so 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 changed 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. Since 2001 Statistics Norway has also prepared an index that excludes energy products and is adjusted for tax changes, the CPI-ATE. Norges Bank has developed a third measure of inflation, CPIXE, which is a combination of the CPI-ATE and an estimated trend growth in energy prices. Since 2005, Norges Bank has published its own interest rate forecast, along with forecasts for inflation, the output gap, and other key variables. It presents a probability distribution ("fan chart") around the point forecast. When inflation targeting was formally adopted, a fiscal rule was also created and made effective as of 2002, specifying a gradual phasing in of oil money into the fiscal budget. The objective was to stabilize future fiscal spending pressure and thereby stabilize real exchange rate expectations and the current exchange rate. It was thus noted by the OECD (2007) that the current macroeconomic stabilization framework in Norway assigns to monetary policy the task of anchoring inflation and stabilizing growth, while fiscal policy looks after 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 $I(1)$ 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), *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

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

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 $\varepsilon_1,\cdots,\varepsilon_T$ 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 $\Pi, \Gamma_i, \Sigma \in \boldsymbol{R}^{p\times p}$, $\Phi \in \boldsymbol{R}^{p\times s}$ and $\Pi_l, \mu \in \boldsymbol{R}^p$, 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 should 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 neither explosive nor seasonally cointegrated. The second condition, which is often called a reduced rank condition, is given by

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

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^* is known 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 $\Gamma = I_p - \sum_{i=1}^{k-1} \Gamma_i$, 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. This condition prevents the process from being $I(2)$ or of higher order. If these three conditions given above are fulfilled, 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 subsequent cointegration analysis and model reduction.

Let us note that the cointegrating rank r is unknown in practice, so it needs to be estimated from the data. A log-likelihood ratio ($\log LR$) test statistic, formulated by the null hypothesis of r rank or $\mathrm{H}(r)$ against the alternative hypothesis $\mathrm{H}(p)$, is designated as $\log LR[\mathrm{H}(r)|\mathrm{H}(p)]$, 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 the method of gamma approximations to calculate 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 interpretable restrictions that can be placed in the estimation of β^*.

Next, a brief review is made with regard to a partial CVAR model and a model reduction procedure. Let us break down equation (2) into $X_t = (Y_t', Z_t')'$ for $Y_t \in R^m$ 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:

$$\alpha = \begin{pmatrix} \alpha_y \\ \alpha_z \end{pmatrix}, \ \Gamma_i = \begin{pmatrix} \Gamma_{y,i} \\ \Gamma_{z,i} \end{pmatrix}, \ \mu = \begin{pmatrix} \mu_y \\ \mu_z \end{pmatrix}, \ \Phi = \begin{pmatrix} \Phi_y \\ \Phi_z \end{pmatrix}, \ \varepsilon_t = \begin{pmatrix} \varepsilon_{y,t} \\ \varepsilon_{z,t} \end{pmatrix}, \ \Sigma = \begin{pmatrix} \Sigma_{yy} & \Sigma_{yz} \\ \Sigma_{zy} & \Sigma_{zz} \end{pmatrix},$$

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:

$$\Delta Y_t = \omega \Delta Z_t + \alpha_y \beta^{*\prime} X_{t-1}^* + \sum_{i=1}^{k-1} \Gamma_{y,i}^* \Delta X_{t-i} + \mu_y^* + \Phi_y^* D_t + \varepsilon_{y,t}^*, \quad (3)$$

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

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 or not $\alpha_z = 0$ holds 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 (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. Equilibrium in the Money Market

With price homogeneity imposed, demand for real broad money may be postulated 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 $i = 1, \cdots, 4$. Equation (5) is thus seen as a representation of the equilibrium condition for the money market.

