# The Cointegrated Vector Autoregressive Model: Concepts and Evidence

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To Torjus and Markus

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# Contents

# 1 Chapter 1

## Introduction

# 1 Introduction[1]

*"It is fairly familiar knowledge that we sometimes obtain between quantities varying with the time (time-variables) quite high correlations to which we cannot attach any physical significance whatever, although under the ordinary test the correlation would be held to be certainly "significant.""*
— Yule (1926, p. 2)

Standard estimation and inference theory assumes that time series are stationary processes with their first and second order moments, mean and variance, not depending on time. However, many macroeconomic time series, such as real GDP, household consumption and real disposable income, tend to grow over time with non-constant mean and variance. They are thus non-stationary times series which, if not correctly handled in empirical modelling, may lead to misleading results and conclusions.[2]

Around 60 years after the analysis of Yule (1926), who showed that non-stationary independent time series create nonsense correlations between them, Engle and Granger (1987) introduced the concept of *cointegration* as the statistical counterpart of equilibrium correction. The concept of cointegration has since been of tremendous importance for empirical modelling of time series as it contains the key to handling non-stationarity, and hence also the key to avoiding the problem of nonsense correlations or spurious regressions in the terminology of Granger and Newbold (1974). Roughly speaking, non-stationary time series are said to be cointegrated if some linear combination(s) of them is(are) stationary.[3] Such linear combinations are typically interpreted in light of economic theory and are often referred to as long-run equilibrium relationships in the literature. Likewise, the discrepancy from long-run equilibrium is called an equilibrium correction mechanism, an explanatory variable which was sometimes included in dynamic regression models long before the concept of cointegration of Engle and Granger (1987); see for instance Sargan (1964) and Davidson *et al.* (1978). Importantly, the Engle and Granger (1987) representation theorem says that non-stationary time series that are cointegrated have a stationary equilibrium correction representation and vice versa. Because the property of stationarity is achieved, the equilibrium correction representation is amenable to standard estimation and inference theory.

The cointegrated vector autoregressive (CVAR) model includes differences of variables and the cointegration between them in a multivariate context in order to analyse both short-run and long-run effects within the same model. The CVAR model thus represents a system where variables are pushed away from long-run relationships by exogenous shocks and where short-run adjustments force them back towards long-run relationships. Assumptions from economic theory can also be translated into testable hypotheses on the cointegration relationships of key variables and their associated adjustment coefficients in the CVAR model. Perhaps for the above reasons, a vast literature on the

[1] I am grateful to Thor Andreas Aursland, Thomas von Brasch, Ådne Cappelen, Eilev Jansen, Terje Skjerpen and Anders Rygh Swensen for valuable comments on a draft of this introductory chapter.

[2] In the following, the terms stationary and non-stationary time series are used interchangeably with the terms integrated time series of orders zero and one, denoted $I(0)$ and $I(1)$, respectively.

[3] More formally, when two time series that are $I(1)$ are cointegrated, denoted CI(1,1), then a linear combination is $I(0)$; see for instance Banerjee *et al.* (1993, p. 3).

CVAR model, both theoretical and empirical, has emerged over the last three decades. In fact, a Google Scholar search on the term "the cointegrated VAR model" provides around 68 thousand hits on articles and books, of which Ericsson (1992), Banerjee *et al.* (1993), Hendry (1995), Johansen (1995), Juselius (2006), Lütkepohl (2007), Bjørnland and Thorsrud (2015), Juselius (2015, 2019) and Nymoen (2019) are some examples of overviews of the CVAR model in theory and practice.

The purpose of this thesis is to contribute to the existing empirical literature by using the CVAR model to evaluate the empirical performance of several theoretical aspects of economics ranging from consumer behaviour, inflation dynamics, exchange rate pass-through, export behaviour to OPEC behaviour.

The rest of this introductory chapter is organised as follows: Subsection 1.1 provides an outline of the main concepts of the CVAR model as a point of reference from which the thesis evolves. Subsection 1.2 gives a summary of the main empirical evidence from each of the chapters in the thesis. Subsection 1.3 provides concluding remarks including some unanswered questions that arise from the thesis and that are left for future research.

## 1.1 Outline of concepts

To provide background for the concept of cointegration in the VAR model, I begin this subsection with an illustrative example of a spurious regression between two independent $I(1)$-variables. Then, I outline the main concepts of cointegration analysis in a multivariate setting. Three different cases are considered: a full system, in which all $I(1)$-variables are modelled, a partial system, in which some $I(1)$-variables that are conditional upon some other $I(1)$-variables are modelled, and a system, in which some $I(1)$-variables are modelled together with some $I(0)$-variables.[4] Finally, I outline important econometric issues that need to be addressed in any CVAR analysis to ensure that all systematic aspects of the data are satisfactorily described by the underlying VAR model.

### 1.1.1 Spurious regression

Consider the simple regression model

$$x_{1,t} = \tau_0 + \tau_1 x_{2,t} + \epsilon_t, \tag{1}$$

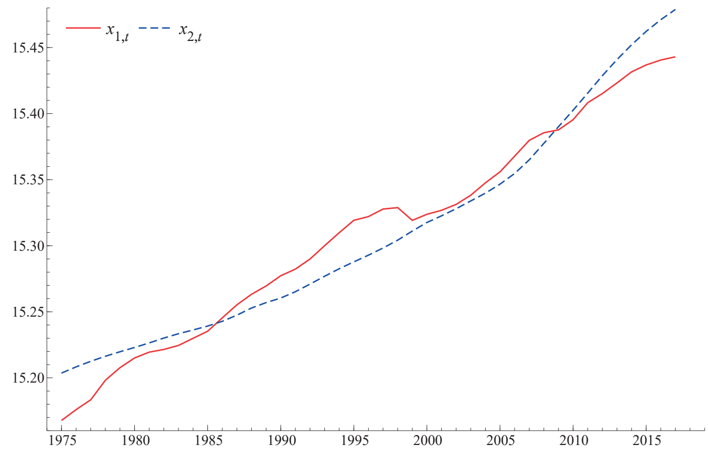
where $x_{1,t}$ and $x_{2,t}$ are the logarithms of the real GDP of Colombia and the population of Norway, respectively, $\epsilon_t$ is the residual and $\tau_0$ and $\tau_1$ are the regression coefficients. Figure 1 shows that neither $x_{1,t}$ nor $x_{2,t}$ are stationary time series as they trend upwards over the sample period. There also seems to be a relationship between the two time series, even though neither economic theory nor common sense tell us so. As such, we expect $\hat{\tau}_1$ to be zero.

Nonetheless, estimating (1) by ordinary least squares (OLS) produces

$$\hat{x}_{1,t} = \underset{(2.98)}{-61.29} + \underset{(0.19)}{5.29} x_{2,t}, \tag{2}$$

[4]Here and throughout the thesis, we do not consider situations in which a variable is of a higher order than one; see for instance Juselius (2006, chapters 16-18) for analyses of $I(2)$-variables.

Figure 1: The real GDP of Colombia ($x_{1,t}$) and the population of Norway ($x_{2,t}$)




*Notes*: Sample period: 1975 − 2017. The mean and range of the logarithms of the real GDP of Colombia are matched to the mean and range of the logarithms of the population of Norway. Sources: OECD and Statistics Norway.

with estimated standard errors in parentheses. Clearly, $\hat{\tau}_1$ is far from zero and highly significant. Accordingly, we are tempted to conclude that there exists a statistically significant relationship between the real GDP of Colombia and the population of Norway, whereas none in fact exists. The regression in (2) is thus an example of a spurious regression along the lines of Granger and Newbold (1974).[5] Because the problem of spurious regression between unrelated non-stationary time series persists with deterministically detrended series, the inclusion of a time trend in (2) is not a remedy; see Banerjee *et al.* (1993, p. 83) and the references therein.

Generally speaking, regression models with unrelated non-stationary time series are likely to give spurious relationships. However, if one or more cointegration relationships exist between non-stationary time series, the concept of cointegration can be regarded as the opposite of spurious regression. The VAR representation of cointegrated variables, which we now turn to, is known as the CVAR model.

[5]Granger and Newbold (1974) point out the likely outcome of spurious regressions using Monte Carlo simulations by means of two independent random walks. A random walk is a special case of an AR(1) process, defined as $x_t = \phi x_{t-1} + \epsilon_t$, where $\epsilon_t \sim i.i.d.N(0, \sigma^2)$, when $\phi = 1$. Hence, such a time series is said to contain a unit root, that is a stochastic trend defined as $x_t = \sum_{i=0}^{t-1} \epsilon_{t-i}$ when $x_0 = 0$, and is thus non-stationary. However, the first difference of a random walk, $\Delta x_t = \epsilon_t$, is stationary, that is $\Delta x_t \sim I(0)$. The AR(1) process, for its part, is stationary when $\phi < 1$. Another source of non-stationarity in a time series is a deterministic trend. Nevertheless, $x_t = \phi_0 + \phi_1 t + \epsilon_t$, where $t$ is a deterministic trend, is called a trend stationary process because stationarity can be achieved by subtracting $t$ from $x_t$, that is $x_t^* = x_t - \phi_1 t$.

### 1.1.2 Full system cointegration analysis

For expositional simplicity, we consider a trivariate VAR ($n = 3$) of second order ($k = 2$) of the form

$$y_t = \sum_{i=1}^{2} \Pi_i y_{t-i} + \epsilon_t, t = k + 1, ..., T, \tag{3}$$

where $y_t = (y_{1,t}, y_{2,t}, y_{3,t})'$ is a (3 x 1) vector of $I(1)$-variables at time $t$, $\Pi_1$ and $\Pi_2$ are (3 x 3) coefficient matrices of $y_{t-1}$ and $y_{t-2}$, respectively, $\epsilon_{k+1}, ..., \epsilon_T$ are independent Gaussian innovations with expectation zero and (3 x 3) covariance matrix $\Omega$, $y_1, ..., y_k$ are kept fixed as initial observations and $T$ is the total number of observations. We have provisionally simplified matters in (3) by not including constants, deterministic trends or dummy variables. Following Johansen (1988), (3) is the point of departure for cointegration analysis. We can rewrite (3) as a CVAR model of the form

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \Pi y_{t-1} + \epsilon_t, t = k + 1, ..., T, \tag{4}$$

where $\Delta$ denotes the difference operator, $\Gamma_1 = -\Pi_2$ is a (3 x 3) coefficient matrix of the first lag of $\Delta y_t$ and $\Pi = (\Pi_1 + \Pi_2 - I)$ is the so-called impact matrix, where $I$ is the identity matrix. The cointegration analysis amounts to determining how many cointegrating vectors are present in $\Pi$. With three variables in the VAR, $\Pi$ has rank $0 \leq r \leq 3$, where $r$ denotes the number of cointegrating vectors. When $\Pi$ has full rank, that is $r = 3$, then (3) is stationary and can be estimated in levels. However, if $\Pi$ has reduced rank two possibilities exist. The first is when $r = 0$, meaning that no cointegrating vectors exist, that is $\Pi = 0$. Then (3) is non-stationary and can be estimated in first differences without loss of information. The second is when $1 \leq r \leq 2$, implying that the impact matrix can be decomposed as $\Pi = \alpha\beta'$, where $\alpha$ and $\beta$ are (3 x $r$) matrices of adjustment coefficients and cointegration coefficients, respectively. Although $y_t$ is $I(1)$, the linear combination $\beta' y_{t-1}$, in which each of the $r$ rows comprises a cointegrating vector, is $I(0)$.[6] The variables in $y_t$ are then said to be cointegrated. Because $\Delta y_t$ is also $I(0)$ it follows that (4) is a stationary CVAR model.[7]

A formal test of the null hypothesis that there are at most $r$ cointegrating vectors in (4) can be based on the likelihood ratio test, the so-called trace test, expressed as

$$\eta_r = -T \sum_{i=r+1}^{3} \ln(1 - \hat{\lambda}_i), r = 0, 1, 2, \tag{5}$$

where $1 \geq \hat{\lambda}_1 \geq \hat{\lambda}_2 \geq \hat{\lambda}_3 \geq 0$ are the estimated eigenvalues from a reduced rank regression of $\Delta y_t$ on $y_{t-1}$ corrected for $\Delta y_{t-1}$ by OLS; see for instance Johansen (1995, chapter 6). The testing procedure is sequential and the cointegration rank is determined as zero if $\eta_0$ is not significant and as $r + 1$ if the last significant statistic is $\eta_r$ using available critical values in the case of no deterministic terms; see for instance Juselius (2006, appendix A,

[6] Also, variables in $y_t$ that are cointegrated will share one or more common stochastic trends which are eliminated by the linear combination $\beta' y_{t-1}$.

[7] A model in which the variables on both sides of the equality sign are of the same order of integration is sometimes referred to as a balanced model; see Banerjee *et al.* (1993, p. 166).

case 1).[8] Estimates of $\beta$ are then obtained as the eigenvectors associated with the $r$ largest eigenvalues.

Generally speaking, significant cointegrating vectors are not uniquely identified. In order to see this, we can introduce a non-singular ($r$ x $r$) matrix $Q$, such that

$$\Pi = \alpha\beta' = \alpha Q Q^{-1}\beta' = \alpha^*\beta^{*\prime}, \tag{6}$$

where $\alpha^* = \alpha Q$, $\beta^* = \beta Q'^{-1}$ and $QQ^{-1} = I$. As $\beta^*$ is not equal to $\beta$ when $Q' \neq I$, the cointegrating vectors are not unique and restrictions must be imposed on $\beta$ and/or $\alpha$ in order to achieve identification. From (6), we see that identification of the cointegrating vectors is driven by $Q$. Thus $r^2$ restrictions are needed to obtain unique identification of the cointegrating vectors. As an example, we consider the case where $r = 2$ in (4) and choose $Q$ such that it equals the ($r$ x $r$) upper block of $\beta$. Then the identifying restrictions become

$$\beta^* = \begin{bmatrix} I_r \\ \beta_{n-r} \end{bmatrix}, \tag{7}$$

where $\beta_{n-r}$ is $((n-r)$ x $r)$. We observe that (7) implies $\beta_{11} = \beta_{22} = 1$ (normalisation) and $\beta_{12} = \beta_{21} = 0$, four restrictions in total, such that the two uniquely identifying cointegrating vectors become $\beta_1^{*\prime} y_{t-1} = y_{1,t-1} + \beta_{31}y_{3,t-1}$ and $\beta_2^{*\prime} y_{t-1} = y_{2,t-1} + \beta_{32}y_{3,t-1}$. In practice, economic theory can also often help in identifying cointegrating vectors. A thorough discussion on how to identify cointegrating vectors can be found in Lütkepohl (2007) and Juselius (2006).

Once the cointegration rank is determined, we can perform likelihood ratio tests of various restrictions on $\beta$ and $\alpha$ in (4) by means of the formula

$$LR = -2(L_{max,R} - L_{max,U}), \tag{8}$$

where $L_{max,R}$ and $L_{max,U}$ are the maximum log-likelihood values of the restricted and the unrestricted model, respectively. The $LR$ statistic is asymptotically $\chi^2$-distributed with degrees of freedom equal to the number of restrictions imposed on the CVAR model; see Johansen (1995, chapter 7). An example of this test is a long-run exclusion restriction of a variable in $y_t$, say $y_{3,t}$, given as a zero restriction on $\beta$. If accepted, $y_{3,t}$ can be omitted from the cointegrating relationships altogether. Another example is a test of weak exogeneity of a variable in $y_t$, again say $y_{3,t}$, with respect to $\beta$ given as a zero restriction on $\alpha$.[9] If accepted, $y_{3,t}$ is not equilibrium correcting or is not reacting to deviations from the long-run relationships.

The likelihood ratio test can also be carried out to test conditional expectations of future variables, which are often inherent in economic models, as restrictions on the coefficients in (4). For instance, consider a simple new Keynesian Phillips curve of the form

$$\Delta y_{1,t} = \delta E_t \Delta y_{1,t+1} - \lambda(y_{1,t} - \psi_1 y_{2,t} - \psi_2 y_{3,t}), \tag{9}$$

[8] In general, the critical values depend on deterministic terms (constant and trend) and how they are included (unrestricted or restricted) in the VAR; see Juselius (2006, appendix A, cases 2-5). We shall return to these issues below. An alternative test for determining the rank order is the so-called max eigenvalue test, where the null hypothesis is that of a specific number of cointegration vectors, say $r = 1$, against the alternative of that of $r = 2$; see for instance Nymoen (2019, chapter 10). However, the trace test seems now to be the preferred test in practice.

[9] See for instance Ericsson (1992) for a formal definition of a weakly exogenous variable.

where $y_{1,t}, y_{2,t}$ and $y_{3,t}$ denote the price level, the unit labour costs and the unit import costs, respectively. Using vector notation and assuming the rank to be one, (9) can be written as

$$c' E_t \Delta y_{t+1} = d'_{-1} \Delta y_t + d' y_t, \tag{10}$$

where $c = (1, 0, 0)'$, $d_{-1} = (1/\delta, 0, 0)'$ and $d = (\lambda/\delta, -\lambda/\delta\psi_1, -\lambda/\delta\psi_2)'$. By leading (4) one period and taking the conditional expectations at time $t$ yields $E_t \Delta y_{t+1} = \Gamma_1 \Delta y_t + \Pi y_t$. Inserting this expression into (10) implies that the restrictions $c'\Gamma_1 = d'_{-1}$ and $c'\Pi = d'$ must be satisfied in order for non-rejection of the new Keynesian Phillips curve to be the outcome of the likelihood ratio test. Accordingly, the conditional expectations of future inflation in (9) involve restrictions on both the short-run and the long-run parameters in (4). In order to construct the likelihood ratio test by means of (8), we thus have to work out the maximum log-likelihood value of the trivariate CVAR model, both with and without the conditional expectations restrictions imposed; see Johansen and Swensen (1999, 2008) for details.

### 1.1.3 Partial system cointegration analysis

It is common practice to model some $I(1)$-variables conditional on some other $I(1)$-variables in partial systems when the total number of variables is too large for a full system to be feasible. Harbo *et al.* (1998) discuss the problem of rank determination in such partial systems. To illustrate, suppose that $y_t$ is decomposed into $y'_t = (x'_t, z'_t)$, where $x'_t = (y_{1,t}, y_{2,t})$ and $z'_t = (y_{3,t})$, and that the coefficient matrices are decomposed conformably with $y_t$. Then, the full system in (4) can be expressed as a conditional model of $\Delta x_t$ of the form

$$\Delta x_t = \omega \Delta z_t + (\Gamma_{11} - \omega\Gamma_{21})\Delta y_{t-1} + (\alpha_1 - \omega\alpha_2)\beta' y_{t-1} + \epsilon_{1,t} - \omega\epsilon_{2,t}, \tag{11}$$

where $\omega = \Omega_{12}\Omega_{22}^{-1}$, and a marginal model of $\Delta z_t$ of the form

$$\Delta z_t = \Gamma_{21} \Delta y_{t-1} + \alpha_2 \beta' y_{t-1} + \epsilon_{2,t}. \tag{12}$$

According to Harbo *et al.* (1998), a sufficient condition for efficient inference of the cointegration rank is that the conditioning variable, $z_t$, is weakly exogenous for $\beta$, the parameters of interest. When $\alpha_2 = 0$ in (12), there is no information about $\beta$ in the marginal model and $z_t$ is thus weakly exogenous. The determination of the rank can then be based solely on the conditional model in (11) as a partial system without loss of information. By conditioning on $z_t$, we may achieve more stable parameters in (11) than in (4), especially if the parameters in (12) are non-constant; see Juselius (2006, p. 198). Ideally, we should formally test $\alpha_2 = 0$ in the full system, rather than impose weak exogeneity from the outset, in the manner described above. However, in those cases when $z_t$ is likely to be weakly exogenous, for instance due to a priori beliefs, formal testing may not be necessary.

Nevertheless, cointegration in (11) is consistent with $0 < r \leq 2$, which means that the maximum rank order may be equal to the number of modelled $I(1)$-variables in $y_t$. If so, the impact matrix has full rank. Notably, $z_t$ is by assumption a common stochastic trend in the partial system, that is it corresponds to a unit root in the full system; see Juselius (2006, p. 198). As shown by Harbo *et al.* (1998), the asymptotic critical values for the trace test generally depend on the number of weakly exogenous variables as well as the specification of deterministic components (constant and trend) in a partial system. For this reason, a set of tables of critical values for different specifications of a partial system is also available in, for instance, Pesaran *et al.* (2000) and Doornik (2003).

### 1.1.4 Cointegration analysis with stationary variables

An issue which often arises in cointegration analysis is how to treat stationary explanatory variables in the VAR model. Rahbek and Mosconi (1999) address this issue when testing for cointegration rank. The approach of Rahbek and Mosconi (1999) may be illustrated by means of an extended version of (4) of the form

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \alpha\beta' y_{t-1} + \sum_{i=0}^{2} \Psi_i z_{t-i} + \rho + \epsilon_t, t = k+1, ..., T, \tag{13}$$

where $z_t = y_{4,t}$ is a supposedly stationary explanatory variable and $\rho$ is a constant. As pointed out by Rahbek and Mosconi (1999), the asymptotic distribution of the trace test statistic generally depends on nuisance parameters due to the presence of stationary regressors and the inclusion of deterministic components. The suggested approach is therefore to analyse an augmented version of (13) of the form

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \alpha\beta^{*\prime} \begin{pmatrix} y_{t-1} \\ \sum_{i=1}^{t} z_i \\ t \end{pmatrix} + \sum_{i=0}^{2} \Psi_i z_{t-i} + \rho + \epsilon_t, t = k+1, ..., T, \tag{14}$$

which makes the inference on the cointegrating rank asymptotically free of nuisance parameters because of the inclusion of the accumulated level of $z_t$ in the cointegration relations. The linear trend is also included in the cointegration relations since (13) contains a constant. Once the rank is determined by means of (14) using critical values tabulated in Harbo *et al.* (1998), we may test the linear restrictions that there is no accumulated level of $z_t$ and no linear trend in the cointegrating relations by considering the hypothesis $\beta^* = (\beta', 0')'$. The likelihood ratio test for this hypothesis, which is asymptotically $\chi^2(2r)$-distributed with $2r$ degrees of freedom, may in line with Rahbek and Mosconi (1999) be regarded as a misspecification test of (13).

### 1.1.5 Well-specified underlying VAR

As stressed by Juselius (2019) among others, a CVAR analysis has as an underlying premise that all systematic aspects of the data are satisfactorily described by the underlying VAR model. Accordingly, a convincing CVAR analysis ought to address important econometric issues such as the time series properties of the variables involved, the validity of the assumption of identically and independently normally distributed residuals, the role of deterministic components (constant and trend) in the VAR model and changes in the structure of the VAR model due to extraordinary events and economic regime shifts. As will become clear in the various applications of the CVAR model in this thesis, we search for statistically well-specified underlying VAR models along these lines as premises for valid statistical inference.

Generally speaking, we let economic theory dictate the variables to be included in the VAR model. Once the VAR model has been specified in this way, we examine time series properties of the chosen variables by means of both graphical inspections and the widely used augmented Dickey-Fuller (ADF) unit root test.[10] Because most of the macroeconomic time series considered in this thesis exhibit a clear trending behaviour, we rely on the ADF-test of the form

$$\Delta y_t = \vartheta_0 + \vartheta_1 t + \varphi y_{t-1} + \sum_{i=1}^{p-1} \theta_i \Delta y_{t-i} + \epsilon_t, \tag{15}$$

where $y_t$ is a single time series, say $y_{1,t}$, $\vartheta_0$ is an intercept, $\vartheta_1$ is a slope coefficient, $\theta_i$ are coefficients of $\Delta y_{t-i}$ and $p$ is the lag length in the underlying AR($p$)-process to take into account significant autocorrelations in the residuals. A test for a unit root in (15) involves testing the null hypothesis $H_0$: $\varphi = 0$, that is non-stationarity, against the alternative hypothesis $H_1$: $\varphi < 0$, that is stationarity, by comparing the $t$-statistic of the OLS estimate for $\varphi$ with the relevant critical values from the Dickey-Fuller distribution; see for instance Ericsson and MacKinnon (2002) for tabulated critical values.[11]

Next, we decide on the appropriate lag length of the VAR model to ensure no serious departures from the assumption of identically and independently normally distributed residuals. As a guide to choosing the appropriate lag length, we rely on the Akaike's information criterion, likelihood ratio tests for sequential model reduction and tests for residual autocorrelation, heteroscedasticity and non-normality (skewness and excess kurtosis); see for instance Juselius (2006, chapter 4) for details. Since satisfactory empirical fit is important for statistical inference in this thesis, we are generally reluctant to include too few lags in the VAR model, which may cause loss of valuable information and autocorrelated residuals. That said, we should bear in mind that too many lags may produce an overparameterised model, and thus the possibility of additional parameter estimate uncertainty.

As previously noted, the presence of deterministic components (constant and trend) and how they are specified, in a full system, a partial system or a system with stationary explanatory variables, affects the asymptotic distribution of the trace test statistic, and thus the rank determination. We briefly illustrate how such components are incorporated in the VAR models in this thesis by means of an extended version of the full system in (3) expressed as

$$y_t = \sum_{i=1}^{2} \Pi_i y_{t-i} + \varpi_0 + \varpi_1 t + \epsilon_t, t = k+1, ..., T, \tag{16}$$

where $\varpi_0$ and $\varpi_1$ are (3 x 1) vectors of intercepts and slope coefficients, respectively. A common practice with trending behaviour in macroeconomic time series, which we follow in the empirical analyses, is to *restrict* the linear trend to lying within the cointegration space, whereas the intercepts are kept *unrestricted* in the VAR model. This means that we allow for linear, but not quadratic trends, in the variables. In addition, if the trend turns

[10]See for instance Patterson (2011) for an overview of various other tests for a unit root in a time series.

[11]As a rule of thumb, the Dickey-Fuller regression should generally include the order of deterministic trends necessary to take into account the trending behaviour of a series under both the null and the alternative hypothesis; see Patterson (2011, Chapter 6). The trending behaviour of the macroeconomic time series in our cases is taken into account under the null hypothesis by the inclusion of the intercept in (15), whereas under the alternative hypothesis the deterministic trend is part of the data generating process.

out to be *in*significant in the corresponding CVAR model using a likelihood ratio test as outlined above, then the linear combination $\beta' y_{t-1}$ eliminates both the common stochastic trends and the deterministic trends in the variables. The cointegration relations in such cases are stationary and not trend-stationary.

Finally, large outliers in the data, and hence the possibility of violation of the normality assumption for the residuals, sometimes show up in empirical work due to extraordinary events and economic regime shifts. A sudden huge drop in exchange rates and in electricity prices (due to changes in temperature) are examples of the former, whereas financial deregulation of credit markets and a monetary policy change from exchange rate to inflation targeting are examples of the latter. Failing to account for outliers by means of dummy variables is likely to produce structural misspecification of the VAR model. Using dummy variables, the full system in (4), as an example, may be reformulated as

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \Pi y_{t-1} + \Phi_0 D_{0,t} + \Phi_1 D_{1,t} + \Phi_2 D_{2,t} + \epsilon_t, t = k+1, ..., T, \tag{17}$$

where $D_{0,t}$, $D_{1,t}$ and $D_{2,t}$ are vectors of permanent intervention dummy variables $(..., 0, 0, 1, 0, 0, ...)$, transitory shock dummy variables $(..., 0, 0, 1, -1, 0, 0, ...)$ and mean-shift dummy variables $(..., 0, 0, 1, 1, ...)$, respectively, in the terminology of Juselius (2006, chapter 6). If $D_{0,t}$, $D_{1,t}$ and $D_{2,t}$ enter (17) unrestrictedly, then the permanent intervention dummy accounts for a large impulse in $\Delta y_t$ and cumulates to a level shift in $y_t$, the transitory shock dummy accounts for two successive impulses of opposite sign in $\Delta y_t$ and cumulates to a single impulse in $y_t$, and the mean-shift dummy accounts for a mean shift in $\Delta y_t$ and cumulates to a broken trend in $y_t$. However, if $D_{0,t}$, $D_{1,t}$ and $D_{2,t}$ are restricted to the cointegration relations, then these dummy variables do not cumulate in $y_t$. For example, $D_{2,t}$ restricted to the cointegration relations means that a mean-shift in $\beta' y_{t-1}$ is modelled without a broken trend in $y_t$. Although permanent intervention and transitory shock dummy variables do not affect the asymptotic distribution of the trace test statistic, they may do so in finite samples; see Juselius (2006, chapter 8). However, as discussed by Doornik *et al.* (1998) among others, the asymptotic distribution is influenced by mean-shift dummy variables in the VAR model just as it is in the cases of a constant and a trend.

A related issue to mean-shift dummy variables concerns structural breaks, which are sometimes of relevance for macroeconomic time series to render a well-specified underlying VAR model for the cointegration analysis. For example, the household saving ratio in Norway increased considerably in the wake of the financial crisis in 2008, possibly because of a downward shift in the equilibrium level of consumption as a result of increased uncertainty. The approach by Johansen *et al.* (2000) allows for such a structural break in the long-run relationship by means of a full system which takes into account the possibility of separate deterministic trends in the sub-period prior to the financial crisis, $1, \ldots, T_1$, and in the sub-period after the financial crisis, $T_1 + 1, \ldots, T$.[12] The basic idea is to permit a CVAR model for each sub-period, so that the parameters of the stochastic components are the same for both sub-periods, while the parameters of the deterministic trend are not the same, that is there is a structural break. Formally, let $T_0 = 0$ and $T_2 = T$. If $ID_{j,t} = 1$ for $t = T_{j-1}$ and $ID_{j,t} = 0$ otherwise, such that $ID_{j,t-i}$ is an indicator for the $i$th observation in the $j$th period, $j = 1, 2$. Then, it follows that $SD_{j,t} = \Sigma_{i=k+1}^{T_j - T_{j-1}} ID_{j,t-i} = 1$

[12]Generally, the approach by Johansen *et al.* (2000) allows for any pre-specified number of sub-periods, $l$, of length $T_j - T_{j-1}$ for $j = 1, \ldots, l$ and $0 = T_0 < T_1 < T_2 < \ldots < T_l = T$. Thus, the last observation in the $j$th sub-period is $T_j$, whereas $T_j + 1$ is the first observation in sub-period number $j + 1$.

for $t = T_{j-1} + k + 1, \ldots, T_j$ and $SD_{j,t} = 0$ otherwise. We can now reformulate (4) as an illustrative example as

$$\begin{aligned} \Delta y_t &= \Gamma_1 \Delta y_{t-1} + \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} y_{t-1} \\ tSD_t \end{pmatrix} + \mu SD_t \\ & \quad + \kappa_{2,1} ID_{2,t-1} + \kappa_{2,2} ID_{2,t-2} + \epsilon_t, t = k+1, \ldots, T \end{aligned} \tag{18}$$

where $SD_t = (SD_{1,t}, SD_{2,t})'$, $\gamma = (\gamma_1', \gamma_2')'$, $\mu = (\mu_1, \mu_2)$ and $\kappa_{2,1}$ and $\kappa_{2,2}$ are coefficient vectors of the indicator. Suppose that the sample period runs from the first quarter of 1984 to the fourth quarter of 2016 ($1984q1-2016q4$) and that the structural break occurs in the fourth quarter of 2008. Then, $SD_{1,t}$ is a step indicator which equals one in the period $1983q3-2008q3$, $SD_{2,t}$ is a step indicator which equals one in the period $2009q2-2016q4$ and $ID_{2,t}$ is an impulse indicator which equals one in $2008q4$ and $2009q1$, and otherwise equals zero.

The cointegration analysis in our example amounts to allowing $SD_t$ and $ID_{2,t}$ unrestricted entry to (18), whereas $tSD_t$ is restricted to lie in the cointegration space. We thereby allow for a structural break in the cointegration relationships. Again, as emphasised by Johansen *et al.* (2000), the asymptotic distribution of the trace test statistic depends *inter alia* on the number of break points, the location of the break points and the specification of the deterministic trend. Consequently, the trace statistic has to be compared in each single case with relevant critical values calculated by means of the response surface analysis in Johansen *et al.* (2000, Table 4).

Recently, Kurita and Nielsen (2019) proposed a class of partial systems in which structural breaks are allowed for in a similar manner as in the full system considered by Johansen *et al.* (2000).

## 1.2 Summary of evidence

This thesis builds on the concepts of the CVAR model outlined above, using mainly Norwegian macroeconomic time series data in the various applications of the model. In this subsection, I give a summary of the main empirical findings from each of the subsequent chapters, starting with Norwegian consumer behaviour (chapter 2) followed by Norwegian inflation dynamics (chapters 3 and 4), the effects of the distribution sector and trade liberalisation on exchange rate pass-through in Norway (chapters 5 and 6), export behaviour in Norway (chapter 7) and OPEC behaviour (chapter 8).

### 1.2.1 Chapter 2: The consumption Euler equation or the Keynesian consumption function?[13]

Whereas the Keynesian consumption function asserts that changes in household income affect consumption markedly, both the permanent income hypothesis of Friedman (1957) and the life-cycle hypothesis of Ando and Modigliani (1963) imply that consumption depends on unanticipated and not on anticipated income shocks, with a much stronger response to permanent than to transitory shocks. These hypotheses are typically formulated as consumption Euler equations, where consumption of a representative agent does not respond appreciably to transitory income changes. Consumption Euler equations in

[13]This chapter has been co-authored with Ådne Cappelen, Eilev Jansen and Anders Rygh Swensen and is forthcoming in *Oxford Bulletin of Economics and Statistics*; see Boug *et al.* (2020).

various forms have found little support in aggregate data, however; see Flavin (1981), Campbell and Deaton (1989), Muellbauer and Lattimore (1995), Yogo (2004), Palumbo *et al.* (2006) and Canzoneri *et al.* (2007).

Extended versions of the standard forward-looking theory that allow for precautionary savings, liquidity constraints and habit formation can explain some of the empirical results found in the literature. Campbell and Mankiw (1991) among others account for precautionary savings and liquidity constraints in an aggregate consumption model that assumes constant relative risk aversion utility preferences and that some of the households are current income consumers. Deaton (1991) explains consumer behaviour by means of the so-called buffer-stock model in which households facing liquidity constraints use liquid assets to buffer against temporary income shocks. Kaplan and Violante (2014) introduce trading costs to explain evidence of current income consumers even for those who are wealthy due to illiquid assets and credit constraints. The consumption model of Smets and Wouters (2003), upon which many DSGE models typically are based, includes habit formation in that current consumption is proportional to past consumption.

This chapter contributions to the literature in three ways. First, we formulate a CVAR model that nests both a class of consumption Euler equations and various Keynesian type consumption functions. The former include a version of the martingale hypothesis of Hall (1978) and the equations of precautionary savings and liquidity constraints as in Campbell and Mankiw (1991) and of habit formation as in Smets and Wouters (2003). Using likelihood methods, one can test the properties of *cointegration* between consumption and income and of *equilibrium correction* in the nesting CVAR. Drawing upon Eitrheim *et al.* (2002), the former property represents the common ground for a Keynesian type consumption function and a consumption Euler equation, while the latter represents the feature that distinguishes between them.

Second, we study aggregate Norwegian consumer behaviour within the context of the nested CVAR using quarterly data that span the period from the early 1980s to the end of 2016. We find support for cointegration between consumption, income and wealth once a structural break around the financial crisis in 2008 is taken into account. Our finding that consumption cointegrates with both income and wealth and not only with income is evidence against a consumption Euler equation. Likelihood ratio tests further show that consumption equilibrium corrects to changes in income and wealth and not that income equilibrium corrects to changes in consumption, as would be the case if an Euler equation were true.

Third, we consider conditional expectations of future consumption and income in CVAR models within the context of Johansen and Swensen (1999, 2008). Since, as pointed out by Tinsley (2002), "empirical rejection of rational expectations is the rule rather than the exception in macroeconomics", we divide the parameters of well fitted CVAR models into two parts: parameters of interest, which are the parameters describing rational expectations, and nuisance parameters, which are the parameters necessary to ensure satisfactory empirical fit. Using this strategy it is possible to focus on economically interesting parameters stemming from the class of Euler equations. Our treatment of the role of conditional expectations of future consumption and income is quite similar to what has been done in the new Keynesian literature on pricing behaviour; see Boug *et al.* (2010, 2017). We find that when conditional expectations of future consumption and income are considered in CVAR models, most of the parameters stemming from the class of Euler equations are not corroborated by the data. Only habit formation in line with Smets and Wouters (2003) seems to play an important role in explaining Norwegian consumer behaviour.

### 1.2.2 Chapter 3: Inflation dynamics in a small open economy[14]

Forward-looking models, based on rational expectations, optimizing agents, and imperfect competition in markets for goods, have long played a central role in the understanding of inflation dynamics in the economics profession and in central banks conducting inflation targeting. Since the influential papers by Roberts (1995), Fuhrer and Moore (1995), Galí and Gertler (1999), Galí *et al.* (2001), and Sbordone (2002), a large number of studies have been devoted to testing the role of expectations in inflation dynamics, using data from both closed and open economies, see Mavroeidis *et al.* (2014) and An and Schorfheide (2007) for reviews of the literature. Studies differ with respect to the sample period studied and the econometric methods applied. The supportive evidence for forward-looking behaviour in price formation is rather mixed.

In this chapter, we evaluate the empirical performance of forward-looking models based on Norwegian data that run from the first quarter of 1982 to the fourth quarter of 2011. Building on Sbordone (2002), our forward-looking models relate current inflation to expected future inflation and the difference between actual price and steady-state value of levels as a theory-consistent forcing variable. The steady-state value is specified as a mark-up over marginal costs, which, in turn, are determined by the costs of both labor and imported intermediate goods along the lines of the open economy models in McCallum and Nelson (1999), Kara and Nelson (2003), and Batini *et al.* (2005). We contribute to the empirical literature by studying both the exact formulation, in the sense of Hansen and Sargent (1991), and the inexact formulation, in which a stochastic error term is included in the model. For the exact formulation, we employ, as in chapter 2, the likelihood-based testing procedures suggested by Johansen and Swensen (1999, 2008) within the context of a CVAR model. Because similar treatment of the inexact formulation is more complicated to handle, see Boug *et al.* (2010), we rely on a test based on a minimum distance approach along the lines of Sbordone (2002) and Magnusson and Mavroeidis (2010). Consequently, we are able to shed some light on the importance of introducing a stochastic error term into the empirical model, an econometric issue that is often neglected in the literature. Unlike most related studies, see for instance Galí and Gertler (1999), Galí *et al.* (2001) and Batini *et al.* (2005), we pay particular attention to the time series properties of the variables involved, and the possible existence of unit roots. As a premise for the validity of the statistical inference of the forward-looking models, we search for statistically well-specified underlying VAR models. In our study, we also compare and contrast the specifications of a reduced-form forward-looking model with a backward-looking counterpart model as two competing models of inflation dynamics, both in-sample and out-of-sample in a forecasting competition.

Our empirical investigation produces several noteworthy findings. First, we establish a well-specified empirical counterpart to the theory-consistent link between consumer prices and marginal costs. Second, we demonstrate that the exact formulation of the forward-looking model is at odds with the data. The rational expectations hypothesis is not rejected statistically but, when only economically meaningful parameters are allowed,

[14]This chapter has been co-authored with Ådne Cappelen and Anders Rygh Swensen and is published in *The Scandinavian Journal of Economics*; see Boug *et al.* (2017).

the model is not supported by the data. In addition, a plot of the likelihood surface reveals that some of the parameters might not be well-identified. We also show that the inexact formulation of the forward-looking model does not yield radically different results. In particular, the identification problem seems also to be present in this model. Third, we establish a well-specified competing backward-looking model of inflation dynamics in a sample containing a major monetary policy regime shift. Finally, we find that the backward-looking model forecasts somewhat better than a reduced-form forward-looking model during and after the financial crisis in 2008.

### 1.2.3 Chapter 4: Expectations and regime robustness in price formation: evidence from vector autoregressive models and recursive methods[15]

While the usefulness of cointegration analysis for studying long run pricing behaviour is hardly controversial, the modelling of inflation dynamics is, as pointed out in chapter 3, more of an unsettled issue. Data-based methods are often used to develop conditional equilibrium correction models (EqCM), but theory-based models with forward-looking behaviour in order to explain short run dynamics are also widely applied in the literature, see Mavroeidis *et al.* (2014).

As explained by Roberts (1995), there are several routes from a theoretical set up of firm's pricing behaviour that lead to the new Keynesian Phillips curve, including the linear quadratic adjustment cost (LQAC) model of Rotemberg (1982) and the models of staggered contracts developed by Taylor (1979, 1980) and Calvo (1983). Proponents of these theory-based models often refer to the Lucas critique[16] as a reason for discarding conditional EqCM models and argue that the new Keynesian Phillips curve may circumvent the critique if economic agents are indeed forward-looking. However, estimation of such a model is not in itself compelling evidence for or against the Lucas critique, even if the proposed model is not rejected. One should also demonstrate the weaknesses of competing models and verify that those models are impacted by the Lucas critique. Conversely, rejection of a particular new Keynesian Phillips curve does not necessarily preclude economic agents from acting on alternative expectation-based models; see for instance Ericsson and Irons (1995).

The formation of export prices is an area in which the LQAC-model under rational expectations has been studied by Cuthbertson (1986, 1990) and Price (1991, 1992). Cuthbertson (1990) and Price (1991, 1992) conclude that the empirical performance of the model using UK manufacturing export price data compares favourably with an EqCM-model. However, not much evidence is given in these studies about the constancy or otherwise of the conditional and/or marginal models. One may then in accordance with Hendry (1988), see also Ericsson and Hendry (1999), argue that the LQAC-models and the EqCM-models reported in these studies are hard to distinguish empirically.

This chapter evaluates the empirical performance of both the LQAC-model and the EqCM-model, using Norwegian data on machinery export prices that span the period from the first quarter of 1978 to the fourth quarter of 2004. The Norwegian central bank

[15]This chapter has been co-authored with Ådne Cappelen and Anders Rygh Swensen and is published in *Empirical Economics*; see Boug *et al.* (2006).

[16]Lucas (1976) argues that the parameters of econometric models depend crucially on agents expectations and are unlikely to remain stable in the event of policy regime changes. See Favero and Hendry (1992) and Ericsson and Irons (1995) for comprehensive reviews and interpretations of the Lucas critique.

followed a policy of exchange rate targeting in various forms for many years before changing fundamentally to inflation targeting in 2001. In principle, this could have caused export price formation to shift in accordance with the Lucas critique. Our point of departure is the standard assumption of imperfectly competitive markets, where the optimal export price is determined as a mark-up over marginal costs. We estimate marginal costs using a fully specified cost-function consistent with economic theory. The mark-up in turn is modelled as depending on relative Norwegian and competing prices. To evaluate the LQAC-model, we again use the testing procedure of Johansen and Swensen (1999). The results show that there is overwhelming empirical evidence against the LQAC-model. We then follow Hendry (1988) and show by means of recursive methods that there exists a stable EqCM-model of export price behaviour (despite regime changes) along with an unstable marginal process for one of the conditioning variables in that model. Finally, we demonstrate that the estimated EqCM-model performs well post-sample in spite of the known monetary policy regime change with the introduction of inflation targeting. Hence, the invariance property of monetary policy regimes shows that the EqCM-model in our case is not subject to the Lucas critique.

### 1.2.4 Chapter 5: Exchange rate pass-through in a small open economy: the importance of the distribution sector[17]

Much of the literature on the new open economy macroeconomics is based on models that feature rational expectations, optimizing agents and imperfect competition in markets for goods. Small new Keynesian open economy models typically include these ingredients when analysing exchange rate pass-through, the responsiveness of import prices to changes in the exchange rate, and the effects of monetary policy; see for instance Svensson (2000), Galí and Monacelli (2005), Atkeson and Burstein (2008) and Bugamelli and Tedeschi (2008).

Some studies in this literature draw a distinction between producer currency pricing (PCP) and local currency pricing (LCP) when analysing exchange rate pass-through to domestic prices; see for instance Devereux and Engel (2003). According to PCP, prices for internationally traded goods are set in the currency of the producer (exporter). If PCP holds, producers do not change their prices frequently, whereas consumers (and importers) face prices that vary one-for-one with exchange rate changes (due to full pass-through). In this framework, changes in the exchange rate are passed on to the terms of trade and consumer demand for domestically relative to foreign produced goods. LCP, on the other hand, is a price setting strategy in which prices are set in the currency of the consumer, with no (or limited) pass-through of exchange rate changes to import prices, at least in the short run. Thus, exchange rate changes may have only limited effects on producer costs (to the extent that production is based on imported materials) or on consumer prices (to the extent that consumption is based directly on imported goods and services).

In this chapter, we present empirical evidence on exchange rate pass-through for the Norwegian economy by means of a CVAR model of trade margins in the distribution sector. The econometric modelling of trade margins is based on data that span the period from the first quarter of 1970 to the third quarter of 2010. The degree and speed of exchange rate pass-through to retailers' trade margins are important for inflation dynamics, as trade

[17]This chapter has been co-authored with Ådne Cappelen and Torbjørn Eika and is published in *Open Economies Review*; see Boug *et al.* (2013).

margins make up close to 30 per cent of the official consumer price index. We assume monopolistically competitive pricing behaviour when modelling prices, but do not consider forward-looking behaviour as this hypothesis is found to be clearly at odds with Norwegian data; see Bjørnstad and Nymoen (2008), Bårdsen *et al.* (2005, p. 145) and Boug *et al.* (2006). The estimated CVAR model of trade margins is then analysed within a large-scale macroeconometric model of the Norwegian economy, assuming 10 per cent depreciation of the Norwegian krone on a permanent basis. By using the macroeconometric model, which includes the pricing-to-market hypothesis of Krugman (1987), price-wage and wage-wage spirals across industries, we are able to examine exchange rate pass-through to import prices, production costs, mark-ups and consumer prices for a large number of commodities and industries. Unlike studies in the new open economy literature, which are typically based on partial analyses of aggregated single-equation models, we thus take into account numerous channels through which the exchange rate is likely to operate in a small, open economy like the Norwegian one.