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

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

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

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

Or, if it turns out that the interest rates do not play any critical role *empirically*, we may find $\gamma_2 = 0$ as follows:

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

Thus,

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

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. Usually the velocity of money and inflation show a co-movement, with other variables 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 analysis conducted in this paper, we consider the money market function (6) or (7) as a candidate for one of the underlying long-run economic linkages embedded in the data.

4.2. Interest-Rate-Based Monetary Policy

Let us turn to monetary policy. Taylor (1993) suggests a simple policy rule in which the short-term policy rate responds 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^* is the natural logarithm of potential output; π_t and π_t^* are actual inflation and target inflation, respectively; and η_0 is 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.

It is possible to reckon the above Taylor rule as a way to implement inflation targeting. The Taylor rule may be seen as a short-run operating procedure for a medium-run inflation target, while inflation targeting itself yields a medium-run plan. The output gap included in the rule can be counted as a measure of inflation pressure. If we could assume that the potential output shows a log linear deterministic trend, then $y_t^* = \theta_1 + \theta_2 t$ with $\theta_i > 0$ for $i = 1,2$. Furthermore, supposing that the inflation target could be assumed to be time-varying, approximated as $\pi_t^* = \theta_3 + \theta_4 t$ with $\theta_3 > 0$ and $\theta_4 \neq 0$, 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 $\nu_3 = -\nu_1\theta_2 - \nu_2\theta_4$. Laurent (1988) as well as Bernanke and Blinder (1992) argue that the spread between the Federal funds rate and yield on government bonds is a useful indicator of the stance of monetary policy. The reasons for such an argument are summarized as follows: (i) the long-term rate should incorporate inflation expectations of all interest rates but is relatively insensitive to short-run variations in monetary ease or tightness and (ii) a conceivable predominant force underlying the long-term rate behavior comes from expected inflation. Bernanke and Blinder (1992) further note that the Federal funds rate, not the bond rate, dominates movements in the spread. The same holds for the interest rates and 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 a modified Taylor-type target, where the bond rate is used instead of inflation, seem congruent with the data in the U.S. Assuming that the long-term rate moves along with inflation and noting that inflation is already present in equation (9), we add to equation (9) the information as measured by the real long-term rate, $i_t^l - \pi_t$, to obtain

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

where $\bar{r}^l$ is the mean of the real long-term rate, which is assumed to be time-invariant, and $\nu_0^* = \nu_0 - \nu_4\bar{r}^l$; it is simply assumed that the factor productivity of the economy, which may directly affect the real long-term interest rate, grows along a long-term trend over the period. If there is a one-to-one effect from the long-term rate to the short-term rate, then $\nu_4 = 1$, and thus we find:

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

It follows from $\nu_3 = -\nu_1\theta_2 - \nu_2\theta_4$ that

$$i_t^s = \nu_0^* + \nu_1(y_t - \theta_5 t) + \nu_2\pi_t + i_t^l, \tag{11}$$

where $\theta_5 = \theta_2 + (\nu_2/\nu_1)\theta_4$. Note that the last terms $\nu_2\pi_t + i_t^l$ show the combined influence of inflation and expected inflation (embedded in the long-term interest rate) on the short-term policy rate, whose magnitude is near $\nu_2 + 1$. Rearranging the equation above, the following is obtained:

$$i_t^s - i_t^l = \nu_0^* + \nu_1(y_t - \theta_5 t) + \nu_2\pi_t. \tag{12}$$

Thus equation (12) shows a monetary policy target expressed in terms of the interest rate spread, $i_t^s - i_t^l$. In the following empirical analysis performed in this paper, we regard the monetary policy rule function (12) as a candidate for one of the long-run economic relationships in the 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:

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

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 estimates 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 for estimation runs 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 thus judged that the model's short-run dynamics should be adjusted in such a manner as suggested by Kurita and Nielsen (2009) with the selection of $k = 2$ for the model's lag length. Namely, a (three-quarter) lagged second-order dynamic term for inflation,