Model simulations show that exchange rate pass-through takes place quite rapidly with import prices and fairly slowly with consumer prices in the Norwegian economy. We demonstrate that pass-through to consumer prices is not complete even within a ten-year horizon, a finding which may support the LCP hypothesis. The importance of the distribution sector is clearly apparent, as trade margins act as cushions to exchange rate fluctuations in the short run, thereby limiting the extent of exchange rate pass-through to consumer prices. If domestic inputs to the distribution sector are quantitatively important, then tradable goods sold to consumers include national value added (retail services), which may explain why there is incomplete pass-through. Likewise, imports as intermediate goods that together with domestic inputs produce final goods sold to consumers may also contribute to limited pass-through of exchange rate movements to consumer prices. We also present evidence that the exchange rate pass-through in the retailers' price setting has not changed significantly since the shift in monetary policy to inflation targeting in 2001 and the financial crisis in 2008.

### 1.2.5 Chapter 6: Trade liberalisation and exchange rate pass-through: the case of textiles and wearing apparels[18]

A key topic in monetary economics of interest for policymakers in general and inflation targeting central banks in particular is the responsiveness of prices of internationally traded goods to changes in exchange rates. Typically, empirical studies find evidence of incomplete pass-through, which is often explained by pricing-to-market behaviour under conditions of imperfect competition and segmented markets along the lines of Krugman (1987); see for instance Menon (1996), Naug and Nymoen (1996), Goldberg and Knetter (1997), Kenny and McGettigan (1998), Doyle (2004), Campa and Goldberg (2005) and Gust *et al.* (2010).

However, previous studies usually ignore the Bhagwati hypothesis that the presence of non-tariff barriers to trade may affect pass-through; see Bhagwati (1991). The hypothesis asserts that in the presence of quantity constraints on imports, a small depreciation of the exchange rate is likely to be absorbed into the quota rents extracted by the exporter rather than being reflected in import prices. If, on the other hand, the deprecia-

[18]This chapter has been co-authored with Andreas Benedictow and is published in *Empirical Economics*; see Benedictow and Boug (2013).

tion is large enough to push import prices above the point where the quantity constraints are no longer binding, then pass-through will be positive, but incomplete.

In this chapter, we estimate a CVAR model for Norwegian import prices on clothing using data from the first quarter of 1986 to the first quarter of 2008 and controlling for the gradual removal of non-tariff barriers to trade experienced in the clothing industry since the mid-1990s. The empirical analysis is prompted by the fact that low consumer price inflation observed over several years in Norway coincides well with a simultaneous fall in import prices for clothing. Developments in import prices for clothing may partly be explained by factors such as shifts in exchange rates, international prices (measured in foreign currency) and domestic market conditions in line with the pricing-to-market hypothesis of Krugman (1987). However, they should also be viewed in light of trade liberalisation, which led to a massive increase in imports of clothing from China and other low-cost countries at the expense of imports from high-cost countries, the euro area in particular. The significant deflationary effect on traded goods prices of shifts in the country composition of imports has been dubbed the China effect and is likely to be important for quantifying exchange rate pass-through in regression models.

We construct three different measures of foreign prices which are used in estimating pass-through. Our first two measures are based on the Törnqvist and Fischer price indices. The fact that available data on foreign prices for clothing are price indices and not price levels makes the Törnqvist and Fischer price indices not directly usable for numerical calculations in our context. If the available set of price indices is used directly by means of the Törnqvist and Fischer price indices, only inflationary impulses are taken into account in the final inflation aggregate. For constructing the first two measures of foreign prices, we therefore suggest a data calibration method based on purchasing power parities to take account of not only inflationary differences, but also varying import shares and differences in price levels, that is the China effect, among trading partners. Our third measure of foreign prices is based on the often used geometric mean price index with constant import shares as weights, a measure which thus fails to take account of the China effect. By comparing the pass-through estimates that emerge from modelling the import prices of clothing with the alternative measures of foreign prices, we are able to shed some light on the potential problem of omitted variable bias in the empirical tests of the pricing-to-market hypothesis.

We find that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that, on average, the shift in imports from high- to low-cost countries has reduced the international price impulses generated by clothing imports by around 2 percentage points annually. Controlling for these effects, we estimate the pass-through and pricing-to-market elasticities to be 0.44 and 0.56, respectively. By way of contrast, we find that using the geometric mean price index as a measure of foreign prices biases the estimates due to a substantial overestimation of international price impulses. We also establish that the preferred estimated dynamic model is reasonably stable within sample and exhibits no serious forecasting failures around the dates of the shifts in trade policy. That no serious forecasting failures are detected may reflect the fact that likely pass-through effects of changes in trade policy are controlled for through the suggested measures of foreign prices. Consequently, once the effect of shifts in imports towards low-cost countries is controlled for in the estimation, we find little evidence that the properties of the import price equation have changed alongside trade liberalisation.

After the publication of this chapter in *Empirical Economics* in 2013, Benedictow and Boug (2017) have suggested an alternative approach based on a geometric mean of *price levels* in order to construct a measure of foreign prices. Although the calculated China effect will become qualitatively similar to the one in this chapter based on the Törnqvist and Fischer price indices, Benedictow and Boug (2017) show how to *exactly* decomposing a geometric mean of price levels into inflationary effects and the China effect. In this sense, the approach in Benedictow and Boug (2017) may be considered as an improvement of how the measures of foreign prices are calculated in this chapter.

### 1.2.6 Chapter 7: Exchange rate volatility and export performance: a cointegrated VAR approach[19]

Since the breakdown of the Bretton-Woods agreement and the transition to floating exchange rates, the nature and magnitude of the relationship between exchange rate volatility and trade flows has been a subject of major interest to economists. A number of theoretical models exist showing that the impact of exchange rate volatility on trade may be positive or negative depending on the assumptions made with respect to risk preferences, the availability of (forward) capital markets and the time horizon of trade transactions; see for instance Ethier (1973), Hooper and Kohlhagen (1978), De Grauwe (1988), Franke (1991), Viaene and de Vries (1992) and Sercu (1992).

The empirical evidence is no less inconclusive. Some studies, such as Chowdhury (1993), Arize (1995), Arize *et al.* (2000) and De Vita and Abbott (2004), provide evidence that increased exchange rate volatility has an adverse effect on trade due to risk-averse traders. In other words, higher exchange rate volatility leads to higher costs for risk-averse traders and thus to less trade volume. On the other hand, Asseery and Peel (1991), Holly (1995) and Bredin *et al.* (2003) are among those who find that exchange rate volatility affects trade positively. When trade is considered as an option held by firms, the trade option value, and hence also the export supply, may rise with exchange rate volatility. Others find no evidence that exchange rate volatility has any significant impact on trade; see for instance Aristotelous (2001) and Solakoglu *et al.* (2008). Given today's financial markets for currency hedging, one may argue that traders should be able (at least to some extent) to reduce or hedge uncertainty associated with exchange rate volatility. The relationship between exchange rate volatility and trade may therefore be weak, if not completely absent.

McKenzie (1999) and Bahmani-Oskooee and Hegerty (2007) provide literature reviews and discuss several empirical issues that may be important for determining the impact of exchange rate volatility on trade. These issues are mainly related to which exchange rate volatility measure to use, which sample period to consider, which countries to study, which data frequency and aggregation level to employ and which methodology to apply in each specific study at hand. Any of these issues and the manner in which they are handled may provide part of the explanation for the inconclusive findings in the literature.

In this chapter, we aim to provide further evidence of the impact of exchange rate volatility on exports while trying to take into account some of the issues related to previous contributions. Specifically, we study exchange rate volatility and Norwegian exports within a standard demand-type model. Knowledge of the impact of exchange rate

[19]This chapter has been co-authored with Andreas Fagereng and is published in *Applied Economics*; see Boug and Fagereng (2010).

volatility on exports is of vital importance for policymakers in a small open economy like the Norwegian one, which depends heavily on trade with the outside world. The empirical analysis is based on the CVAR model and data that span the period from the first quarter of 1985 to the fourth quarter of 2005. The CVAR model is particularly relevant in our context as different properties of the time series involved, which are often neglected in related studies, can be handled using essentially the same methodology.

We follow Johansen (1995) and Rahbek and Mosconi (1999) and conduct cointegration rank inference by means of a VAR model in which all variables, including exchange rate volatility, are non-stationary, and a VAR model in which volatility is a stationary regressor, respectively. To measure volatility, we make use of the conditional variance of the exchange rate from a Generalized Autoregressive Conditional Heteroscedasticity (GARCH) model. As a test of robustness, we consider GARCH-measures of volatility based on both the nominal and the real exchange rate. Unlike some related studies, which have used aggregated data for an economy, we model exports of machinery and equipment as one important group of goods. Finally, we pay attention to special exchange rate events and monetary policy regime shifts during the sample period. In particular, we test by means of out-of-sample forecasting whether the shift to inflation targeting in 2001 did have significant effects on exporters' behaviour, and thus on the estimated coefficients of a parsimonious model.

Our empirical findings suggest that a reduced rank VAR, in which exports, relative prices and world market demand represent the modelled variables, explains the data quite well. However, we are unable to identify any statistically significant cointegration relationship among the variables when the information set also includes a GARCH-measure of exchange rate volatility, treated as either a stationary or a non-stationary variable in the VAR. Rather, we find that volatility changes by means of impulse dummies connected to the monetary policy change from a fixed to a managed floating exchange rate and the Asian financial crises during the 1990s enter significantly into a dynamic model in which exports, relative prices and world market demand form a cointegration relationship. We also demonstrate that the dynamic model performs well out-of-sample, a finding which contradicts the hypothesis that increased exchange rate volatility following the introduction of inflation targeting has had a significant impact on export performance.

Since the publication of this chapter in *Applied Economics* in 2010, some investigators have followed up the analysis of exchange rate volatility on trade flows for other countries with some similar modelling features. For instance, Sharma and Pal (2018) find that nominal exchange rate volatility, measured by means of GARCH-type models, has a dampening long run impact on India's commodity exports to the US, Germany and China within a panel data setting. Pino *et al.* (2016) also find a negative long run relationship between different measures of exchange rate volatility and export flows of Indonesia, Malaysia, South Korea, Thailand and the Philippines within a univariate cointegration approach.

### 1.2.7 Chapter 8: Did OPEC change its behaviour after the November 2014 meeting?[20]

The growing importance of OPEC during the 1970s and the increased perception that the organisation could affect world oil prices initiated a lot of empirical work about the oil

[20] This chapter is submitted to *Empirical Economics*; see Boug (2020).

market, and in particular, about the behaviour of OPEC. Yet, there is no clear consensus in the literature on the exact nature of the behaviour of OPEC and its ability to influence world oil prices. Some studies rely on a competitive model while others specify models of imperfect competition, often assuming that OPEC functions like a monopoly, an oligopoly or a cartel in some way; see for example Smith (2009), Almoguera *et al.* (2011), Fattouh and Mahadeva (2013) and Alkhathlan *et al.* (2014) for reviews of the literature.

The study of Griffin (1985), which *inter alia* tests the various cartel models and the competitive model for the period 1971 to 1983, is the starting point for much of the empirical work in this field of research. Among the various competing hypotheses about OPEC behaviour, Griffin (1985) finds some support for a cartel model where OPEC as a group is a dominant producer setting the oil price, while non-OPEC countries behave as a competitive fringe. A similar conclusion is reached in Böckem (2004), relying on data for the 1990s. Other studies find that a core group within OPEC, or even Saudi Arabia alone, fits the description of a dominant producer; see for example Dahl and Yücel (1991), Alhajji and Huettner (2000a, 2000b) and Hansen and Lindholt (2008). Another study by Almoguera *et al.* (2011) concludes that the behaviour of OPEC over the period from 1974 to 2004 is best described as a Cournot competition with a competitive fringe. Recently, Golombek *et al.* (2018) find support for a dominant firm-competitive fringe model, in which both the demand and the supply side of the oil market are taken into account, using data over the period from 1986 to 2016. Smith (2005) argues for a model of the OPEC countries as a "bureaucratic production syndicate" and finds strong evidence of collusion, but with significant transaction costs regarding redistribution of output among members of the cartel. By way of contrast, Spilimbergo (2001) finds no support for the hypothesis that OPEC is a market sharing cartel and Okullo and Reynès (2016) show that heterogeneity within OPEC and non-OPEC supply create strong incentives against collusion.

The empirical evidence on OPEC behaviour is thus rather mixed. Kaufmann *et al.* (2008) argue that the failure to agree on a benchmark model is not a weakness, but rather a strength, as OPEC behaviour does not fit easily into any single model. The disagreement and plethora of models may be due to changing OPEC behaviour over time, as pointed out by Almoguera *et al.* (2011), Fattouh and Mahadeva (2013) and Alkhathlan *et al.* (2014). Several oil market analysts have also argued that at the November 2014 meeting OPEC significantly changed its behaviour by deciding to keep its supply unchanged despite the huge oil price drop in advance. Recently, Behar and Ritz (2016) argue that this decision was taken by OPEC mainly to limit the role of competitors like American producers of shale oil.

In this chapter, I contribute to the literature by studying the behaviour of OPEC as a group for the period from 1992 to 2018, with a special focus on the possibility of changes in actions taken by the organisation before and after the November 2014 meeting. First, I set up a theoretical model of OPEC behaviour that encompasses several of the possibilities discussed in the literature, among them the competitive model, various forms of the imperfect competition model and the Hotelling rule. Unlike some related empirical studies, my model includes a possible role for resource stock to affect OPEC extraction or production costs. Then, in the empirical analysis, I contribute to the literature by paying particular attention to time series properties of variables involved and searching for statistically well-specified underlying models as premises for valid inference about the OPEC behaviour.

By estimating a CVAR model based on data running from the first quarter of 1992 to the fourth quarter of 2013, I find support for the imperfect competition hypothesis regarding the output decisions of OPEC. The implied average estimate of the price elasticity of demand for OPEC oil is less than minus unity, consistent with the dominant producer model. Moreover, OPEC's oil reserves seem to affect production costs positively, contradicting Hotelling's (1931) rule. I also find that a dynamic equilibrium correction model with imperfect competition is reasonably stable in-sample and has a somewhat better fit than an alternative dynamic model with weaker theoretical underpinnings. However, a forecasting exercise based on data running from the first quarter of 2014 to the fourth quarter of 2018 reveals that the preferred model breaks down following the OPEC meeting in November 2014. The model systematically underpredicts OPEC production over the most part of the forecasting period and by as much as nearly 3 million barrels per day at the end of 2016. A sequence of parameter constancy forecast Chow tests also reveals a significant structural break around the fourth quarter of 2014. During 2018, however, the model forecasts OPEC production quite well. I therefore conclude that OPEC's behaviour did indeed change significantly after the November 2014 meeting.

## 1.3 Concluding remarks

The purpose of this thesis has been to use the CVAR model to evaluate the empirical performance of theoretical models of consumer behaviour, inflation dynamics, exchange rate pass-through, export behaviour and OPEC behaviour.

As established in chapter 2, the empirical findings lend support for Keynesian type consumption functions rather than consumption Euler equations to explain Norwegian consumer behaviour. The results from chapters 3 and 4 show that different types of forward-looking models are at odds with Norwegian inflation dynamics, and that competing dynamic backward-looking models fit the data pretty well. The outcome from chapter 5 is that the exchange rate pass-through is quite rapid with respect to import prices and fairly slow with respect to consumer prices in the Norwegian economy, supporting the local currency hypothesis. The importance of the distribution sector is clearly demonstrated, as trade margins act as cushions to exchange rate fluctuations in the short run, thereby limiting the extent of exchange rate pass-through to consumer prices. The findings from chapter 6 point to some degree of incomplete exchange rate pass-through in Norwegian import prices for clothing in accordance with both the pricing-to-market and the Bhagwati hypotheses. The upshot of chapter 7 is that exchange rate volatility, measured by means of a GARCH model, is not able to explain export behaviour in Norway. By way of contrast, the results show that volatility, measured by means of impulse dummies related to monetary policy changes and the Asian financial crises during the 1990s, plays a significant part in a demand type model of exports. Finally, the findings from chapter 8 support a dynamic imperfect competition model of OPEC oil supply. A forecasting exercise reveals, however, that the model breaks down following the November 2014 meeting at which OPEC, surprisingly enough, decided to keep its supply unchanged despite the huge oil price drop in advance.

Although this thesis has evaluated basic assumptions underlying the various theoretical models by means of testable hypotheses within fairly well-specified CVAR models, some unanswered questions still arise from the empirical analyses. One unanswered methodological question is related to the examination of parameter constancy in the estimated models. Most of the chapters have relied on recursive estimation of coefficients and sequential test statistics for examining instabilities within the sample period. While these are useful techniques for establishing parameter constancy, there are now more powerful approaches available, such as the so-called impulse indicator saturation tests; see for instance Hendry and Doornik (2014). In particular, impulse indicator saturation tests have the advantage of being able to assess parameter constancy early in the sample period, whereas recursive methods require an initial estimation period. Hence, this thesis fails to evaluate parameter constancy for about a fifth of the sample period in many cases. Another unanswered methodological question is related to the failure to impose a "true" weakly exogenous variable when testing conditional expectations within the context of Johansen and Swensen (1999, 2008). For instance, in chapter 2, household income is found to be weakly exogenous with respect to the long run parameters, a property which is not imposed in the testing of conditional expectations of future consumption and income, as the testing procedure of Johansen and Swensen (1999, 2008) requires a full system. The question of whether efficiency gains could be achieved in the finite sample by testing conditional expectations within a partial rather than within a full system is thus left unanswered. The question of whether weak exogeneity of income has implications for the analysis of the structural break around the financial crisis in 2008, which could be illuminated by the framework of Kurita and Nielsen (2019), is also not answered in this thesis.

Yet, some other unanswered questions are related to the operationalisation of theoretical variables and to the underlying assumptions imposed on the theoretical models from the outset. For instance, in chapter 2 the consumption variable includes durable goods in addition to non-durable goods and services. Because the implications of the permanent income hypothesis are tested, one may argue that the empirical findings are to some extent influenced by the investment element of durable goods, and should therefore be excluded from the consumption variable under study. Similarly, in chapter 3 the deflator for total imports is used as a measure of the unit import costs of intermediate goods used in production. Since total imports also include final consumer goods, and not just intermediate goods, one may argue that the correspondence between the operational variable and the theoretical variable is not as close as it should be. When it comes to assumptions imposed on the theoretical models, chapter 2 imposes the *ex post* real interest rate as a simplifying assumption in the model of Campbell and Mankiw (1991). Thus this thesis fails to test for conditional expectations regarding the *ex ante* real interest rate in addition to conditional expectations regarding future consumption and income. Another simplifying assumption is made in chapter 8, in that the theoretical model for estimation purposes is basically related to the supply behaviour of OPEC. While important, the demand side of the oil market, that is the residual demand facing OPEC, is not an explicit part of the empirical model. The question of whether these untested simplifying assumptions are crucial to the conclusions of the empirical analyses, and thus to their reliability, is left for future research. Thus, the so-called *Law of Decreasing Credibility* by Charles F. Manski may put some of the conclusions in this thesis, if not all of them, in the right perspective:

"*There is a tension between the strength of assumptions and their credibility. I have called this...: The Law of Decreasing Credibility: The credibility of inference decreases with the strength of the assumptions maintained. This Law implies that analysts face a dilemma as they decide what assumptions to maintain: stronger assumptions yield conclusions that are more powerful but less credible.*"
— Manski (2011, p. 262)

## 1.4 References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# 2 Chapter 2

## The consumption Euler equation or the Keynesian consumption function?

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# The consumption Euler equation or the Keynesian consumption function?

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**Abstract**

We formulate a general cointegrated vector autoregressive (CVAR) model that nests both a class of consumption Euler equations and various Keynesian type consumption functions. Using likelihood-based methods and Norwegian data, we find support for cointegration between consumption, income and wealth once a structural break around the financial crisis is taken into account. That consumption cointegrates with both income and wealth and not only with income points to the empirical irrelevance of an Euler equation. Moreover, we find that consumption equilibrium corrects to changes in income and wealth and not that income equilibrium corrects to changes in consumption, which would follow from an Euler equation. We also find that when conditional expectations of future consumption and income are considered in CVAR models, most of the parameters stemming from the class of Euler equations are not corroborated by the data. Only habit formation seems important in explaining Norwegian consumer behaviour. Our estimated conditional Keynesian type consumption function implies a first year marginal propensity to consume (MPC) out of income close to 40 per cent.

**Keywords:** Consumption Euler equation, Keynesian consumption function, financial crisis, structural break, conditional expectations

**JEL classification:** C51, C52, E21

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# 1 Introduction

Economists have long been concerned with how households react to changes in fiscal policy. The financial crisis in 2008 led to renewed interest in how household asset composition, liquidity and credit market conditions may affect consumption; see for instance Muellbauer (2016) and Kaplan *et al.* (2018). The effects of fiscal policy depend on the marginal propensity to consume (MPC) out of shocks to income. A new consensus seems to be emerging on the size of the MPC that is much larger than what used to be common in many DSGE models. For instance, the heterogeneity-augmented model by Carroll *et al.* (2017) predicts an aggregate MPC of around 20 per cent compared to roughly 5 per cent implied by macroeconomic models with representative agents.

Whereas the Keynesian consumption function asserts that changes in household income affect consumption markedly, both the permanent income hypothesis of Friedman (1957) and the life-cycle hypothesis of Ando and Modigliani (1963) imply that consumption depends on unanticipated and not on anticipated income shocks, with a much stronger response to permanent than to transitory shocks. These hypotheses are typically formulated as consumption Euler equations, where consumption of a representative agent does not respond appreciably to transitory income changes. Consumption Euler equations in various forms have found little support in aggregate data, however; see Flavin (1981), Campbell and Deaton (1989), Muellbauer and Lattimore (1995), Yogo (2004), Palumbo *et al.* (2006) and Canzoneri *et al.* (2007). Recent microeconometric studies also find that households react much more strongly to transitory income shocks than predicted by the standard forward-looking theory of consumption. For instance, Jappelli and Pistaferri (2014) estimate an average MPC of 48 per cent using Italian data and Fagereng *et al.* (2019) find an MPC that ranges between 35 and 70 per cent using Norwegian data.

Extended versions of the standard forward-looking theory that allow for precautionary savings, liquidity constraints and habit formation can explain some of the empirical results found in the literature. Campbell and Mankiw (1991) among others account for precautionary savings and liquidity constraints in an aggregate consumption model that assumes constant relative risk aversion utility preferences and that some of the households are current income consumers. Deaton (1991) explains consumer behaviour by means of the so-called buffer-stock model in which households facing liquidity constraints use liquid assets to buffer against temporary income shocks. Kaplan and Violante (2014) introduce trading costs to explain evidence of current income consumers even for those who are wealthy due to illiquid assets and credit constraints. The consumption model of Smets and Wouters (2003), upon which many DSGE models typically are based, includes habit formation in that current consumption is proportional to past consumption.

The contributions of the present paper are threefold. First, we formulate a general cointegrated vector autoregressive (CVAR) model that nests both a class of consumption Euler equations and various Keynesian type consumption functions.

The former include a version of the martingale hypothesis of Hall (1978) and the equations of precautionary savings and liquidity constraints as in Campbell and Mankiw (1991) and of habit formation as in Smets and Wouters (2003). Using likelihood methods, one can test the properties of *cointegration* between consumption and income and of *equilibrium correction* in the nesting CVAR. Drawing upon Eitrheim *et al.* (2002), see also Anundsen and Nymoen (2019) for an application to American data, the former property represents the common ground for a Keynesian type consumption function and a consumption Euler equation while the latter represents the feature that distinguishes between them. The joint implication of a consumption Euler equation and the existence of cointegration between consumption and income is that saving today predicts income decline tomorrow, the so-called "saving for a rainy day" hypothesis of Campbell (1987).

Second, we study aggregate Norwegian consumer behaviour within the context of the nested CVAR using seasonally unadjusted quarterly data that span the period from the early 1980s to the end of 2016. We find support for cointegration between consumption, income and wealth once a structural break around the financial crisis in 2008 is taken into account. Our finding that consumption cointegrates with both income and wealth and not only with income is evidence against a consumption Euler equation. Likelihood ratio tests further show that consumption equilibrium corrects to changes in income and wealth and not that income equilibrium corrects to changes in consumption, as would be the case if an Euler equation were true. Our estimated conditional Keynesian type consumption function implies a first year MPC of around 40 per cent, which is in line with recent microeconometric evidence.

Third, we consider conditional expectations of future consumption and income in CVAR models within the context of Johansen and Swensen (1999, 2008). Since, as pointed out by Tinsley (2002), "empirical rejection of rational expectations is the rule rather than the exception in macroeconomics", we divide the parameters of well fitted CVAR models into two parts: parameters of interest, which are the parameters describing rational expectations, and nuisance parameters, which are the parameters necessary to ensure satisfactory empirical fit. Using this strategy it is possible to focus on economically interesting parameters stemming from the class of Euler equations. Our treatment of the role of conditional expectations of future consumption and income is quite similar to what has been done in the new Keynesian literature on pricing behaviour; see Boug *et al.* (2010, 2017). We find that when conditional expectations of future consumption and income are considered in CVAR models, most of the parameters stemming from the class of Euler equations are not corroborated by the data. Only habit formation in line with Smets and Wouters (2003) seems to play an important role in explaining Norwegian consumer behaviour.

The rest of the paper is structured as follows: Section 2 discusses the theoretical background and how the various hypotheses of consumer behaviour are nested within a general CVAR. Section 3 presents the data used in the empirical analysis. Section 4 reports findings from the cointegration analysis and the estimation of consumption short run dynamics. Section 5 presents results from considering conditional expectations of consumption and income in CVAR models. Section 6 provides a conclusion.

# 2 Theoretical background

As a useful benchmark for the empirical analysis, we first outline the martingale hypothesis and the "saving for a rainy day" hypothesis. Then, we present the consumption Euler equations with precautionary savings, liquidity constraints and habit formation based on CRRA utility preferences. Finally, we formulate a CVAR that nests the various hypotheses from the set of Euler equations as well as Keynesian type consumption functions.

## 2.1 The martingale and the "saving for a rainy day" hypotheses

The main idea behind the permanent income hypothesis and the life-cycle hypothesis, which both the martingale and the "saving for a rainy day" hypotheses build upon, is that aggregate consumption can be modelled as the intertemporal optimisation decision under uncertainty by a representative consumer.

If the utility function of the consumer is quadratic and the riskless rate of real return is constant and equal to the subjective discount rate, the martingale hypothesis by Hall (1978) can be formulated as

$$E_t C_{t+1} = C_t, \tag{1}$$

where $E_t$ and $C_{t+1}$ denote expectations conditional on information at time $t$ and consumption at time $t+1$, respectively. According to (1), no other variable than consumption at time $t$ should help predict consumption at time $t+1$. This in turn implies that $\Delta C_{t+1} = \varepsilon_{t+1}$, where $\varepsilon_{t+1}$ is an unforecastable innovation in permanent income. The change in consumption is thus unforecastable.

An alternative formulation of the permanent income hypothesis is the "saving for a rainy day" hypothesis by Campbell (1987). As shown by Campbell and Deaton (1989) and used by Palumbo *et al.* (2006) among others, a logarithmic version of this hypothesis can be written as $\frac{S_t}{YL_t} \approx -\sum_{i=1}^{\infty} \rho^i E_t \Delta yl_{t+i} + \varsigma$, where $S_t \equiv Y_t - C_t$, $Y_t \equiv R(1+R)^{-1}W_t + YL_t$, $R$ denotes the riskless rate of real return, $W_t$ is financial wealth at time $t$, $YL_t$ is labour income at time $t$, $\rho$ denote a discount factor and $\varsigma$ is a constant.[1] Hence, the saving ratio, $\frac{S_t}{YL_t}$, and the expected future labour income growth, $E_t \Delta yl_{t+i}$, are negatively related so that savings are increasing today when the consumer anticipates income to decline tomorrow. The consumer "saves for a rainy day".

We follow Eitrheim *et al.* (2002) and Anundsen and Nymoen (2019) and

[1] Here and below lower case letters denote the logarithms of a variable.

formulate the "saving for a rainy day" hypothesis as

$$y_t - c_t \approx -\sum_{i=1}^{\infty} \rho^i E_t \Delta y_{t+i} + \varsigma, \tag{2}$$

where the saving ratio, $\frac{S_t}{YL_t}$, is approximated by the logarithms of income to consumption ratio, $y_t - c_t$, and labour income, $yl_{t+i}$, is replaced by income, $y_{t+i}$. An important time series property, which we shall utilise in the nesting CVAR, is that the saving ratio is stationary, $I(0)$, and thus that income and consumption are cointegrated with a coefficient equal to one, when income is non-stationary, $I(1)$.

## 2.2 Euler equations with CRRA preferences

To allow for precautionary savings, liquidity constraints and habit formation, we now turn to consumption Euler equations with CRRA preferences. Campbell and Mankiw (1991) apply the utility function $U(C_t) = (1-\delta)^{-1} C_t^{1-\delta}$ for $1 \neq \delta > 0$, where $\delta$ is the inverse of the intertemporal elasticity of substitution, $\sigma$. Assuming that the logarithms of consumption is normally distributed with mean $E_t \ln C_{t+1}$ and time varying variance $\eta_{t+1}^2$, we may write the consumption Euler equation as

$$E_t \Delta c_{t+1} = \frac{\eta_{t+1}^2}{2\sigma} - \sigma\theta + \sigma R_t, \tag{3}$$

where $\theta$ and $R_t$ denote the subjective discount rate, assumed constant, and the real interest rate, assumed ex post, respectively. If the consumer faces more uncertainty, that is a larger $\eta_{t+1}^2$, consumption is expected to increase from this period to the next. Thus, the consumer reduces consumption now in response to increased uncertainty to have a larger safety buffer, that is precautionary savings, for more consumption in the next period. According to (3), savings by the consumer is also associated with intertemporal substitution in consumption. An increase in the real interest rate makes savings more profitable due to relatively costly consumption today, hence consumption is expected to increase from this period to the next. When the variance, $\eta^2$, is constant, (3) simplifies to $E_t \Delta c_{t+1} = \phi + \sigma R_t$, where $\phi = \frac{\eta^2}{2\sigma} - \sigma\theta$.

Campbell and Mankiw (1991) account for liquidity constraints in a simple way by assuming that aggregate consumption is equal to a weighted average of rule of thumb consumers and permanent income consumers with weights $\mu$ and $1-\mu$, respectively. In addition, Campbell and Mankiw (1991) assume that rule of thumb consumers determine consumption growth as a weighted average of current and one period lag of income growth with weights $\lambda$ and $1-\lambda$. We can then formulate an augmented version of (3) with constant variance as

$$E_t \Delta c_{t+1} = (1-\mu)\phi + \mu[\lambda E_t \Delta y_{t+1} + (1-\lambda)\Delta y_t] + (1-\mu)\sigma R_t, \tag{4}$$

where $\Delta y_{t+1}$ and $\Delta y_t$ are disposable income growth at time $t+1$ and $t$. As stressed by Basu and Kimball (2002) and later by Galí *et al.* (2007), the interpretation of the results in Campbell and Mankiw (1991) hinges on the assumption of utility preferences that are separable in consumption and labour. Otherwise, due to high correlation between changes in disposable income and hours worked, a significant $\mu$ may be the outcome from estimation of (4) even if all consumers are fully permanent income consumers.

We may also formulate an augmented version of (4) by adding lagged change in consumption, $\Delta c_t$, and an equilibrium correction term, $(c_t - \nu y_t)$, such that

$$(5)\quad E_t\Delta c_{t+1} = (1-\mu)\phi+\mu[\lambda E_t\Delta y_{t+1}+(1-\lambda)\Delta y_t]+(1-\mu)\sigma R_t+\tau\Delta c_t+\varrho(c_t-\nu y_t),$$

where consumption and income are cointegrated with the parameter $\nu$. As pointed out by Campbell and Mankiw (1991), $\Delta c_t$ would appear in (5) with $\tau > 0$ if there are important quadratic adjustment costs in consumption whereas $(c_t - \nu y_t)$ would appear with $\varrho < 0$ in a disequilibrium model of consumption and income.[2]

The consumption Euler equation by Smets and Wouters (2003), typically included in DSGE models, is also based on CRRA preferences. However, the utility function now also includes habit formation, $hC_{t-1}$, that is proportional to past consumption through the parameter $h$.[3] In line with Smets and Wouters (2003), we may log-linearize the Euler equation around a non-stochastic steady state such that consumption obeys $c_t = (1 - \omega_1)c_{t-1} + \omega_1 E_t c_{t+1} - \omega_2 \hat{r}_t$, where $\omega_1 = (1 + h)^{-1}$, $\omega_2 = \frac{(1-h)}{(1+h)\delta}$ and $\hat{r}_t$ is the log deviation of the ex ante real interest rate from its non-stochastic steady state. Consumption thus depends on a weighted average of past and expected future consumption and the ex ante real interest rate. The higher the degree of habit formation, the smaller is the impact of the real interest rate on consumption for a given elasticity of substitution. We can also write expected consumption growth when $h \neq 0$ as

$$(6)\qquad E_t\Delta c_{t+1} = h\Delta c_t + (1 - h)\delta^{-1}\hat{r}_t.$$

Hence, effects on expected consumption growth of lagged change in consumption can either be attributed to habit formation, as in (6), or to quadratic adjustment costs in consumption, as in (5).

## 2.3 A nesting CVAR

Thus far we have focused on various consumption models based on Euler equations. There exists, however, a huge empirical literature initiated by Davidson *et al.* (1978) based on a theoretical framework, which goes back to Keynes (1936), saying that current aggregate income is an important determinant of current aggregate consumption. The consumption models by Brodin and Nymoen (1992), Eitrheim *et*

[2]Campbell and Mankiw (1991) impose $\nu = 1$ and find both factors to be insignificant for a number of countries. However, for the UK the rejection of the equilibrium correction term is contested by Hendry (1991).

[3]We simplify matters by disregarding shocks to the subjective discount rate which is another element entering the utility function in Smets and Wouters (2003).

*al.* (2002), Erlandsen and Nymoen (2008) and Jansen (2013), which are all based on Norwegian data, belong to this literature. These studies have in common a Keynesian type long run consumption function of the form

$$c_t = \beta_y y_t + \beta_w w_t, \tag{7}$$

where $c_t$, $y_t$ and $w_t$ denote real consumption, real disposable income and real household wealth, respectively. If $c_t$, $y_t$ and $w_t$ are integrated series of order one, I(1), (7) implies cointegration between the three variables with the cointegration parameters $\beta_y$ and $\beta_w$ for income and wealth. Both Erlandsen and Nymoen (2008) and Jansen (2013) augment (7) by the real after tax interest rate as a separate variable to capture the possibility of long run substitution effects in consumption. An increase in the real after tax interest rate is assumed to make consumption today more expensive relative to consumption tomorrow. Hence, consumption is expected to decline. Noticeably, (7) and the "saving for a rainy day" hypothesis in (2) share the same cointegration property between consumption and income in the special case when $\beta_y = 1$ and $\beta_w = 0$.

We now formulate a general CVAR that nests all the Euler equations considered above as well as the various Keynesian type consumption functions inherent in (7). Our point of reference is a *full* CVAR representation of a $p$-dimensional VAR of order $k$ written as

$$\Delta X_t = \Pi X_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \Delta X_{t-j} + \gamma t + \vartheta + \Phi D_t + \epsilon_t, \tag{8}$$

where $\Delta$ is the difference operator, $X_t = (c_t, y_t, w_t, R_t)'$ comprises real consumption, $c_t$, real disposable income, $y_t$, real household wealth, $w_t$, and the real after tax interest rate, $R_t$, as the modelled variables, $t$ is a deterministic trend, $\Gamma_j$ and $\Phi$ are matrices of coefficients, $\gamma$ is a vector of coefficients, $\vartheta$ is a vector of intercepts, $D_t$ is a vector of centered seasonal dummies, and $\epsilon_t$ are normally distributed random variables with expectation zero and unrestricted covariance matrix $\Omega$. The initial observations $X_1, \ldots, X_k$ are considered as given. The impact matrix $\Pi$ has rank $0 \leq r \leq p$, and therefore can be written $\Pi = \alpha\beta'$, where $\alpha$ and $\beta$ are $p \times r$ matrices of adjustment coefficients and cointegration coefficients, respectively, of full rank $r$.

Drawing upon the analysis in Eitrheim *et al.* (2002), the Euler equation approach implies that consumption, wealth and the real after tax interest rate are *not* equilibrium correcting and that income alone, in line with the "saving for a rainy day" hypothesis in (2), *is* equilibrium correcting. These properties and the various hypotheses considered in Subsections 2.1 and 2.2 are nested in the CVAR only when $r = 2$. By leading (8) one period and taking the conditional expectations of $\Delta X_{t+1}$, we can write out the CVAR when $k = 2$ for notational simplicity as

$$
(9) \qquad E_t \begin{pmatrix} \Delta c_{t+1} \\ \Delta y_{t+1} \\ \Delta w_{t+1} \\ \Delta R_{t+1} \end{pmatrix} = \begin{pmatrix} \alpha_{c1} & \alpha_{c2} \\ \alpha_{y1} & \alpha_{y2} \\ \alpha_{w1} & \alpha_{w2} \\ \alpha_{R1} & \alpha_{R2} \end{pmatrix} \begin{pmatrix} 1 & \beta_{y1} & 0 & \beta_{R1} \\ -1 & 1 & \beta_{w2} & \beta_{R2} \end{pmatrix} \begin{pmatrix} c_t \\ y_t \\ w_t \\ R_t \end{pmatrix}
$$

$$
+ \begin{pmatrix} \gamma_{1,11} & \gamma_{1,12} & \gamma_{1,13} & \gamma_{1,14} \\ \gamma_{1,21} & \gamma_{1,22} & \gamma_{1,23} & \gamma_{1,24} \\ \gamma_{1,31} & \gamma_{1,32} & \gamma_{1,33} & \gamma_{1,34} \\ \gamma_{1,41} & \gamma_{1,42} & \gamma_{1,43} & \gamma_{1,44} \end{pmatrix} \begin{pmatrix} \Delta c_t \\ \Delta y_t \\ \Delta w_t \\ \Delta R_t \end{pmatrix}
$$

$$
+\gamma t + \vartheta + \Phi D_{t+1},
$$

where $E_t \epsilon_{t+1} = 0$ and $\beta_{y1} = -\nu$ from (5). Exact identification of the two cointegrating vectors is achieved by imposing $\beta_{c1} = 1$ and $\beta_{w1} = 0$ in the first row of $\beta'$ and $\beta_{c2} = -1$ and $\beta_{y2} = 1$ in the second row of $\beta'$, all dictated from the theory of cointegration between consumption and income. The consumption Euler equation and the "saving for a rainy day"hypothesis together impose $\beta_{y1} = -1$ and $\beta_{w2} = \alpha_{w1} = \alpha_{w2} = \alpha_{R1} = \alpha_{R2} = 0$ as additional restrictions on the cointegrating part of (9), which makes the two cointegrating vectors not identifiable. Still (9) provides important insights by deriving some of the single equation relationships in Subsection 2.2 from it. In particular, it is only when $\alpha_{c1} = \alpha_{c2}$ that consumption is *not* equilibrium correcting and this restriction can be tested empirically once the two cointegrating vectors are *exactly* identified. When $\alpha_{c1} = \alpha_{c2}$ the consumption Euler equation in the case of no rule of thumb consumers is given by $E_t \Delta c_{t+1} = \vartheta_c + \alpha_{c1}(\beta_{R1} + \beta_{R2})R_t$, where $\Gamma_1 = 0$, $\gamma = 0$, $\Phi = 0$, $\vartheta_c = \phi$ and $\alpha_{c1}(\beta_{R1} + \beta_{R2}) = \sigma$, in accordance with (3) with constant variance.

The "saving for a rainy day" hypothesis is likewise given by $E_t \Delta y_{t+1} = \vartheta_y + (\alpha_{y1} - \alpha_{y2})(c_t - y_t) + (\alpha_{y1}\beta_{R1} + \alpha_{y2}\beta_{R2})R_t$, where $\vartheta_y = \varsigma$ and $(\alpha_{y1} - \alpha_{y2})^{-1} = \rho$, in line with (2). The additional term $(\alpha_{y1}\beta_{R1} + \alpha_{y2}\beta_{R2})R_t$ makes the "saving for a rainy day" hypothesis in (9) somewhat less restrictive than (2) in the sense that the real after tax interest rate is allowed to vary over time. The additional term is easy to handle such that the CVAR also nests all the hypotheses in (5) with some rule of thumb consumers. To see this, we multiply (9) by the matrix $c' = (1, -\mu\lambda, 0, 0)$, still assume $\alpha_{c1} = \alpha_{c2}$, $\beta_{y1} = -1$ and $\beta_{w2} = 0$, and rearrange terms to obtain

$$
\begin{aligned}
(10)\; E_t \Delta c_{t+1} - \mu\lambda E_t \Delta y_{t+1} \;=\;& \vartheta_c - \mu\lambda\vartheta_y + (\gamma_{1,12} - \mu\lambda\gamma_{1,22})\Delta y_t \\
& + [\alpha_{c1}(\beta_{R1} + \beta_{R2}) - \mu\lambda(\alpha_{y1}\beta_{R1} + \alpha_{y2}\beta_{R2})]R_t \\
& + (\gamma_{1,11} - \mu\lambda\gamma_{1,21})\Delta c_t - \mu\lambda(\alpha_{y1} - \alpha_{y2})(c_t - y_t),
\end{aligned}
$$

where $\gamma = 0$, $\Phi = 0$, $\vartheta_c - \mu\lambda\vartheta_y = (1 - \mu)\phi$, $\gamma_{1,12} - \mu\lambda\gamma_{1,22} = \mu(1 - \lambda)$, $\alpha_{c1}(\beta_{R1} + \beta_{R2}) - \mu\lambda(\alpha_{y1}\beta_{R1} + \alpha_{y2}\beta_{R2}) = (1 - \mu)\sigma$, $\gamma_{1,11} - \mu\lambda\gamma_{1,21} = \tau$ and $-\mu\lambda(\alpha_{y1} - \alpha_{y2}) = \varrho$.

The theories we have discussed above entail different outcomes for subsequent empirical estimation of the consumption equation. First, a logarithmic version of the martingale hypothesis by Hall (1978), $E_t \Delta c_{t+1} = 0$, implies that $\mu\lambda$ equals zero and that no significant terms appear on the right hand side of (10). Second, precautionary savings in response to uncertainty are reflected in the intercept, $\vartheta_c - \mu\lambda\vartheta_y$. Third, a significantly positive estimate of $[\alpha_{c1}(\beta_{R1} + \beta_{R2}) - \mu\lambda(\alpha_{y1}\beta_{R1} + \alpha_{y2}\beta_{R2})]$ can be interpreted as the intertemporal elasticity of substitution in consumption. Fourth, a significantly positive estimate of $(\gamma_{1,11} - \mu\lambda\gamma_{1,21})$ points to quadratic adjustment costs or habit formation in consumption. Fifth, significantly positive estimates of $\mu\lambda$ and $(\gamma_{1,12} - \mu\lambda\gamma_{1,22})$ indicate a substantial portion of rule of thumb consumers responding to current and one period lag in income growth, respectively. Finally, a significantly positive estimate of $\mu\lambda(\alpha_{y1} - \alpha_{y2})$ can be interpreted as the coefficient of speed of adjustment in a disequilibrium model of consumption and income.