$\Delta^2\pi_{t-3}$, is added to the VAR(2) model unrestrictedly, with the result that the autocorrelation problem is addressed while the standard asymptotic theory for the cointegrating rank test remains valid. Furthermore, the data seem to be much influenced by outliers in 1998.3, 2008.1 and 2009.1. As described in Section 2, the first outlier is possibly due to a policy response to an 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, taking either 0 or 1, are included unrestrictedly in the model: $D_{1,t} = 1(1998.3)$, $D_{2,t} = 1(2008.1)$, $D_{3,t} = 1(2009.1)$ 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*: *k*th-order autocorrelation (F_{ar}: see Godfrey, 1978, Nielsen, 2006), *k*th-order autoregressive conditional heteroskedasticity or ARCH (F_{arch}: see Engle, 1982), 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*

	π_t	$i_t^s - i_t^l$	$m_t - p_t$	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*

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

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 standard 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 required in the econometric exploration.

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. Prior to the exploration of such a parsimonious model, we address the issue of identifying equilibrium correction mechanisms in the next sub-section.

Table 3. *Weak Exogeneity Tests*

π_t	$i_t^s - i_t^l$	$m_t - p_t$	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.

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), thus encouraging 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 we find the revealed structure can be expressed as follows:

$$\hat{\alpha}\,\hat{\beta}^{*\prime} = \begin{bmatrix} \underset{(0.048)}{-0.269} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(0.042)}{-0.207} \\ \underset{(-)}{0} & \underset{(-)}{0} \\ \underset{(-)}{0} & \underset{(-)}{0} \end{bmatrix} \begin{bmatrix} \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.117)}{0.291} & \underset{(-)}{-0.291} & \underset{(0.00028)}{-0.00063} \\ \underset{(-)}{0} & \underset{(-)}{1} & \underset{(-)}{0} & \underset{(0.050)}{-0.134} & \underset{(0.00035)}{0.00083} \end{bmatrix}, \qquad (13)$$

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

$$\pi_{t-1} = 0.291(p_{t-1} - m_{t-1} + y_{t-1}) + 0.00063t + v_{1,t-1}, \tag{14}$$

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

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

where $y_{t-1}^{*} = 0.00619t$ and $y_{t-1} - y_{t-1}^{*}$ is thus 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 economic 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 important factors for the dynamics of a parsimonious VECM pursued in the next section. Before moving on, 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. The revealed structure may thus be seen as a set of common features of economies that in principle obey monetary policy rules, although further empirical explorations of various countries and regions are required to reach a decisive conclusion on this hypothetical view.

6. A Reduced Vector Equilibrium Correction Model

This section aims to achieve a parsimonious VECM for $\Delta\pi_t$ and $\Delta(i_t^s - i_t^l)$. 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:

$$\Delta\pi_t = -\underset{(0.062)}{0.335} ecm_{1,t-1} + \underset{(0.111)}{0.240}\Delta m_{t-1} + \underset{(0.081)}{0.327}\Delta\pi_{t-1} + \underset{(0.059)}{0.323}\Delta^2\pi_{t-3} + \underset{(0.015)}{0.074}$$

$$- \underset{(0.025)}{0.098}\, D_{3,t} + \hat{\varepsilon}_{1,t},$$

$$\Delta(i_t^s - i_t^l) = -\underset{(0.042)}{0.188} ecm_{2,t-1} + \underset{(0.074)}{0.320}\Delta(i_{t-1}^s - i_{t-1}^l) - \underset{(0.014)}{0.042}\Delta(m_{t-1} - p_{t-1})$$

$$-\underset{(0.006)}{0.025} + \underset{(0.004)}{0.028}\, D_{1,t} - \underset{(0.004)}{0.016}\, D_{2,t} + \hat{\varepsilon}_{2,t}, \tag{16}$$

where

$$ecm_{1,t-1} = \pi_{t-1} + 0.291(m_{t-1} - p_{t-1} - y_{t-1}) - 0.00063t,$$
$$ecm_{2,t-1} = i_{t-1}^s - i_{t-1}^l - 0.134y_{t-1} + 0.00083t,$$

$$\text{Autocorr.}\,[F_{ar}(20,166)] = 0.74\ [0.78],$$
$$\text{Hetero.}\,[F_{het}(69,204)] = 0.92\ [0.65],$$
$$\text{Normality}[\chi_{nd}^2(4)] = 4.23\ [0.38].$$

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 Fig. 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 Fig. 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, Fig. 2 (a) and (c) report scaled residuals, while Fig. 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.

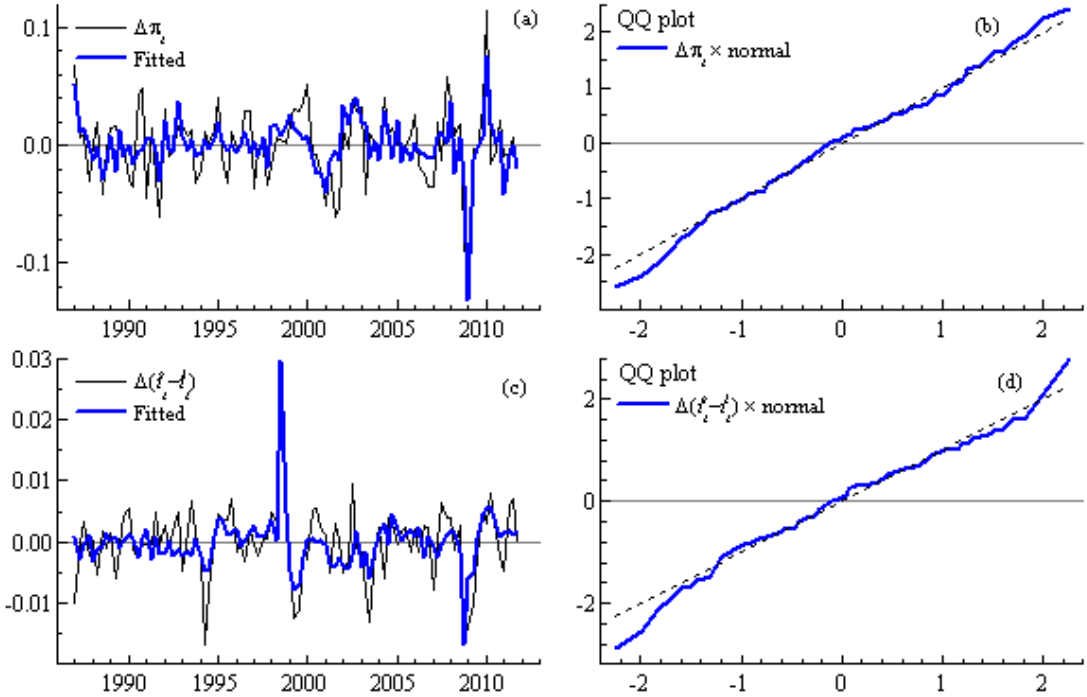

Fig. 1. Fitted Values and Residual QQ Plots

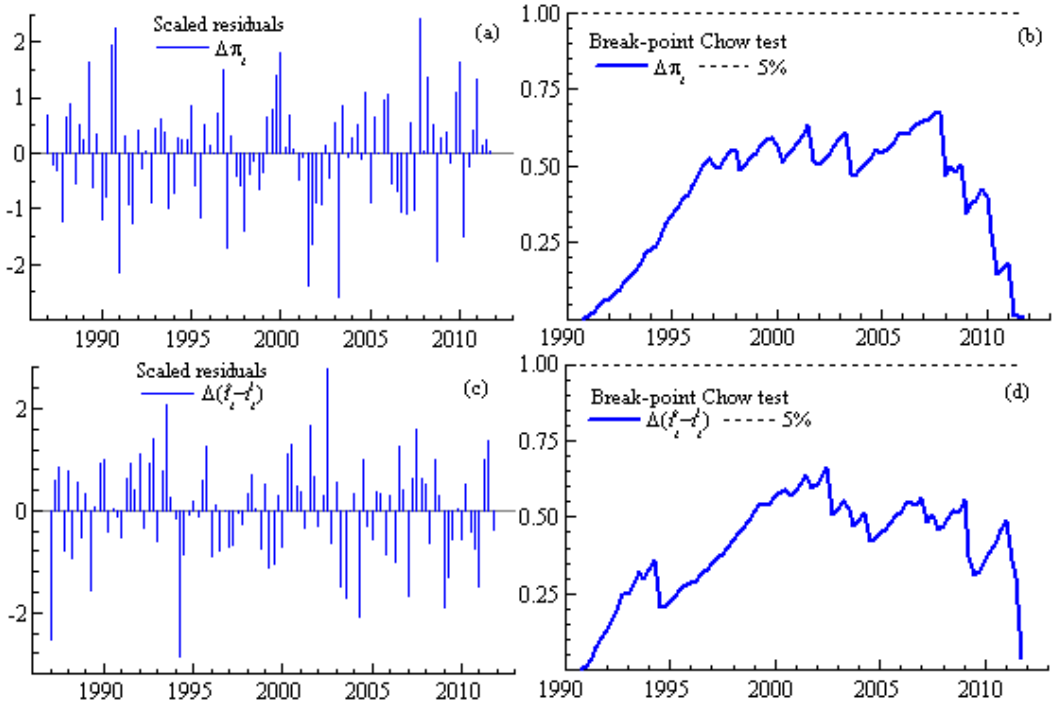

Fig. 2. Scaled Residuals and Recursive Chow Tests

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 $\Delta(i_t^s - i_t^l)$, 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 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 $\Delta(i_t^s - i_t^l)$ 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.