The Keynesian consumption function approach, as opposed to the Euler equation approach, implies that consumption *is* equilibrium correcting in the CVAR. To simplify the exposition, we now assume that $r = 1$. When the cointegration vector is normalised with respect to consumption and $k = 2$, the CVAR in (8) becomes

$$
\begin{aligned}
(11) \qquad \begin{pmatrix} \Delta c_t \\ \Delta y_t \\ \Delta w_t \\ \Delta R_t \end{pmatrix} &= \begin{pmatrix} \alpha_c \\ \alpha_y \\ \alpha_w \\ \alpha_R \end{pmatrix} [c_{t-1} - \beta_y y_{t-1} - \beta_w w_{t-1} - \beta_R R_{t-1}] \\
&\quad + \Gamma_1 \begin{pmatrix} \Delta c_{t-1} \\ \Delta y_{t-1} \\ \Delta w_{t-1} \\ \Delta R_{t-1} \end{pmatrix} + \gamma t + \vartheta + \Phi D_t + \epsilon_t.
\end{aligned}
$$

It follows that consumption is equilibrium correcting when $-1 < \alpha_c < 0$. However, income, wealth and the real after tax interest rate may also be equilibrium correcting if the corresponding value of alpha is positive and less than one. If, on the other hand, $\alpha_y = \alpha_w = \alpha_R = 0$, then income, wealth and the real after tax interest rate are all weakly exogenous with respect to $\beta$ and the conditional Keynesian consumption function from (11) becomes

$$
\begin{aligned}
(12) \Delta c_t &= \alpha_c[c_{t-1} - \beta_y y_{t-1} - \beta_w w_{t-1} - \beta_R R_{t-1}] + \omega_y \Delta y_t + \omega_w \Delta w_t + \omega_R \Delta R_t \\
&\quad + \tilde{\gamma}_{1,11}\Delta c_{t-1} + \tilde{\gamma}_{1,12}\Delta y_{t-1} + \tilde{\gamma}_{1,13}\Delta w_{t-1} + \tilde{\gamma}_{1,14}\Delta R_{t-1} \\
&\quad + \tilde{\gamma}_c t + \tilde{\vartheta}_c + \tilde{\Phi}_c D_t + \tilde{\epsilon}_{ct},
\end{aligned}
$$

where the inclusion of the contemporaneous variables, $\Delta y_t$, $\Delta w_t$ and $\Delta R_t$, follows from the properties of the multivariate normal error distribution and where the coefficients are linear functions of the coefficients in (11) and the parameters from the multivariate normal error distribution; see for instance Johansen (1995, p. 122).

We have seen that cointegration in (8) represents the common ground between the consumption Euler equation approach and the Keynesian consumption function approach and that the theoretical predictions from the two approaches put different restrictions with respect to weak exogeneity on consumption and income.

In the empirical analysis, we shall therefore consider hypotheses of cointegration and equilibrium correction as restrictions on $\Pi = \alpha\beta'$ in order to discriminate between the two approaches. Because CVAR models considering conditional expectations of future consumption and income may corroborate parameters stemming from the class of Euler equations, we shall also examine the empirical relevance of such models within the context of Johansen and Swensen (1999, 2008).

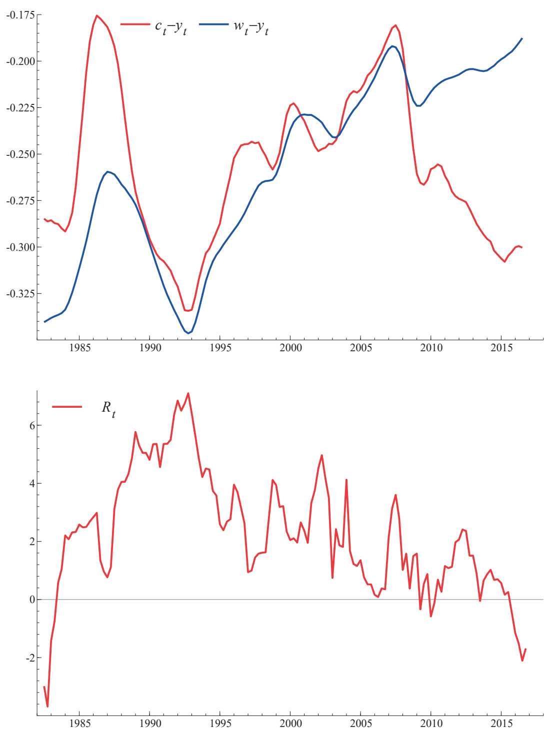
# 3 Data

For comparison reasons, we maintain the data set from Jansen (2013) *as is* and extend it by using quarterly growth rates from the final national accounts for the period $2008q3-2016q4$, keeping $2008q2$ fixed. Because the capital markets in Norway were heavily regulated during the 1970s and early 1980s, which likely prevented many consumers from acting freely in accordance with a consumption Euler equation, we choose $1984q1$ as the starting point of our sample period. However, due to lags in the CVAR, the sample period for estimation purposes includes data points from $1982q3$ to $2016q4$. The sample period is thus consistent with the period of liberalisation of what was believed to be the most binding regulations of credit for households, namely the bond market which was deregulated in several steps between 1982 and 1985 to allow for competition among banks and other lending institutions in the household market. We also choose $2008q4$ as the starting point of the financial crisis. Albeit the bankruptcy of Lehman Brothers took place the 15th of September 2008, we believe that the main effects on the Norwegian economy, and hence on the households' consumer behaviour, emerged in the fourth quarter of 2008 onwards.

The consumption variable is defined as real consumption excluding expenditures on health services and housing. The income variable is real disposable income excluding equity income. The wealth variable is measured in real terms net of household debt and thus consists of the value of housing as well as total net financial wealth. Finally, the real after tax interest rate is defined as the average nominal interest rates on bank loans faced by households net of marginal income tax and adjusted for inflation. In Online Appendix A, we give more precise definitions of all the variables entering the empirical models in Sections 4 and 5.

Figure 1 shows the consumption to income and the wealth to income ratios together with the real after tax interest rate for the sample period $1982q3-2016q4$. We observe a strong co-movement between the two ratios in the sample period before the financial crisis hit the Norwegian economy and this is *prima facie* evidence for cointegration between the three variables involved. However, a break in the cointegration relationship seems evident in the subsequent period as the two ratios then diverge and move in opposite directions. The real after tax interest rate for its part reached a historical high level in the early 1990s in the wake of the huge boom in consumption during the second half of the 1980s. Since then the real after tax interest rate has shown a downward trend and reached negative levels as in the early 1980s at the end of the sample period. These features of the data are the premises for the cointegration analysis and the modelling of short run dynamics.

Figure 1: The consumption to income $(c_t - y_t)$ ratio, the wealth to income $(w_t - y_t)$ ratio and the real after tax interest rate $(R_t)$




*Notes*: Sample period: 1982$q$3$-$2016$q$4. Left panel shows moving averages of the two ratios in logarithms, with one quarter lag and two quarters lead. Mean and range of the logarithms of wealth to income are matched to mean and range of the logarithms of consumption to income. Right panel shows the real after tax interest rate measured in per cent per annum.

# 4 Cointegration and dynamics[4]

In this section, we first carry out a multivariate cointegration analysis with a structural break around the financial crisis in 2008, applying the models and methods in Johansen *et al.* (2000). Then, we estimate consumption short run dynamics within a partial CVAR, following the modelling strategy in Harbo *et al.* (1998), to calculate the magnitude of the MPC.

## 4.1 Test results

A preliminary analysis of $\Pi = \alpha\beta'$, using (8) as the underlying model with $k = 6$ guided from both Akaike's information criterion (AIC), likelihood ratio tests of sequential model reduction and diagnostic tests of the residuals, confirms a significant structural break in the long run relationship around the financial crisis.[5] We are therefore motivated to follow Johansen *et al.* (2000) and capture a structural break in the long-run relationship by a model which takes into account the possibility of separate trends in the two periods $1, \ldots, T_1$ and $T_1 + 1, \ldots, T$. The idea is to allow for two VAR models where the $k$ first observations are conditioned upon, but where the parameters of the stochastic components are the same for both models and where the parameters of the deterministic components may differ corresponding to a broken trend. Formally, let $T_0 = 0$ and $T_2 = T$. If $ID_{j,t} = 1$ for $t = T_{j-1}$ and $ID_{j,t} = 0$ else so that $ID_{j,t-i}$ is the indicator for the $i$th observation in the $j$th period, $j = 1, 2$, it follows that $SD_{j,t} = \Sigma_{i=k+1}^{T_j - T_{j-1}} ID_{j,t-i} = 1$ for $t = T_{j-1} + k + 1, \ldots, T_j$ and $SD_{j,t} = 0$ else. The CVAR in (8) is then reformulated for $t = k + 1, \ldots, T$ as

$$(13) \quad \Delta X_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ tSD_t \end{pmatrix} + \mu SD_t + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-(k-1)} + \Phi D_t + \kappa_{2,1} ID_{2,t-1} + \ldots + \kappa_{2,k} ID_{2,t-k} + \epsilon_t,$$

where $SD_t = (SD_{1,t}, SD_{2,t})'$, $\gamma = (\gamma_1', \gamma_2')'$ and $\mu = (\mu_1, \mu_2)$. We assume, referring to Figure 1 and the discussion in Section 3, that the break occurs in 2008$q$4. Accordingly, $SD_{1,t}$ is a step dummy which equals one in the period 1982$q$3−2008$q$3, $SD_{2,t}$ is a step dummy which equals one in the period 2010$q$2−2016$q$4 and $ID_{2,t}$ are impulse indicators which equal one for $t$ = 2008$q$4,...,2010$q$1, otherwise zero for $k = 6$.

We can now conduct a cointegration analysis with the augmented VAR, letting $SD_t$ and $ID_{2,t}$ enter unrestrictedly, whereas $tSD_t$ is restricted to lie in the cointegration space. Again, according to both the AIC and the series of model reduction tests, the VAR in our case should include six lags as the premise for the cointegration analysis. Also, with $k = 6$ there are no serious departures from white noise residuals according to the diagnostic tests. Juselius (2006, p. 72) suggests as

[4]The econometric modelling in this section is carried out with PcGive 14; see Doornik and Hendry (2013).

[5]Detailed results can be found in Online Appendix B.

Table 1: Trace test results for cointegration with a structural break[1]

| Eigenvalue ($\lambda_i$) | $H_0$ | $H_A$ | $\lambda_{trace}$ |
|---|---|---|---|
| 0.287 | $r = 0$ | $r \geq 1$ | 101.21 [0.001] |
| 0.188 | $r \leq 1$ | $r \geq 2$ | 56.49 [0.050] |
| 0.129 | $r \leq 2$ | $r \geq 3$ | 28.96 [0.207] |
| 0.078 | $r \leq 3$ | $r = 4$ | 10.74 [0.372] |

| Diagnostics[2] | Test statistic | $p$-value |
|---|---|---|
| Vector autocorrelation 1-5 test: | F(80,286)=1.37 | 0.034 |
| Vector normality test: | $\chi^2(8) = 9.02$ | 0.341 |
| Vector heteroscedasticity test: | F(224,266)=1.03 | 0.407 |

Sample period: 1982$q$3$-$2016$q$4. [1] See Johansen *et al.* (2000), VAR of order 6, modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_t$ (restricted), $SD_t$ (unrestricted), $ID_{2t}$ (unrestricted) and centered seasonal dummies (unrestricted), $r$ denotes the rank order of $\Pi = \alpha\beta'$ and $\lambda_{trace}$ is the trace statistic with $p$-value in square brackets, which are calculated by means of the estimated response surface in Johansen *et al.* (2000, Table 4). [2] See Doornik and Hendry (2013, p. 172).

a rule of thumb to use a VAR with $k = 2$ in a tentative cointegration analysis. We find, however, that such a specification suffers from severe residual autocorrelation. Because (13) controls for a structural break around the financial crisis and the fact that both the fourth and the fifth lag of consumption dynamics are strongly significant in the estimated models in Subsections 4.2 and 5.2, we argue that the severe residual autocorrelation is associated with left-out dynamics rather than structural misspecification. We therefore claim that cointegration results based on VAR(6) are more likely to represent the underlying data structure.[6]

Table 1 shows trace test results for cointegration together with the diagnostic tests of the underlying VAR. The trace tests support marginally the hypothesis of two cointegrating vectors between $c_t$, $y_t$, $w_t$ and $R_t$ at the 5 per cent significance level. We shall therefore consider two cases, $r = 2$ and $r = 1$, when testing restrictions on $\Pi = \alpha\beta'$ to discriminate between the consumption Euler equation and the Keynesian consumption function. Starting with $r = 2$, we find, once the two cointegrating vectors are exactly identified as described below (9), that the restriction $\alpha_{c1} = \alpha_{c2}$ is strongly rejected by a likelihood ratio statistic, which becomes $\chi^2(1) = 10.65$ with a $p$-value of 0.001. We may conclude already at this stage of the analysis that the data overwhelmingly reject a consumption Euler equation.

Table 2 reports main likelihood ratio tests of restrictions on $\Pi = \alpha\beta'$ assuming $r = 1$. The hypothesis of homogeneity between consumption, income and wealth is not rejected by the data in Model (*ii*). Note that the trend variable for the first period is *not* excluded from the model as the estimate of $\gamma_1$ is a borderline case at the 10 per cent significance level ($p$-value = 0.103). A likelihood ratio test, $\chi^2(1) = 1.94$ and $p$-value = 0.16, supports reduction from Model (*ii*) to Model (*iii*)

[6] A comparative cointegration analysis with VAR(2) together with a summary of the AIC, the model reduction tests and diagnostic tests are given in Online Appendix C.

Table 2: Likelihood ratio test results for restrictions on $\Pi = \alpha\beta'$ with a structural break[1]

Model $(i)$: $\beta_c = 1$

$$c_{t-1} - \underset{(0.23)}{1.06} y_{t-1} - \underset{(0.039)}{0.16}\, w_{t-1} + \underset{(0.32)}{1.72} R_{t-1} + \underset{(0.0013)}{0.0023} tSD_{1,t} + \underset{(0.0020)}{0.0066} tSD_{2,t}$$

$$\hat{\alpha}_c = \underset{(0.09)}{-0.31},\ \hat{\alpha}_y = \underset{(0.07)}{0.003},\ \hat{\alpha}_w = \underset{(0.11)}{-0.27},\ \hat{\alpha}_R = \underset{(0.03)}{-0.15}$$

$\log L = 1558.35$

Model $(ii)$: $\beta_c = 1$, $\beta_y + \beta_w = 1$

$$c_{t-1} - 0.84 y_{t-1} - \underset{(0.04)}{0.16} w_{t-1} + \underset{(0.35)}{1.97} R_{t-1} + \underset{(0.00028)}{0.00089} tSD_{1,t} + \underset{(0.001)}{0.0051} tSD_{2,t}$$

$$\hat{\alpha}_c = \underset{(0.08)}{-0.26},\ \hat{\alpha}_y = \underset{(0.07)}{-0.008},\ \hat{\alpha}_w = \underset{(0.10)}{-0.24},\ \hat{\alpha}_R = \underset{(0.03)}{-0.14}$$

$\log L = 1557.71$

$\chi^2(1) = 1.28[0.26]$[2]

Model $(iii)$: $\beta_c = 1$, $\beta_y = 1$, $\beta_w = 0$

$$c_{t-1} - y_{t-1} + \underset{(0.63)}{4.28} R_{t-1} + \underset{(0.00035)}{0.00026} tSD_{1,t} + \underset{(0.0021)}{0.0074} tSD_{2,t}$$

$$\hat{\alpha}_c = \underset{(0.04)}{-0.10},\ \hat{\alpha}_y = \underset{(0.03)}{-0.026},\ \hat{\alpha}_w = \underset{(0.05)}{-0.15},\ \hat{\alpha}_R = \underset{(0.01)}{-0.07}$$

$\log L = 1556.74$

$\chi^2(2) = 3.21[0.20]$[3], $\chi^2(1) = 2.16[0.14]$[4]

Model $(iv)$: $\beta_c = 1$, $\beta_y + \beta_w = 1$, $\alpha_y = 0$

$$c_{t-1} - 0.83 y_{t-1} - \underset{(0.04)}{0.17} w_{t-1} + \underset{(0.35)}{1.93} R_{t-1} + \underset{(0.00028)}{0.00091} tSD_{1,t} + \underset{(0.00099)}{0.0050} tSD_{2,t}$$

$$\hat{\alpha}_c = \underset{(0.08)}{-0.26},\ \hat{\alpha}_w = \underset{(0.10)}{-0.24},\ \hat{\alpha}_R = \underset{(0.03)}{-0.15}$$

$\log L = 1557.70$

$\chi^2(2) = 1.28[0.53]$[5], $\chi^2(1) = 0.002[0.97]$[6]

Sample period: $1982q3-2016q4$. [1] See Johansen *et al.* (2000), VAR of order 6 with a structural break in $2008q4$, $r = 1$, modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_{1,t}$ and $tSD_{2,t}$ (restricted), $SD_{1,t}$ and $SD_{2,t}$ (unrestricted), $ID_{2,t}$ (unrestricted) and centered seasonal dummies (unrestricted), standard errors in parenthesis, $p$-values in square brackets. [2] $\beta_y + \beta_w = 1$. [3] $\beta_y = 1$, $\beta_w = 0$. [4] $\beta_w = 0$. [5] $\beta_y + \beta_w = 1$, $\alpha_y = 0$. [6] $\alpha_y = 0$.

in which homogeneity between consumption and income and exclusion of the wealth variable are imposed. However, the $p$-value of the individual hypothesis $\beta_w = 0$ equals 0.14 and the associated $t$-value is as high as 4 in magnitude in Model ($ii$). In addition, the estimates of $\beta_R$ and $\alpha$ change significantly when imposing homogeneity between consumption and income only. We therefore keep the wealth variable in the cointegrating vector and find that the estimated adjustment coefficients, except from the estimate of $\alpha_y$ ($p$-value = 0.97), are all highly significant in Model ($iv$). Accordingly, consumption, and not income, equilibrium corrects in the CVAR, which clearly contradicts the consumption Euler equation. When imposing homogeneity between consumption, income and wealth and weak exogeneity of income, the restricted long-run relationship becomes

$$\text{(14)}\quad \widehat{eqcm}_t = c_{t-1} - 0.83y_{t-1} - 0.17w_{t-1} + 1.93R_{t-1} + 0.00091tSD_{1,t} + 0.0050tSD_{2,t}.$$

Recursively estimated coefficients of $y_t$, and hence also of $w_t$, as well as of $R_t$, are stable before, during and after the financial crisis once the structural break is allowed for. Also, recursive likelihood ratio tests support the joint hypothesis of $\beta_y + \beta_w = 1$ and $\alpha_y = 0$.[7] A comparison with equation (2) in Jansen (2013) shows that the estimated long run coefficients of income and wealth are almost perfectly reproduced on the sample period ending in 2016$q$4. We also find that the deterministic trend in (14) is significantly shifting equilibrium consumption downwards both before and after the financial crisis. However, the shift is much larger after 2008$q$4, with a factor of 5.5 according to model ($iv$). A possible interpretation may be that the broken trend reflects increased uncertainty and thus increased precautionary savings in the wake of the financial crisis. The fact that the households' saving ratio increased from nearly 4 per cent in 2008 to more than 10 per cent in 2015 supports this conjecture.

## 4.2 Short run dynamics

To facilitate a comparison of the magnitude of the MPC implied by Model A1 in Table 4 in Jansen (2013), we perform a reduced rank regression for a partial model following the modelling strategy of equation (10) in Harbo *et al.* (1998). Since the hypothesis of weak exogeneity of income with respect to the long-run parameters is supported by the data, we can without loss of information condition on this variable when estimating a partial CVAR for consumption, wealth and the real after tax interest rate. Our point of departure is therefore the partial model written in vector form as

$$\text{(15)}\qquad \Delta X_t^* = \Theta_D D_t + \sum_{j=0}^{5} \omega_{Z,j} \Delta Z_{t-j} + \sum_{j=1}^{5} \Theta_{X^*,j} \Delta X_{t-j}^* + \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ tSD_t \end{pmatrix} + \varepsilon_t,$$

[7] See Online Appendix D.

where $X_t^* = (c_t, w_t, R_t)'$, $Z_t = y_t$, $X_t = (c_t, y_t, w_t, R_t)'$ and $D_t$ includes the centered seasonal dummies and all the dummies for the structural break around the financial crisis. First, we estimate (15) by constrained full information maximum likelihood (CFIML) whereby the rank is restricted to one and the hypothesis of homogeneity between consumption, income and wealth is imposed in accordance with the evidence above. Then, we simplify the model, general-to-specific, by deleting the most *in*significant short run dynamics one-by-one within the system as a whole. The simplified dynamic consumption model with respect to the stochastic variables and the broken trend in the long run relationship becomes

$$
\begin{aligned}
(16)\ \Delta\widehat{c}_t \quad = \quad & -\underset{(0.09)}{0.34}\Delta c_{t-1} - \underset{(0.07)}{0.14}\Delta c_{t-2} + \underset{(0.07)}{0.45}\Delta c_{t-4} + \underset{(0.08)}{0.25}\Delta c_{t-5} + \underset{(0.09)}{0.21}\Delta y_t \\
& + \underset{(0.10)}{0.27}\Delta y_{t-3} + \underset{(0.11)}{0.24}\Delta y_{t-4} + \underset{(0.07)}{0.26}\Delta w_{t-1} - \underset{(0.07)}{0.15}\Delta w_{t-4} - \underset{(0.10)}{0.34}c_{t-1} \\
& + \underset{(0.07)}{0.26}y_{t-1} + \underset{(0.03)}{0.08}w_{t-1} - \underset{(0.11)}{0.28}R_{t-1} - \underset{(0.0001)}{0.0004}tSD_{1,t} - \underset{(0.0005)}{0.0016}tSD_{2,t}.
\end{aligned}
$$

The consumption model seems quite well-specified as judged by both diagnostics tests and plots of model fit and residuals.[8] Interestingly, (16) implies a first year MPC of around 40 per cent, which is quite close to 30 per cent implied by Model A1 in Jansen (2013) and recent microeconometric evidence referred to in Section 1. These findings are in line with the argument in Doornik and Hendry (1997) that the main source of forecast failure is deterministic shifts in equilibrium means, e.g. the equilibrium saving ratio, and not shifts in the derivative coefficients, e.g. the marginal propensity to consume, that are of primary interest for policy analysis.

So far we have modelled the financial crisis as a structural break in the long run relationship between consumption, income, wealth and the real interest rate. As a final exercise, we test for parameter constancy of the short run dynamic wealth effects, which one could argue have changed after the financial crisis. Indeed, adding lagged effects of the interaction between $SD_{2,t}$ and $\Delta w_t$ in the whole specific system, the fourth lag turned out with a significant and positively signed parameter estimate in the consumption model only.[9] A possible interpretation of the fact that the short run response of consumption to wealth has increased after the financial crisis is the following: After the financial crisis households have faced increased credit constraints by *inter alia* lending criteria based on payment-to-income ratios due to increased credit risk in the economy. Households have thus not been able to borrow at the observed lending interest rates as easily as before the financial crisis because of tightening of the credit practises. As a consequence, households credit worthiness, as measured by total wealth, has become increasingly important for mortgage and other loan security after the financial crisis, and thus also for the ability to borrow for consumption purposes in the short run.

Before turning to conditional expectations of future consumption and income, one may also interpret the rejection of a consumption Euler equation as being due to

[8]Detailed estimation results, diagnostic tests and plots of model fit and residuals of the whole specific system are reported in Online Appendix E.

[9]Detailed results are given in Online Appendix F.

increased credit constraints after the financial crisis which has made it much more difficult for households to smooth consumption. However, redoing our cointegration analysis with a sample period ending in 2008$q$3, as shown in Online Appendix B, leads to the same conclusions regarding the validity of the consumption Euler equation. For example, the hypothesis $\alpha_{c1} = \alpha_{c2}$ is even more strongly rejected for this shorter sample period than for the whole sample. Our results are thus quite robust to the choice of sample period.

# 5 Conditional expectations[10]

Although the findings from testing restrictions on $\Pi = \alpha\beta'$ do not support a consumption Euler equation, the question whether conditional expectations of future consumption and income play a role in explaining the consumer behaviour is left unanswered. We recall from Subsection 2.3 that the conditional expectations of future consumption and income in (10) nests all the hypotheses in (5) with some rule of thumb consumers. Conditional expectations of future consumption and income can be treated within the context of Johansen and Swensen (1999, 2008). First, we outline the estimation and testing procedure, paying particular attention to the conditional expectations restrictions on the stochastic part of the CVAR. Then, we estimate CVAR models using (5) with conditional expectations as reference and examine whether data can corroborate parameters stemming from the class of Euler equations. We shall throughout the analysis specify the *exact* form of CVAR models in which the rank order of the impact matrix is one and income is not weakly exogenous as in Subsection 4.2.

## 5.1 Outline of the estimation and testing procedure

To simplify the exposition, we outline the estimation and testing procedure by means of (8). The consumption Euler equations involving expectations of future variables can generally be expressed as $c'E_t\Delta X_{t+1} = d'X_t$, which implies restrictions on the coefficients in (8). For instance, a bivariate system where the variables satisfy a martingale hypothesis can be written $E_tX_{1,t+1} = (1,0)E_t(X_{1,t+1}, X_{2,t+1})' = (1,0)(X_{1,t}, X_{2,t})' = X_{1,t}$ or $(1,0)E_t\Delta(X_{1,t+1}, X_{2,t+1})' = 0$. It is often convenient to have a more general specification of the form

$$(17) \quad c'E_t\Delta X_{t+1} - d'X_t + d'_{-1}\Delta X_t + \ldots + d'_{-k+1}\Delta X_{t-k+2} + \vartheta_0 + \gamma_0 t + \Phi_0 D_t = 0$$

where $c$, $d$, $d_{-1}$,$\ldots$,$d_{-k+1}$, $\vartheta_0$, $\gamma_0$ and $\Phi_0$ have known elements.

A flexible formulation arises by assuming that the $p \times q$ matrix $c$ is known and allowing $d$, $d_{-1}$,$\ldots$,$d_{-k+1}$, $\vartheta_0$, $\gamma_0$ and $\Phi_0$ to be treated as matrices of unknown parameters. If they are allowed to vary freely (17) does not imply any constraints. By testing whether any of the matrices $d, d_{-1}, \ldots, d_{-k+1}$, $\vartheta_0$, $\gamma_0$ and $\Phi_0$ in (17) are

[10]The estimation and testing in this section are performed with the statistical package R; see R Core Team (2019).

equal to zero, or any given matrix, one can investigate whether a simplification of the relation of the conditional expectations is possible.

Using the methods described in Johansen and Swensen (1999, 2008) the value of the concentrated likelihood $L_c(d, d_{-1}, \ldots, d_{-k+1}, \vartheta_0, \gamma_0, \Phi_0)$ can be computed. Further maximization over $d, d_{-1}, \ldots, d_{-k+1}, \vartheta_0, \gamma_0$ and $\Phi_0$ yields a value $\max L_c(d, d_{-1}, \ldots, d_{-k+1}, \vartheta_0, \gamma_0,$
$\Phi_0)$ which is equal to the maximal value of the likelihood for (8), denoted as $L_{max}$. The likelihood ratio for a test of a particular hypothesis, for instance $d_{-k+1} = d^0_{-k+1}$, can be found as

$$\frac{\max_{d,d_{-1},\ldots,d_{-k+2},\vartheta_0,\gamma_0,\Phi_0} L_c(d, d_{-1}, \ldots, d^0_{-k+1}, \vartheta_0, \gamma_0, \Phi_0)}{\max_{d,d_{-1},\ldots,d_{-k+2},d_{-k+1},\vartheta_0,\gamma_0,\Phi_0} L_c(d, d_{-1}, \ldots, d_{-k+1}, \vartheta_0, \gamma_0, \Phi_0)} = \frac{\max_{d,d_{-1},\ldots,d_{-k+2},\vartheta_0,\gamma_0,\Phi_0} L_c(d, d_{-1}, \ldots, d^0_{-k+1}, \vartheta_0, \gamma_0, \Phi_0)}{L_{max}}.$$

The maximization with respect to $d_{-1}, \ldots, d_{-k+2}, \vartheta_0, \gamma_0$ and $\Phi_0$ can be performed by ordinary least squares (OLS) and reduced rank regression, while maximizing with respect to $d$ must be carried out using numerical optimization. A more detailed explanation of the procedure can be found in Online Appendix G.

## 5.2 Estimated CVAR models

The conditional expectations, $c'E_t\Delta X_{t+1}$, can also be found from (13) by leading the variables one period and taking expectations at time $t$. Comparing the coefficients from an augmented version of (17), where a broken trend is taken into account, implies the following restrictions on the stochastic variables

$$c'\alpha\beta' = d', c'\Gamma_1 = -d'_{-1}, \ldots, c'\Gamma_{k-1} = -d'_{-k+1}, \tag{18}$$

while the non-stochastic variables are unspecified and have the form $c'\alpha\gamma'(tSD_{t+1} + SD_{t+1}) + c'\Phi D_{t+1} + c'\mu SD_{t+1} + c'\kappa_{2,1}ID_{2,t} + \ldots + c'\kappa_{2,k}ID_{2,t-k+1}$. As indicated in the introduction, it is reasonable to start with considering the coefficients of the stochastic variables as the parameters of interest.

We first investigate the case of conditional expectations involving only consumption, $c = (1, 0, 0, 0)'$, and focus on the consumption equation, not the whole system as in Subsection 4.2, when deleting any insignificant short run dynamics from the model, that is any insignificant coefficients in the first row of the matrices $\Gamma_j$ denoted as $(\gamma_{j,11}, \gamma_{j,12}, \gamma_{j,13}, \gamma_{j,14})$ for $j = 1, \ldots, 5$. Specifically, we use a sequential reduction procedure, starting with the coefficients of the last lag of growth in the real interest rate, that is $\gamma_{5,14}$.

Table 3 shows likelihood ratio test results for simplifying restrictions on $c'\Gamma_1 = -d'_{-1}, \ldots, c'\Gamma_{k-1} = -d_{-k+1}$. As pointed out earlier, when maximizing first over $d'_{-1}, \ldots, d'_{-k+1}$ and the coefficients of the non-stochastic variables and then over the parameters of the cointegration vector, the value of the maximum of the likelihood is the same as for (13), see Table 2 models ($i$ ) and ($ii$). By sequentially simplifying

Table 3: Likelihood ratio test results for simplifying restrictions[1]

| Model | Restrictions | $\log L_i$ | $i-j$[2] | $-2\log\frac{L_j}{L_i}$ | df | p-value |
|---|---|---|---|---|---|---|
| 1 | - | 1558.35 | - | - | - | - |
| 2 | $\beta_y+\beta_w=1$ | 1557.71 | 1-2 | 1.28 | 1 | 0.26 |
| 3 | Model 2, $\gamma_{5,14}=0$ | 1557.69 | 2-3 | 0.04 | 1 | 0.84 |
| 4 | Model 3, $\gamma_{5,13}=0$ | 1556.55 | 3-4 | 2.28 | 1 | 0.13 |
| 5 | Model 4, $\gamma_{5,12}=0$ | 1554.87 | 4-5 | 3.36 | 1 | 0.07 |
| 6 | Model 5, $\gamma_{5,11}=0$ | 1549.34 | 5-6 | 11.06 | 1 | 0.0009 |
| 7 | Model 5, $\gamma_{4,14}=0$ | 1554.08 | 5-7 | 1.58 | 1 | 0.21 |
| 8 | Model 7, $\gamma_{4,13}=0$ | 1550.13 | 7-8 | 7.90 | 1 | 0.005 |
| 9 | Model 7, $\gamma_{3,14}=0$ | 1554.06 | 7-9 | 0.04 | 1 | 0.84 |
| 10 | Model 9, $\gamma_{2,14}=0$ | 1552.99 | 9-10 | 2.14 | 1 | 0.14 |
| 11 | Model 10, $\gamma_{1,14}=0$ | 1551.06 | 10-11 | 3.86 | 1 | 0.05 |

Sample period: 1982$q$3–2016$q$4. [1] See Johansen and Swensen (1999, 2008). [2] $i-j$ denotes the likelihood ratio test for the additional restriction(s) on model $j$ compared to model $i$.

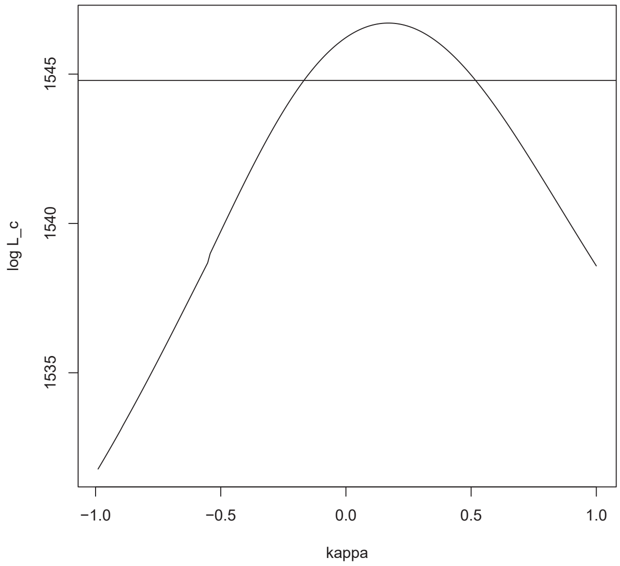
the model, we end up with Model 11 which includes five lags of consumption growth, four lags of income and wealth growth and no lag of growth in the real interest rate. As a further simplification, the remaining short run dynamics of Model 11 having coefficients with a p-value larger than 0.1 are set equal to zero. For the stochastic variables and the broken trend in the long run relationship the result is

$$\widehat{E_t\Delta c_{t+1}} = -0.16(1.0,-0.83,-0.17,2.76,0.0010,0.0059)\begin{pmatrix}X_t\\ tSD_{t+1}\end{pmatrix} \\ -\underset{(0.08)}{0.45}\Delta c_t - \underset{(0.07)}{0.27}\Delta c_{t-1} + \underset{(0.07)}{0.53}\Delta c_{t-3} + \underset{(0.08)}{0.30}\Delta c_{t-4} \\ +\underset{(0.09)}{0.14}\Delta y_{t-3} + \underset{(0.07)}{0.24}\Delta w_t - \underset{(0.07)}{0.20}\Delta w_{t-3}, \tag{19}$$

where $X_t=(c_t,y_t,w_t,R_t)'$ and $tSD_{t+1}=(tSD_{1,t+1},tSD_{2,t+1})'$. The corresponding maximal value of the log likelihood is 1546.23. The further reduction of Model 11 is thus valid by a likelihood ratio test with $\chi^2(6)=9.66$ and $p=0.14$. There are several interesting consequences of (19). The first is a clear rejection of the hypothesis that log consumption is a martingale, $E_t\Delta c_{t+1}=0$, which is a variant of the hypothesis of Hall (1978) in (1). The result found in Jansen (2013) is therefore confirmed for the whole period. The second is a clear rejection of the implicit restrictions of the consumption Euler equation in (3) with constant variance since (19) contains many significant coefficients involving growth in both income and wealth. The third is that the significant coefficients of lags of consumption growth can be interpreted as habit formation in line with the Smets and Wouters (2003) model in (6).

We next investigate the case of conditional expectations involving both future consumption and income, $c=(1,-\mu\lambda,0,0)'$, to shed light on the magnitude of the proportion of rule of thumb consumers. Accordingly, we analyse a simplified version of (5) involving only $\kappa=\mu\lambda$ and not all the economically interesting parameters

Figure 2: Concentrated log likelihood for $\kappa = \mu\lambda$ with a 90 per cent confidence interval




*Notes*: Sample period: $1982q3-2016q4$.

$\mu, \phi, \lambda, \sigma, \tau$ and $\varrho$ separately. Figure 2 plots the concentrated log likelihood for $\kappa = \mu\lambda$ in the region $[-1, 1]$ with $c = (1, -\mu\lambda, 0, 0)'$ using the same variables as in (19). The maximal value of the log likelihood is 1546.71, corresponding to the maximum likelihood estimate $\hat{\kappa} = 0.17$. The indicated 90 per cent confidence interval $(-0.11, 0.46)$ contains parameters which have a sensible economic interpretation. However, since the interval also contains $\kappa = 0$ this parameter cannot be claimed to be different from zero. More precisely, the likelihood ratio statistic $2(1546.71 - 1546.23) = 0.96$ corresponds to a p-value equal to 0.33. The conclusions drawn from (19) with $c = (1, 0, 0, 0)'$ concerning the martingale hypothesis, the consumption Euler equation in (3) and habit formations are therefore still valid.

Treating the coefficients of the non-stochastic variables as parameters of interest also has some conclusions worth mentioning. The step dummies $SD_t$ are introduced to take the financial crisis into account. It is a generally accepted fact that the financial crisis was unanticipated. Therefore, the coefficients of $SD_{t+1}$ should satisfy $c'\alpha\gamma_1 = c'\alpha\gamma_2$ and $c'\mu_1 = c'\mu_2$, that is there is only a linear trend such that the time of the break cannot be known. This hypothesis is, however, overwhelmingly rejected with a likelihood ratio statistic $2(1546.23 - 1540.55) = 11.36$ and a corresponding p-value equal to 0.003 with 2 degrees of freedom. The rejection of a hypothesis that is true, that is that the financial crisis is unanticipated, provides additional support to the conclusion from Section 4 that a consumption Euler equation is not a valid empirical description of the data.

In any case, we may also ask if the habit formation type behaviour in (19) is a consequence of assuming that the agents know the financial shift has occurred in the way expectations are formulated in the model. However, estimating the specific consumption model with a sample period ending in $2008q3$, we find that the lag structure of consumption growth is fairly similar to that in (19) for the whole sample period.[11] These findings indicate that the estimated habit formation type behaviour is not due to the way expectations about the financial shift are formulated in the model.

We conclude from all the findings in this section that when considering conditional expectations of consumption and income in CVAR models, most of the parameters stemming from the class of consumption Euler equations are not supported by the data. Only habit formation seems to play an important role in explaining Norwegian consumer behaviour.

# 6 Conclusions

In this paper, we have formulated a general CVAR that nests both a class of consumption Euler equations and various Keynesian type consumption functions. Using likelihood-based methods and Norwegian data, we found evidence of cointegration between consumption, income and wealth once a structural break around the financial crisis is accounted for. That consumption cointegrates with both income and wealth and not only with income demonstrates the empirical irrelevance of a consumption Euler equation. More importantly, we found that consumption equilibrium corrects to changes in income and wealth and not that income equilibrium corrects to changes in consumption, as would be the case if an Euler equation were true. Finally, we found that when conditional expectations of future consumption and income are considered in CVAR models, most of the parameters stemming from the class of Euler equations are not corroborated by the data. Only habit formation, typically included in DSGE models, seems to be important in explaining Norwegian consumer behaviour. Our estimated conditional Keynesian type consumption function implies a first year MPC of around 40 per cent.

We have relied on a CVAR in which a structural break in the cointegration relationship between consumption, income and wealth around the event of the financial crisis has been accounted for by a broken trend. A possible interpretation may be that the broken trend reflects increased uncertainty and thus increased precautionary savings in the wake of the financial crisis. Another possibility is that the broken trend picks up some important effects of omitted variables necessary to explain the changed consumer behaviour after the financial crisis. For instance, we have neither included a variable capturing the changing credit conditions faced by households nor disaggregated the wealth variable into separate variables for liquid assets, illiquid assets, debt and housing. Such variables may be important in a CVAR to adequately pick up effects of the household financial accelerator on consumption

[11]Detailed results are given in Online Appendix H.

in the wake of the financial crisis. We leave this issue for future work.

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# Online Appendices to "The consumption Euler equation or the Keynesian consumption function?"

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# Appendix A. Data definitions and sources

The original data set used in Jansen (2013) as well as the extended data set are collected from Statistics Norway and Norges Bank, unless otherwise noted.

$C_t$: Consumption expenditures in households and ideal organisations excluding expenditures on health services and housing, fixed 2006 prices. Source: Statistics Norway.

$Y_t$: Households real disposable income, excluding equity income, defined as nominal income deflated by $PC_t$, the price deflator for total consumption expenditures in households and ideal organisations excluding expenditures on health services and housing (2006=1). Source: Statistics Norway.

$W_t$: Real household wealth defined as nominal household wealth ($NW_t$), the sum of financial and housing wealth, deflated by $PC_t$.

$NW_t = [L_{t-1}+ML_{t-1}+NL_{t-1}-CR_{t-1} + (PH/PC)_t a_t K_{t-1}]PC_t$, where $L_t$ is household liquid assets (money stock and deposits), $ML_t$ is household medium liquid assets (equity and bonds), $NL_t$ is household non-liquid assets (insurance claims), $CR_t$ is household debt to banks and other financial institutions, $PH_t$ is housing price index (2006=1), $a_t$ is the fraction of residential housing stock owned by households and $K_t$ is the real value of the residential housing stock (fixed 2006 prices). All financial wealth components and the residential housing stock in the definition of $NW_t$ are in fixed values and they refer to the end of period $t-1$. The residential housing stock is updated each quarter by adding the gross investments in housing capital (in real terms deflated by $PH_t$) and deducting 0.4 per cent depreciation per quarter. Source: Statistics Norway. Data for nominal household wealth for the period 1982$q$3 to 1992$q$3 are from Erlandsen and Nymoen (2008). These data are chained in 1992$q$3 with data from Statistics Norway.

$R_t$: Real after tax interest rate for households defined as $4 \cdot RLB_t(1-\tau_t)-CPI_t/CPI_{t-4}+1$, where $RLB_t$ is average interest rate on households' bank loans, $\tau_t$ is marginal income tax rate faced by households and $CPI_t$ is headline consumer price index (2006=1). Sources: Statistics Norway and Norges Bank.

As mentioned in the paper, for comparison reasons we maintain the data set in Jansen (2013) *as is* and extend it by using quarterly growth rates from the final national accounts for the period 2008$q$3$-$2016$q$4, keeping 2008$q$2 fixed. As such, we follow both Brodin and Nymoen (1992), Jansen (2013) and Eitrheim *et al.* (2002) and work with non-per capita consumption, income and wealth in the empirical analysis. We note that $C_t/N_t = (Y_t/N_t)^{(1-\beta_w)} \cdot (W_t/N_t)^{\beta_w}$, where $N_t$ denotes the population, is equivalent to $c_t = (1-\beta_w)y_t + \beta_w w_t$ in the case of homogeneity between consumption, income and wealth. As shown in the paper, the homogeneity restriction is indeed supported by the data.[1]

The consumption variable is defined as real consumption excluding expenditures on health services and housing. Expenditures on health services are excluded from the consumption variable as almost all of these are refunded by the government. Likewise, the imputed housing consumption is closely related to the imputed value of housing income by construction in the national accounts. Thus it does not make sense to include this component in the consumption variable when one purpose of our study is to estimate the MPC. Since we want to test the theory implications of the permanent income hypothesis, it could be argued that also durable goods other than housing should be excluded from the consumption variable under study, see for instance Deaton (1992, pp. 99-103). However, data inspection reveals that the ratio between consumption of durables and our consumption variable fluctuates around a constant level, which suggests long-run constancy. Taking logarithms means that the difference between the two consumption variables will be captured adequately by the intercept term in the consumption models under study.

The income variable is real disposable income excluding equity income. The latter is left out because of episodes where tax increases on equity incomes were announced for the coming year leading to substantial tax motivated fluctuations in this income component, bearing in mind that equity incomes are likely to be less motivating for consumption than other incomes. Likewise, the wealth variable is measured in real terms net of household debt and thus consists of the value of housing as well as total net financial wealth. These entities differ widely in terms of liquidity and availability for the purpose of consumption of goods and services. We have nonetheless maintained the aggregated wealth measure in the sequel.

As also mentioned in the paper, we choose 2008q4 as the starting point of the financial crisis. We have checked that choosing 2008q3 as the starting point does not alter the empirical findings reported in the paper. These findings are available from the authors upon request.

# Appendix B. A preliminary cointegration analysis

In this appendix, we show that a preliminary cointegration analysis of $\Pi = \alpha\beta'$ using a CVAR without a trend break reveals a significant structural break in the long run relationship around the financial crisis.

## The sample period prior to the financial crisis

Our underlying CVAR, equation (8) in Subsection 2.3 of the paper, may be rewritten

[1] Erlandsen and Nymoen (2008), on the other hand, express their consumption function in per capita terms, but they emphasise that the results obtained do not depend in any substantive way on the per capita formulation.

as

$$\text{(1)} \quad \Delta X_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ t \end{pmatrix} + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-(k-1)} + \vartheta + \Phi D_t + \epsilon_t,$$

where the linear trend, following common practice, is restricted to lie within the cointegrating space, and the parameters $\vartheta$ and $\Phi$ are kept unrestricted. As a guidance for choosing the optimal lag length of the VAR, we rely on the Akaike's information criterion (AIC), likelihood ratio tests of sequential model reduction and diagnostic tests of the residuals. According to both the AIC and the series of model reduction tests, the VAR in our case should include six lags as the premise for the cointegration analysis.[2] We also note that a shorter lag length than $k = 6$ produces significant departures from white noise residuals, especially in the equation for consumption, according to the diagnostic tests.

Table 1: Trace test results for cointegration[1]

| Eigenvalue ($\lambda_i$) | $H_0$ | $H_A$ | $\lambda_{trace}$ |
|---|---|---|---|
| 0.359 | $r = 0$ | $r \geq 1$ | 84.88 [0.000] |
| 0.209 | $r \leq 1$ | $r \geq 2$ | 40.84 [0.078] |
| 0.102 | $r \leq 2$ | $r \geq 3$ | 17.62 [0.377] |
| 0.068 | $r \leq 3$ | $r = 4$ | 7.01 [0.354] |

| Diagnostics[2] | Test statistic | $p$-value |
|---|---|---|
| Vector autocorrelation 1-5 test: | F(80,187)=1.23 | 0.253 |
| Vector normality test: | $\chi^2(8) = 6.95$ | 0.542 |
| Vector heteroscedasticity test: | F(212,170)=1.08 | 0.307 |

Sample period: $1982q3-2008q3$. [1] See Johansen (1995), VAR of order 6, modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: trend (restricted), constant (unrestricted) and centered seasonal dummies (unrestricted), $r$ denotes the rank order of $\Pi = \alpha\beta'$ and $\lambda_{trace}$ is the trace statistic with $p$-value in square brackets, as reported in PcGive 14. [2] See Doornik and Hendry (2013, p. 172).