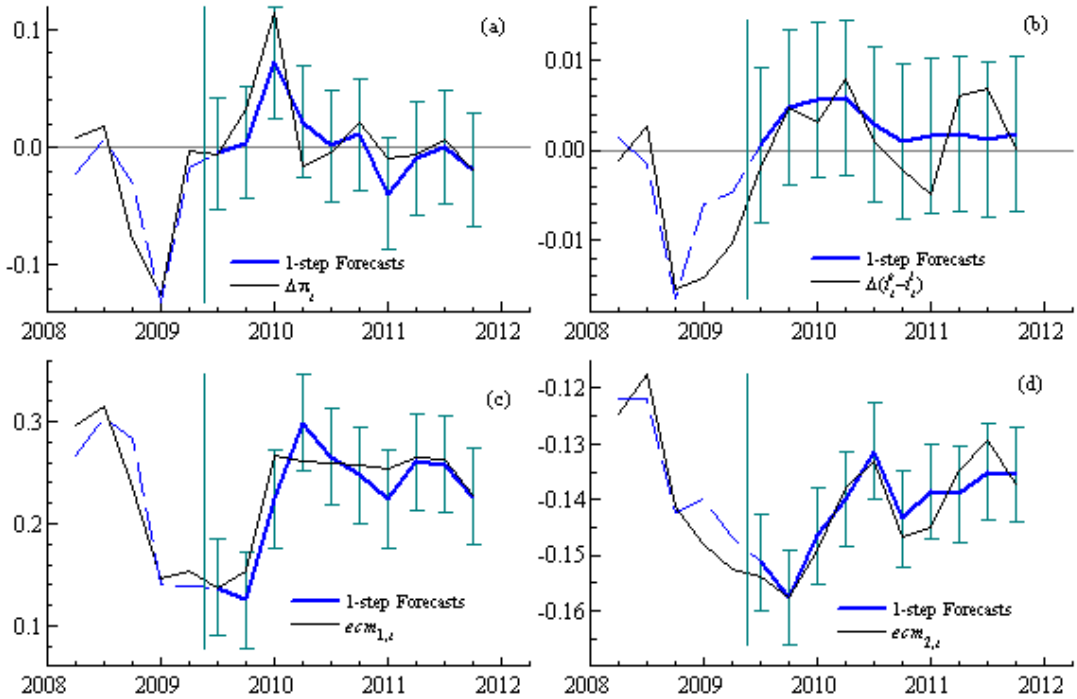

Fig. 3. 1-Step Forecasts Generated from the VECM

In order to demonstrate the practical usefulness of the VECM, model (16) is re-estimated using a truncated sample period, 1986.2 - 2009.2, so that the sequences of 1-step forecasts of $\Delta\pi_t$, $\Delta(i_t^s - i_t^l)$, $ecm_{1,t}$ and $ecm_{2,t}$ for 2009.3 - 2011.4 are derived from it. Note that it turns out that model (16) is not significantly influenced by any contemporaneous regressors; thus the model 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 $\Delta(i_t^s - i_t^l)$. 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.