Table 1 displays results from the trace tests for cointegration and the diagnostic tests of the selected sixth order VAR. The model appears to be well-specified. The trace tests support the hypotheses of one and two cointegrating vector(s) between $c_t$, $y_t$, $w_t$ and $R_t$ at the 5 and 10 per cent significance level, respectively. We shall therefore consider both cases when testing restrictions on $\Pi = \alpha\beta'$ to discriminate between the consumption Euler equation and the Keynesian consumption function.

Assuming $r = 2$, we recall from Subsection 2.3 of the paper that the restriction $\alpha_{c1} = \alpha_{c2}$, which must be satisfied if a consumption Euler equation is true, can be tested empirically once the two cointegrating vectors are exactly identified. Hence, after imposing $\beta_{c1} = 1$ and $\beta_{w1} = 0$ in the first row of $\beta'$ and $\beta_{c2} = -1$ and

[2]Results from the AIC and the model reduction tests are available from the authors upon request.

$\beta_{y2} = 1$ in the second row of $\beta'$ to achieve exact identification, we can compare the log likelihood value of the CVAR with and without the restriction $\alpha_{c1} = \alpha_{c2}$ imposed. The associated likelihood ratio test statistic, $\chi^2(1) = 12.81$ with a $p$-value of 0.0003, points to strong rejection of the restriction. Thus consumption *is* equilibrium correcting in some way. We may therefore already at this stage of the analysis reject a consumption Euler equation and continue testing restrictions on $\Pi = \alpha\beta'$ under the assumption that $r = 1$ since the trace tests also support such an order of the rank.

Table 2 summarises main likelihood ratio tests of restrictions conditioning on the rank being one. We see that the hypotheses of homogeneity between consumption, income and wealth and the irrelevance of the trend variable are not rejected, neither jointly ($p$-value = 0.42) nor individually ($p$-values = 0.19 and 0.22). So is the joint hypothesis $\beta_y = 1$, $\beta_w = 0$ and $\gamma = 0$ ($p$-value = 0.34), the joint hypothesis $\beta_y = 1$ and $\beta_w = 0$ ($p$-value = 0.39) and the individual hypothesis $\beta_w = 0$ ($p$-value = 0.60). Moreover, a likelihood ratio test, $\chi^2(1) = 1.6$ and $p$-value = 0.21, supports reduction from Model (*ii*) to Model (*iii*) and finding homogeneity in this case is compatible with one of the predictions of the consumption Euler equation. However, the estimates of $\beta_R$ and $\alpha$ change considerably when imposing homogeneity between consumption and income only. For these reasons, and the fact that the $t$-value of the estimate of $\beta_w$ in Model (*ii*) is around 2 in magnitude, we continue focusing on the cointegrating vector which also includes the wealth variable. The estimated adjustment coefficients, except from the estimate of $\alpha_y$ ($p$-value = 0.11), are all highly significant in Model (*iv*). Accordingly, consumption, and not income, equilibrium corrects in the CVAR, which clearly contradicts an important prediction of the consumption Euler equation. Imposing the hypotheses of homogeneity, irrelevance of the trend variable and income being weakly exogenous, yields the following restricted long-run relationship

$$\widehat{eqcm}_t = c_t - 0.96y_t - 0.04w_t + 3.30R_t. \tag{2}$$

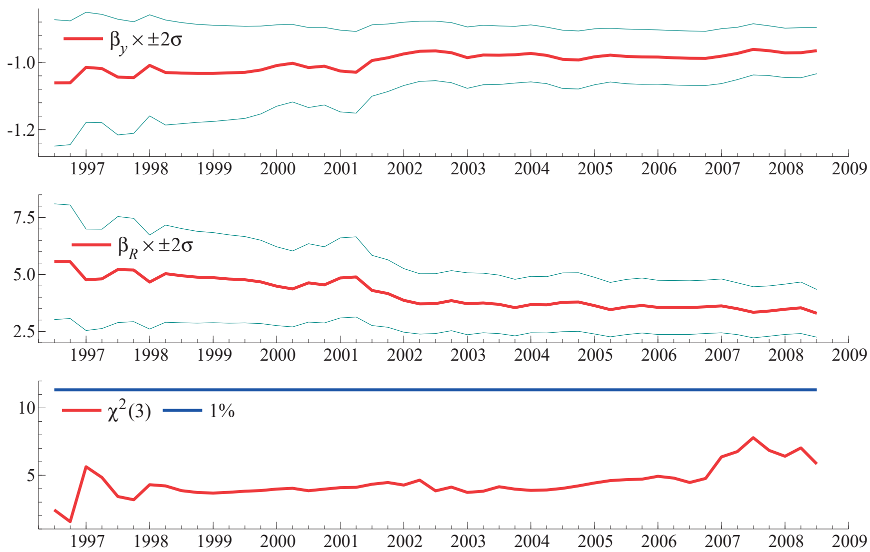
Figure 1 shows recursive estimates of the long-run coefficients in (2). It is evident that the coefficients for income, and hence for wealth, as well as for the real after tax interest rate are reasonably stable. Also, the recursive likelihood ratio tests support the joint hypothesis of $\beta_y + \beta_w = 1$, $\gamma = 0$ and $\alpha_y = 0$. A comparison with equation (2) in Jansen (2013) shows that the estimated coefficient of income (2) is somewhat higher, the estimated coefficient of wealth somewhat lower and, maybe more importantly, that the estimated coefficient of the real after tax interest rate is more than four times as high. Apart from different sample periods, a possible explanation may be that Jansen (2013) considers partial CVAR models in which the real after tax interest rate is conditioned upon at the outset. Because our findings suggest that this variable is far from being weakly exogenous, $\chi^2(1) = 15.35$ and $p$-value = 0.0001, there may be considerable loss of information of not taking into account that property in the cointegration analysis.

Table 2: Likelihood ratio test results for restrictions on $\Pi = \alpha\beta'$[1]

| |
|---|
| Model $(i)$: $\beta_c = 1$ |
| $c_t - \underset{(0.29)}{1.26} y_t - \underset{(0.05)}{0.06} w_t + \underset{(0.44)}{2.63} R_t + \underset{(0.0017)}{0.0022} t$ |
| $\hat{\alpha}_c = \underset{(0.07)}{-0.26}$, $\hat{\alpha}_y = \underset{(0.05)}{-0.09}$, $\hat{\alpha}_w = \underset{(0.09)}{-0.29}$, $\hat{\alpha}_R = \underset{(0.03)}{-0.11}$ |
| $\log L = 1170.04$ |
| |
| Model $(ii)$: $\beta_c = 1$, $\beta_y + \beta_w = 1$, $\gamma = 0$ |
| $c_t - 0.94 y_t - \underset{(0.03)}{0.06} w_t + \underset{(0.47)}{3.07} R_t$ |
| $\hat{\alpha}_c = \underset{(0.06)}{-0.21}$, $\hat{\alpha}_y = \underset{(0.05)}{-0.09}$, $\hat{\alpha}_w = \underset{(0.08)}{-0.25}$, $\hat{\alpha}_R = \underset{(0.02)}{-0.10}$ |
| $\log L = 1169.17$ |
| $\chi^2(2) = 1.75[0.42]^2$, $\chi^2(1) = 1.75[0.19]^3$, $\chi^2(1) = 1.48[0.22]^4$ |
| |
| Model $(iii)$: $\beta_c = 1$, $\beta_y = 1$, $\beta_w = 0$, $\gamma = 0$ |
| $c_t - y_t + \underset{(0.56)}{4.3} R_t$ |
| $\hat{\alpha}_c = \underset{(0.05)}{-0.14}$, $\hat{\alpha}_y = \underset{(0.04)}{-0.05}$, $\hat{\alpha}_w = \underset{(0.06)}{-0.20}$, $\hat{\alpha}_R = \underset{(0.02)}{-0.08}$ |
| $\log L = 1168.37$ |
| $\chi^2(3) = 3.35[0.34]^5$, $\chi^2(2) = 1.91[0.39]^6$, $\chi^2(1) = 0.27[0.60]^7$ |
| |
| Model $(iv)$: $\beta_c = 1$, $\beta_y + \beta_w = 1$, $\gamma = 0$, $\alpha_y = 0$ |
| $c_t - 0.96 y_t - \underset{(0.03)}{0.04} w_t + \underset{(0.52)}{3.30} R_t$, |
| $\hat{\alpha}_c = \underset{(0.06)}{-0.19}$, $\hat{\alpha}_w = \underset{(0.07)}{-0.27}$, $\hat{\alpha}_R = \underset{(0.02)}{-0.09}$ |
| $\log L = 1167.13$ |
| $\chi^2(3) = 5.84[0.12]^8$, $\chi^2(1) = 2.57[0.11]^9$ |

Sample period: $1982q3-2008q3$. [1] See Johansen (1995), VAR of order 6, $r = 1$, modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: trend (restricted), constant (unrestricted) and centered seasonal dummies (unrestricted), standard errors in parenthesis and $p$-values in square brackets. [2] $\beta_y + \beta_w = 1$, $\gamma = 0$. [3] $\beta_y + \beta_w = 1$. [4] $\gamma = 0$. [5] $\beta_y = 1$, $\beta_w = 0$, $\gamma = 0$. [6] $\beta_y = 1$, $\beta_w = 0$. [7] $\beta_w = 0$. [8] $\beta_y + \beta_w = 1$, $\gamma = 0$, $\alpha_y = 0$. [9] $\alpha_y = 0$.

Figure 1: Recursive estimates of restricted long-run coefficients




*Notes*: Sample period: 1982$q$3$-$2008$q$3.

## The extended sample period

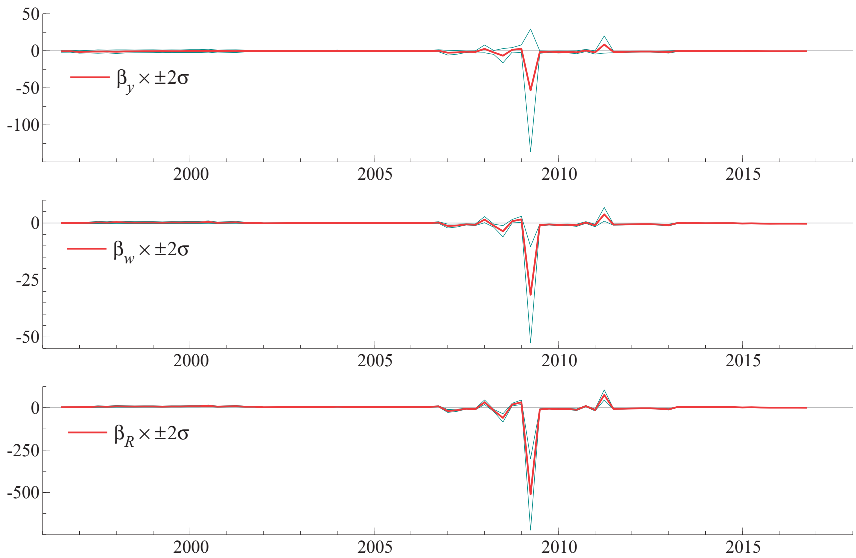
It follows from all the findings above that the data support a Keynesian type consumption function over the sample period prior to the financial crisis. When extending the sample period with 33 additional quarters, up to and including 2016$q$4, we continue to build on (1) with $k = 6$ as the underlying model.[3]

Figure 2 shows recursive estimates of the unrestricted counterpart of (2), assuming the rank to be one, over the extended sample period. Unlike the estimated coefficients over the sample period from the mid 1980s to the end of 2008, we clearly see that the estimated coefficients are unstable and reveal a significant structural break in the long-run relationship around the financial crisis. A possible explanation of the structural break is that the underlying VAR suffers from omitted variables necessary to explain a changing consumer behaviour after the financial crisis hit the Norwegian economy.

We are, therefore, as mentioned in the paper, motivated to follow Johansen *et al.* (2000) and capture a structural break in the long-run relationship by a model which takes into account the possibility of separate trends in the periods before and after the financial crisis.

[3]Again, the AIC and the series of model reduction tests support the choice of six lags in the VAR.

Figure 2: Recursive estimates of unrestricted long-run coefficients




*Notes*: Sample period: $1982q3-2016q4$.

# Appendix C. A comparative cointegration analysis

Our underlying CVAR in the case of a structural break around the financial crisis, as formulated in equation (13) in Subsection 4.1 of the paper, is

$$
\begin{aligned}
(3) \qquad \Delta X_t \quad = \quad & \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ tSD_t \end{pmatrix} + \mu SD_t + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-(k-1)} \\
& + \Phi D_t + \kappa_{2,1} ID_{2,t-1} + \ldots + \kappa_{2,k} ID_{2,t-k} + \epsilon_t.
\end{aligned}
$$

As in the preliminary cointegration analysis above, we rely on the AIC and likelihood ratio tests of sequential model reduction to determine the lag length of the VAR. Table 3 summarises results from the AIC and tests of model reduction based on (3). Again, the VAR should include six lags as the premise for the cointegration analysis.

As noted in the paper, Juselius (2006, p. 72) suggests as a rule of thumb to use a VAR with $k = 2$ in a tentative cointegration analysis. Table 4 summarises results from diagnostic tests in the cases of VAR(6) and VAR(2). We see that the VAR(2), in contrast to the VAR(6), suffers from both severe residual autocorrelation, non-normality and heteroscedasticity. Nevertheless, we conduct a comparative cointegration analysis by means of the two VAR models to see whether the analysis reported in Subsection 4.1 of the paper is robust against the choice of lag length

Table 3: AIC and model reduction tests[1]

| Model | Log$L$ | AIC |
|---|---|---|
| VAR(6) | 1586.59 | $-21.797$ |
| VAR(5) | 1568.60 | $-21.767$ |
| VAR(4) | 1544.80 | $-21.649$ |
| VAR(3) | 1508.43 | $-21.340$ |
| VAR(2) | 1468.47 | $-20.977$ |
| VAR(1) | 1431.59 | $-20.660$ |

| Model reduction | Test statistic | $p$-value |
|---|---|---|
| VAR(6) $\rightarrow$ VAR(5) | $\chi^2(16) = 35.98$ | 0.003 |
| VAR(6) $\rightarrow$ VAR(4) | $\chi^2(32) = 83.58$ | 0.000 |
| VAR(6) $\rightarrow$ VAR(3) | $\chi^2(48) = 156.34$ | 0.000 |
| VAR(6) $\rightarrow$ VAR(2) | $\chi^2(64) = 236.24$ | 0.000 |
| VAR(6) $\rightarrow$ VAR(1) | $\chi^2(80) = 310.01$ | 0.000 |
| | | |
| VAR(5) $\rightarrow$ VAR(4) | F(16,293)=2.30 | 0.003 |
| VAR(5) $\rightarrow$ VAR(3) | F(32,355)=3.12 | 0.000 |
| VAR(5) $\rightarrow$ VAR(2) | F(48,371)=3.74 | 0.000 |
| VAR(5) $\rightarrow$ VAR(1) | F(64,378)=4.13 | 0.000 |
| | | |
| VAR(4) $\rightarrow$ VAR(3) | F(16,306)=3.78 | 0.000 |
| VAR(4) $\rightarrow$ VAR(2) | F(32,370)=4.26 | 0.000 |
| VAR(4) $\rightarrow$ VAR(1) | F(48,387)=4.53 | 0.000 |
| | | |
| VAR(3) $\rightarrow$ VAR(2) | F(16,318)=4.36 | 0.000 |
| VAR(3) $\rightarrow$ VAR(1) | F(32,385)=4.47 | 0.000 |
| | | |
| VAR(2) $\rightarrow$ VAR(1) | F(16,330)=4.15 | 0.000 |

Sample period: 1982$q$3$-$2016$q$4. [1] See Doornik and Hendry (2013). Modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_t$ (restricted), $SD_t$ (unrestricted), $ID_{2t}$ (unrestricted) and centered seasonal dummies (unrestricted).

Table 4: Diagnostic test results[1]

| VAR(6) | Test statistic | $p$-value |
|---|---|---|
| Vector autocorrelation 1-5 test: | F(80,286)=1.37 | 0.034 |
| Vector normality test: | $\chi^2(8) = 9.02$ | 0.341 |
| Vector heteroscedasticity test: | F(224,266)=1.03 | 0.407 |

| VAR(2) | Test statistic | $p$-value |
|---|---|---|
| Vector autocorrelation 1-5 test: | F(80,349)=3.04 | 0.000 |
| Vector normality test: | $\chi^2(8) = 34.48$ | 0.000 |
| Vector heteroscedasticity test: | F(96,390)=1.84 | 0.000 |

Sample period: $1982q3-2016q4$. [1] See Doornik and Hendry (2013, p. 172). Modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_t$ (restricted), $SD_t$ (unrestricted), $ID_{2t}$ (unrestricted) and centered seasonal dummies (unrestricted).

of the underlying VAR. Table 5 shows trace test results for cointegration for the VAR(2). Because the underlying VAR in this case is likely to be misspecified, we only report $\lambda_{trace}$ without calculated p-values. For the same reason, we do not test formally restrictions on the $\Pi = \alpha\beta'$ with VAR(2). Assuming $r = 1$ for simplicity, the unrestricted long run relationship becomes

$$(4)\quad \widehat{eqcm}_t = c_{t-1} - 0.86y_{t-1} - 0.24w_{t-1} + 0.62R_{t-1} + 0.0017tSD_{1,t} + 0.0050tSD_{2,t}.$$

A comparison with equation (14) in the paper, shows that the estimated long run coefficients of income and wealth are quite similar. Also, the estimated deterministic parts are quite close to each other. However, the estimated long run coefficient of the real after tax interest rate drops to almost one third and the estimated adjustment coefficient of consumption ($\hat{\alpha}_c$) increases from around $-0.25$ to $-0.70$ when moving from VAR(6) to VAR(2).

We emphasise in this respect that (3) controls for a structural break around the financial crisis and that both the fourth and the fifth lag of consumption growth are strongly significant in the estimated models in Subsections 4.2 and 5.2 of the paper. For these reasons, we argue that the severe residual autocorrelation in the VAR(2) is associated with left-out dynamics rather than structural misspecification and that the cointegration results based on VAR(6) are more likely to represent the underlying data structure.

Table 5: Trace test results for cointegration with a structural break[1]

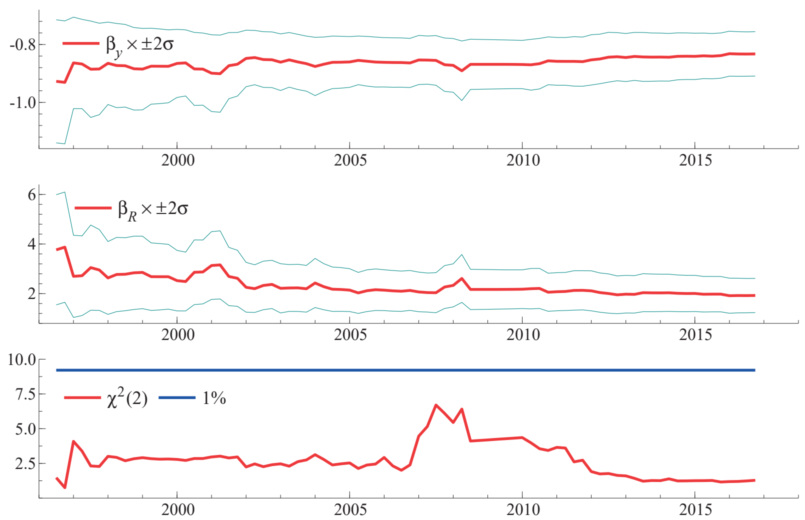
| Eigenvalue ($\lambda_i$) | $H_0$ | $H_A$ | $\lambda_{trace}$ |
|---|---|---|---|
| 0.340 | $r = 0$ | $r \geq 1$ | 104.69 |
| 0.214 | $r \leq 1$ | $r \geq 2$ | 49.77 |
| 0.078 | $r \leq 2$ | $r \geq 3$ | 18.03 |
| 0.054 | $r \leq 3$ | $r = 4$ | 7.38 |

Sample period: 1982$q$3−2016$q$4. [1] See Johansen *et al.* (2000), VAR of order 2, modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_t$ (restricted), $SD_t$ (unrestricted), $ID_{2t}$ (unrestricted) and centered seasonal dummies (unrestricted), $r$ denotes the rank order of $\Pi = \alpha\beta'$ and $\lambda_{trace}$ is the trace statistic.

# Appendix D. Recursive estimation and test statistics

In this appendix, we plot recursive estimates of the restricted long run coefficients in equation (14) in the paper along with recursive likelihood ratio tests for the joint hypothesis of $\beta_y + \beta_w = 1$ and $\alpha_y = 0$. We see that the estimated coefficients are reasonably stable and the joint hypothesis is not rejected neither before nor after the financial crisis.

Figure 3: Recursive estimates of restricted long-run coefficients




*Notes*: Sample period: 1982$q$3−2016$q$4.

# Appendix E. Detailed empirical results

In this appendix, we report detailed estimation results, diagnostic tests and plots of model fit and residuals of the whole specific system reported in Subsection 4.2 of the paper.

## Estimated specific system

The simplified dynamic models for consumption, wealth and the real after tax interest rate with respect to the stochastic variables and the broken trend in the cointegration relationship are (with estimated standard errors in parenthesis)

$$
\begin{aligned}
(5)\quad \Delta\widehat{c}_t \; = \; & -\underset{(0.09)}{0.34}\Delta c_{t-1} - \underset{(0.07)}{0.14}\Delta c_{t-2} + \underset{(0.07)}{0.45}\Delta c_{t-4} + \underset{(0.08)}{0.25}\Delta c_{t-5} + \underset{(0.09)}{0.21}\Delta y_t \\
& + \underset{(0.10)}{0.27}\Delta y_{t-3} + \underset{(0.11)}{0.24}\Delta y_{t-4} + \underset{(0.07)}{0.26}\Delta w_{t-1} - \underset{(0.07)}{0.15}\Delta w_{t-4} - \underset{(0.10)}{0.34}c_{t-1} \\
& + \underset{(0.07)}{0.26}y_{t-1} + \underset{(0.03)}{0.08}w_{t-1} - \underset{(0.11)}{0.28}R_{t-1} - \underset{(0.0001)}{0.0004}tSD_{1,t} - \underset{(0.0005)}{0.0016}tSD_{2,t} \\
& \hat{\sigma}_c = 0.018
\end{aligned}
$$

$$
\begin{aligned}
\Delta\widehat{w}_t \; = \; & \underset{(0.07)}{0.16}\Delta c_{t-3} + \underset{(0.08)}{0.34}\Delta w_{t-1} + \underset{(0.09)}{0.25}\Delta w_{t-4} - \underset{(0.27)}{0.55}\Delta R_{t-4} + \underset{(0.28)}{0.50}\Delta R_{t-5} \\
& + \underset{(0.09)}{0.07}c_{t-1} - \underset{(0.06)}{0.05}y_{t-1} - \underset{(0.02)}{0.02}w_{t-1} + \underset{(0.07)}{0.05}R_{t-1} + \underset{(0.0001)}{0.00008}tSD_{1,t} \\
& + \underset{(0.0004)}{0.0003}tSD_{2,t} \\
& \hat{\sigma}_w = 0.024
\end{aligned}
$$

$$
\begin{aligned}
\Delta\widehat{R}_t \; = \; & \underset{(0.02)}{0.07}\Delta c_{t-5} + \underset{(0.03)}{0.05}\Delta y_t - \underset{(0.08)}{0.31}\Delta R_{t-4} - \underset{(0.03)}{0.15}c_{t-1} + \underset{(0.02)}{0.11}y_{t-1} \\
& + \underset{(0.01)}{0.04}w_{t-1} - \underset{(0.04)}{0.12}R_{t-1} - \underset{(0.00004)}{0.0002}tSD_{1,t} - \underset{(0.0001)}{0.0007}tSD_{2,t} \\
& \hat{\sigma}_R = 0.007
\end{aligned}
$$

## Diagnostic tests

Table 6 shows system diagnostic test results. We see that the system in (5) is quite well specified as no serious residual autocorrelation is detected.