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 was demonstrated as reliable on 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.

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.

A.2. Sources and Notes

The data of π_t, $m_t - p_t$, y_t and $i_t^s - i_t^l$ 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.

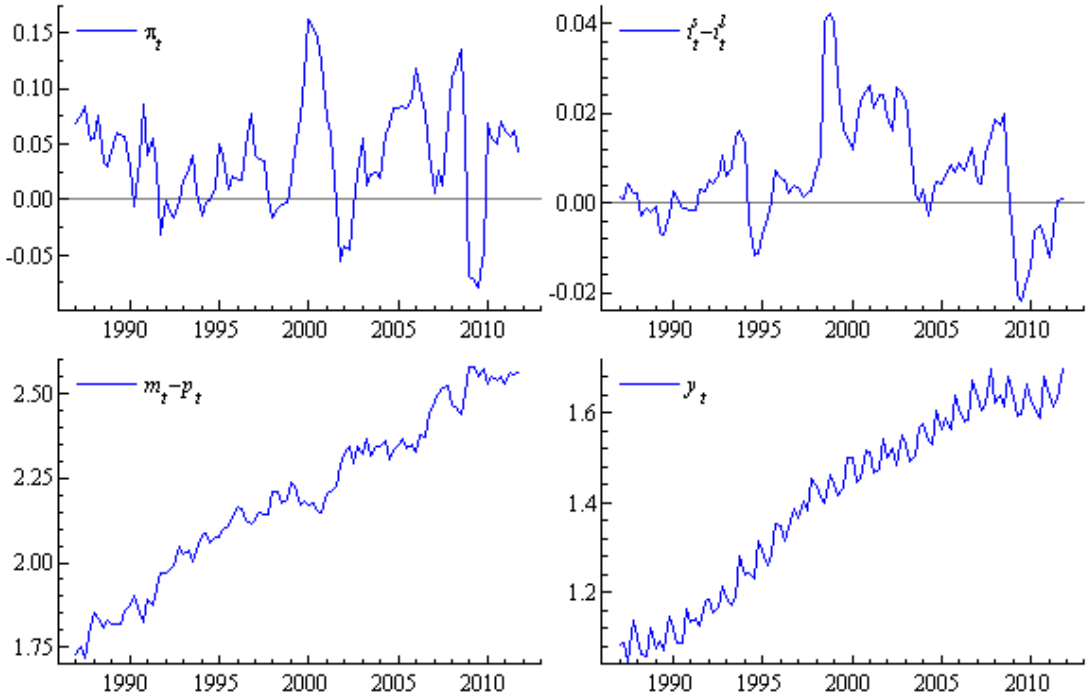

Fig. A. Overview of the Data

	$i_t^s - i_t^l$	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

Table A. Correlation Matrix (1987Q1 - 2011Q4)

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