Table 6: Diagnostic test results[1]

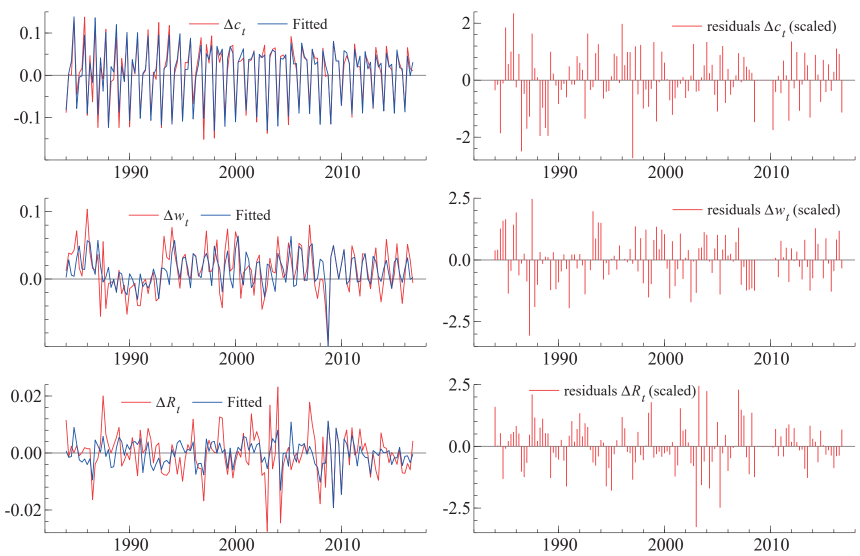
| Test | Test statistic | $p$-value |
|---|---|---|
| Vector autocorrelation 1-5 test: | F(45,285)=1.28 | 0.122 |
| Vector normality test: | $\chi^2(6) = 9.61$ | 0.142 |
| Vector heteroscedasticity test: | F(348,379)=1.32 | 0.004 |

Sample period: 1982$q$3−2016$q$4. [1] See Doornik and Hendry (2013, p. 172). Modelled variables: $c_t$, $y_t$, $w_t$ and $R_t$, deterministic variables: $tSD_t$ (restricted), $SD_t$ (unrestricted), $ID_{2t}$ (unrestricted) and centered seasonal dummies (unrestricted).

## Plots of model fit and residuals

Figure 4 plots model fit and residuals of the whole specific system. We observe that the model fit for consumption, which is of particular interest, is quite satisfactory as the fitted values are close to the historical values over the whole sample period. The estimated residuals of the dynamic consumption model are also quite small for the most part of the sample period.

Figure 4: Model fit and residuals




*Notes*: Sample period: 1982$q$3−2016$q$4.

# Appendix F. Analysis of parameter constancy of short run dynamics

Adding lagged effects of the interaction between $SD_{2,t}$ and $\Delta w_t$, denoted $SD_{2,t-i}\Delta w_{t-i}$ for $i = 1, ..., 5$, in (5) above, we obtain

$$
\begin{aligned}
(6) \quad \Delta\widehat{c}_t \;=\;& \underset{(0.09)}{-0.37}\Delta c_{t-1} - \underset{(0.07)}{0.13}\Delta c_{t-2} + \underset{(0.07)}{0.39}\Delta c_{t-4} + \underset{(0.08)}{0.23}\Delta c_{t-5} + \underset{(0.08)}{0.18}\Delta y_t \\
& + \underset{(0.10)}{0.28}\Delta y_{t-3} + \underset{(0.11)}{0.22}\Delta y_{t-4} + \underset{(0.07)}{0.32}\Delta w_{t-1} - \underset{(0.07)}{0.22}\Delta w_{t-4} - \underset{(0.09)}{0.31}c_{t-1} \\
& + \underset{(0.07)}{0.23}y_{t-1} + \underset{(0.02)}{0.07}w_{t-1} - \underset{(0.11)}{0.30}R_{t-1} - \underset{(0.0001)}{0.0004}tSD_{1,t} - \underset{(0.0004)}{0.002}\,tSD_{2,t} \\
& - \underset{(0.25)}{0.09}SD_{2,t-1}\Delta w_{t-1} + \underset{(0.23)}{0.007}SD_{2,t-2}\Delta w_{t-2} - \underset{(0.26)}{0.23}SD_{2,t-3}\Delta w_{t-3} \\
& + \underset{(0.24)}{0.58}SD_{2,t-4}\Delta w_{t-4} + \underset{(0.24)}{0.09}SD_{2,t-5}\Delta w_{t-5} \\
& \hat{\sigma}_c = 0.017
\end{aligned}
$$

$$
\begin{aligned}
\Delta\widehat{w}_t \;=\;& \underset{(0.08)}{0.26}\Delta c_{t-3} + \underset{(0.09)}{0.38}\Delta w_{t-1} + \underset{(0.09)}{0.15}\Delta w_{t-4} - \underset{(0.28)}{0.50}\Delta R_{t-4} + \underset{(0.28)}{0.55}\Delta R_{t-5} \\
& + \underset{(0.09)}{0.03}c_{t-1} - \underset{(0.06)}{0.02}y_{t-1} - \underset{(0.02)}{0.01}w_{t-1} + \underset{(0.08)}{0.03}R_{t-1} + \underset{(0.0001)}{0.00004}tSD_{1,t} \\
& + \underset{(0.0004)}{0.0001}tSD_{2,t} - \underset{(0.33)}{0.08}SD_{2,t-1}\Delta w_{t-1} - \underset{(0.31)}{0.48}SD_{2,t-2}\Delta w_{t-2} \\
& - \underset{(0.35)}{0.49}SD_{2,t-3}\Delta w_{t-3} + \underset{(0.32)}{0.03}SD_{2,t-4}\Delta w_{t-4} + \underset{(0.32)}{0.08}SD_{2,t-5}\Delta w_{t-5} \\
& \hat{\sigma}_w = 0.024
\end{aligned}
$$

$$
\begin{aligned}
\Delta\widehat{R}_t \;=\;& \underset{(0.03)}{0.07}\Delta c_{t-5} + \underset{(0.03)}{0.06}\Delta y_t - \underset{(0.08)}{0.30}\Delta R_{t-4} - \underset{(0.03)}{0.15}c_{t-1} + \underset{(0.02)}{0.11}y_{t-1} \\
& + \underset{(0.01)}{0.04}w_{t-1} - \underset{(0.04)}{0.14}R_{t-1} - \underset{(0.00004)}{0.0002}tSD_{1,t} - \underset{(0.0001)}{0.0007}tSD_{2,t} \\
& + \underset{(0.09)}{0.02}SD_{2,t-1}\Delta w_{t-1} + \underset{(0.09)}{0.03}SD_{2,t-2}\Delta w_{t-2} + \underset{(0.10)}{0.05}SD_{2,t-3}\Delta w_{t-3} \\
& - \underset{(0.09)}{0.04}SD_{2,t-4}\Delta w_{t-4} - \underset{(0.09)}{0.01}SD_{2,t-5}\Delta w_{t-5} \\
& \hat{\sigma}_R = 0.007
\end{aligned}
$$

As pointed out in the paper, only the fourth lag of $SD_{2,t}\Delta w_t$ is significant in the equation for consumption. We also see that the stochastic part as well as the broken trend component in the cointegration relationship are hardly affected in the consumption equation.

# Appendix G. Details about the estimation and testing procedure

The idea behind the estimation and testing procedure outlined in Subsection 5.1 of the paper can be explained by means of the simple model $\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1}\Delta X_{t-k+1} + \vartheta + \epsilon_t$ with restrictions $c' E_t \Delta X_{t+1} - d' X_t + d'_{-1}\Delta X_t + \ldots + d'_{-k+1}\Delta X_{t-k+2} + \vartheta = 0$. Hence, there are no restrictions on the constant term, which must be estimated. The matrices $d, d_{-1}, \ldots, d_{-k+1}$ are, however, first considered as fixed. The variables are Gaussian. From Johansen and Swensen (1999) it follows that to find the maximum likelihood estimators one has to consider a conditional equation and a marginal equation. The likelihood is of the form $L_{1\cdot 2,max}(d, d_{-1}, \ldots, d_{-k+1})L_{2,max}(d, d_{-1}, \ldots, d_{-k+1})$. The marginal equation takes the form

$$c'\Delta X_t = d' X_{t-1} - d'_{-1}\Delta X_{t-1} - \ldots - d'_{-k+2}\Delta X_{t-k+2} - d'_{-k+1}\Delta X_{t-k+1} + c'\vartheta + c'\epsilon_t.$$

The maximal value of the marginal likelihood can therefore be found by regressing $c'\Delta X_t - d' X_{t-1} + d'_{-1}\Delta X_{t-1} + \cdots + d'_{-k+1}\Delta X_{t-k+1}$ on $c'1$, where 1 is a $p \times 1$ vector, so $L_{2,max}(d, d_{-1}, \ldots$
$, d_{-k+1})$ has a closed form. Let $c_\perp$ be a $p \times p - q$ matrix so $(c, c_\perp)$ has full rank and $c'c_\perp = 0$. and let $\bar{c} = c(c'c)^{-1}$. The conditional equation then takes the form

$$\begin{aligned}
c'_\perp \Delta X_t &= \eta \xi' \bar{d}'_\perp X_{t-1} \\
&- \rho(c'\Delta X_{t-1} - d' X_{t-1} + d'_{-1}\Delta X_{t-1} + \ldots + d'_{-k+1}\Delta X_{t-k+1} - c'\vartheta) \\
&+ \theta\ (d'd)^{-1} d' X_{t-1} + c'_\perp \Gamma_1 \Delta X_{t-1} + \ldots + c'_\perp \Gamma_{k-1}\Delta X_{t-k+1} + c'_\perp \vartheta + u_t,
\end{aligned}$$

where $u_t = (c'_\perp - \rho c')\epsilon_t$. For $d$ and $d_{-1}, \ldots, d_{-k+1}$ fixed, the maximal values of the likelihood can be computed by reduced rank regression. The matrices $\eta$, $\xi$, $\rho = c'_\perp \Omega c (c'\Omega c)^{-1}$ and $\theta$ have dimensions $(p-q) \times (r-q)$, $(p-q) \times (r-q)$, $(p-q) \times q$ and $(p-q) \times q$, respectively.

Now we want to consider maximization over $d$, $d_{-1}, \ldots, d_{-k+1}$. Since these quantities occur in both the marginal and conditional equations the product $L_{1\cdot 2,max}$ $(d, d_{-1}, \ldots, d_{-k+1})L_{2,max}(d, d_{-1}, \ldots, d_{-k+1})$ must be considered. Using a generic numerical optimization procedure is an option, but the number of parameters quickly gets large. We therefore propose another procedure. If $d$ is fixed, new values for $d_{-1}, \ldots, d_{-k+1}$ can be found by regressing $c'\Delta X_t - d' X_{t-1}$ on $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$ and $c'1$. Because $d_{-1}, \ldots, d_{-k+1}$ also occur in the conditional equation it may be that also restrictions arising from this part must be taken into account. Reformulating the conditional equation as

$$\begin{aligned}
c'_\perp \Delta X_t &= \eta \xi' \bar{d}'_\perp X_{t-1} \\
&- \rho(c'\Delta X_{t-1} - d' X_{t-1}) + \theta\ (d'd)^{-1} d' X_{t-1} \\
&+ (c'_\perp \Gamma_1 - \rho d'_{-1})\Delta X_{t-1} - \ldots + (c'_\perp \Gamma_{k-1} - \rho d'_{-k+1})\Delta X_{t-k+1} \\
&+ (c'_\perp - \rho c')\vartheta + u_t,
\end{aligned}$$

one can see that there are no such constraints since $(c_\perp'\Gamma_1 - \rho d_{-1}'), \ldots, (c_\perp'\Gamma_{k-1} - \rho d_{-k+1}')$ and $(c_\perp' - \rho c')\vartheta$ vary freely and the parameters can be estimated by reduced rank regression.

Thus $L_{1\cdot 2,max}(d, d_{-1}, \ldots, d_{-k+1})L_{2,max}(d, d_{-1}, \ldots, d_{-k+1})$ is concentrated and depends only on the values in $d$. The maximum value can be found using a generic optimization procedure. As there are no restrictions on $d, d_1, \ldots, d_{k-1}$ this maximum will be the same as for the reduced rank VAR. Restrictions on $d_{-1}, \ldots, d_{-k+1}$, for instance $d_{-k+1} = d_{-k+1}^0$, can be treated as above, but regressing $c'\Delta X_t - d'X_{t-1} + d_{k-1}^0 \Delta X_{t-k+1}$ on $d_{-1}'\Delta X_{t-1}, \ldots, d_{-k+2}'\Delta X_{t-k+2}$ and $c'1$. The conditional equation takes the form

$$\begin{aligned}
c_\perp'\Delta X_t &= \eta\xi'\bar{d}_\perp' X_{t-1} \\
&- \rho(c'\Delta X_{t-1} - d'X_{t-1} + d_{k-1}^0 \Delta X_{t-k+1}) + \theta\ (d'd)^{-1}d'X_{t-1} \\
&+ (c_\perp'\Gamma_1 - \rho d_{-1}')\Delta X_{t-1} + \ldots + (c_\perp'\Gamma_{k-2} - \rho d_{-k+2}')\Delta X_{t-k+2} \\
&+ c_\perp'\Gamma_{k-1}\Delta X_{t-k+1} + (c_\perp' - \rho c')\vartheta + u_t
\end{aligned}$$

where the parameters can be estimated by reduced rank regression.

# Appendix H. Additional results on conditional expectations

We ask in Subsection 5.2 of the paper if the habit formation type behaviour in equation (19) is a consequence of assuming that the agents know the financial shift has occurred in the way expectations are formulated in the model. However, estimating the specific consumption model with a sample period ending in 2008$q$3 rather than in 2016$q$4, we obtain

$$\begin{aligned}
\widehat{E_t\Delta c_{t+1}} &= -0.26(1.0, -0.83, -0.17, 1.94, 0.0009, 0.0051)\begin{pmatrix} X_t \\ tSD_{t+1} \end{pmatrix} \\
&\quad -\underset{(0.08)}{0.39}\Delta c_t - \underset{(0.07)}{0.25}\Delta c_{t-1} + \underset{(0.07)}{0.53}\Delta c_{t-3} + \underset{(0.08)}{0.30}\Delta c_{t-4} \\
&\quad +\underset{(0.09)}{0.16}\Delta y_{t-3} + \underset{(0.07)}{0.22}\Delta w_t - \underset{(0.07)}{0.19}\Delta w_{t-3},
\end{aligned}$$

where $X_t = (c_t, y_t, w_t, R_t)'$ and $tSD_{t+1} = (tSD_{1,t+1}, tSD_{2,t+1})'$. We see that the lag structure of consumption growth is fairly similar to that in equation (19) for the whole sample period. Hence, the estimated habit formation type behaviour does not seem to be a consequence of the way expectations about the financial shift are formulated in the model.

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# 3 Chapter 3

## Inflation dynamics in a small open economy[21]

Co-authored with Ådne Cappelen and Anders Rygh Swensen.
Published in *The Scandinavian Journal of Economics*.

[21] Equation (11) in the article contains a typo and should read $c_1'\Pi = -(c_0' + c_{-1}' + c_{-2}' - c_1')$, $c_1'\Phi_1 = -\psi_0$.

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# Inflation Dynamics in a Small Open Economy*

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## Abstract

We evaluate the empirical performance of forward-looking models for inflation dynamics in a small open economy. Using likelihood-based testing procedures, we find that the exact formulation is at odds with Norwegian data. Moreover, some of the parameters in the model are not well identified. We also find that the inexact formulation is not rejected statistically using a test based on a minimum distance method. However, confidence regions also reveal an identification problem with this model. Instead, we find a well-specified backward-looking model with imperfect competition underlying the price setting, which is a model that outperforms an alternative forward-looking model in-sample. The backward-looking model also forecasts somewhat better than the alternative forward-looking model, during and after the recent financial crisis.

*Keywords*: Backward-looking; cointegrated vector autoregressive models; equilibrium correction models; forward-looking; likelihood-based methods; minimum distance method

*JEL classification*: *C*51; *C*52; *E*31; *F*31

## I. Introduction

Forward-looking models – based on rational expectations, optimizing agents, and imperfect competition in markets for goods – have long played a central role in understanding inflation dynamics in the economics profession and in central banks conducting inflation targeting. Since the influential papers by Roberts (1995), Fuhrer and Moore (1995), Galí and

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Gertler (1999), Galí *et al.* (2001), and Sbordone (2002), a large number of studies have been devoted to testing the role of expectations in inflation dynamics based on data from both closed and open economies; see Mavroeidis *et al.* (2014) and An and Schorfheide (2007) for reviews of the literature. Studies differ with respect to the data used, the sample period studied, and the econometric methods applied. The supportive evidence on forward-looking behavior in price formation is rather mixed.

In this paper, we follow Mavroeidis *et al.* (2014), among several others, and we evaluate the empirical performance of forward-looking models based on Norwegian data and using a limited-information approach. Building on Sbordone (2002), our forward-looking models relate current inflation to expected future inflation and the difference between the actual price and the steady-state value of levels as a theory-consistent forcing variable. The steady-state value is specified as a mark-up over marginal costs, which, in turn, are determined by costs of both labor and imported intermediate goods along the lines of the open economy models in McCallum and Nelson (1999), Kara and Nelson (2003), and Batini *et al.* (2005). We contribute to the empirical literature by studying both the exact formulation, in the sense of Hansen and Sargent (1991), and the inexact formulation, in which a stochastic error term is included in the model. For the exact formulation, we employ the likelihood-based testing procedures suggested by Johansen and Swensen (1999, 2008). Because a similar treatment of the inexact formulation is more complicated to handle, we rely on a test based on a minimum distance approach along the lines of Sbordone (2002) and Magnusson and Mavroeidis (2010). Consequently, we are able to shed some light on the importance of introducing a stochastic error term to the empirical model, an econometric issue that is often neglected in the literature. Unlike most related studies (e.g., Galí and Gertler, 1999; Galí *et al.*, 2001; Batini *et al.*, 2005), we pay particular attention to time series properties of the variables involved, and the possible existence of unit roots, and we search for statistically well-specified underlying models as premises for valid statistical inference on the forward-looking models (cf. Mavroeidis *et al.*, 2014). Another issue that has been little addressed in the literature is forecasting performance. Recently, Rumler and Valderrama (2010) and King and Watson (2012) have discussed this issue. King and Watson (2012) find a large gap between inflation predicted by the Smets and Wouters (2007) forward-looking model for US and actual inflation. The gap can only be closed by assuming large and exogenous mark-up shocks. They recommend devoting more attention to detailed analysis of the structural inflation equation in order to detect imperfect specifications. In our study, we also compare and contrast specifications of a reduced-form forward-looking model with a backward-looking model counterpart as two competing models of inflation dynamics, both in-sample and out-of-sample in a forecasting competition.

Our empirical investigation, which is based on a dataset not used in this setting earlier, produces several noteworthy findings. First, we establish a well-specified empirical counterpart to the theory-consistent link between consumer prices and marginal costs. Secondly, we demonstrate that the exact formulation of the forward-looking model is at odds with the data. That is, the rational expectations hypothesis is not rejected statistically. However, when only economically meaningful parameters are allowed for, the model is not supported by the data. In addition, a plot of the likelihood surface reveals that some of the parameters might not be well-identified. Using alternative methods to those of Kurmann (2007), we also discuss the inexact formulation of the forward-looking model, and we find no indication that this model yields radically different results. In particular, the identification problem seems also to be present in this model. Thirdly, we establish a well-specified competing backward-looking model of inflation dynamics in a sample containing a major monetary policy regime shift. Finally, we find that the backward-looking model forecasts somewhat better than a reduced-form forward-looking model during and after the recent financial crisis.

The rest of the paper is organized as follows. In Section II, we outline our forward-looking models. In Section III, we describe the data. In Section IV, we report findings from the cointegration analysis. In Section V, we report on the various tests of the forward-looking models. In Section VI, we evaluate the empirical performance of the reduced-form forward-looking model compared with the backward-looking model, and we conduct a forecasting competition between these models. We conclude in Section VII.

## II. Theoretical Framework

Several approaches have been suggested in the literature for the introduction of open economy features into the price formation of firms. One approach treats imports as substitutable final consumer goods for domestically produced goods, and it assumes that the representative firm operates in imperfectly competitive markets facing regular downward sloping demand curves, (e.g., Bårdsen *et al.*, 2005, Chapter 8.7; Galí and Monacelli, 2005). Profit maximization then implies that prices are set as a mark-up over marginal costs, where the mark-up depends on the relative prices of domestic and imported consumer goods. A second approach treats imports as intermediate goods in production rather than as final consumer goods (e.g., McCallum and Nelson, 1999; Kara and Nelson, 2003; Batini *et al.*, 2005). Profit maximization then implies that prices of imports become important determinants of marginal costs and not of the mark-up. A third approach treats imports as intermediate inputs as well as final consumer goods. Hence, prices of imports would be accounted for through both the mark-up and the marginal costs. A disaggregated model would then be necessary.

We rely on the second approach in this paper, and we treat imports as intermediate goods in production. Our main argument is that imports are rarely imported by consumers themselves, but rather by the wholesale and retail trade sector, and they are used in combination with other inputs to supply domestic consumers with goods. These imported goods are usually both intermediate goods and final consumer goods. For instance, an imported agricultural product can be bought in a shop by the consumer, as part of a meal in a restaurant, or as an input to the domestic food industry. Petrol is used both by consumers as a final good and by firms as input in production. In any case, the wholesale and retail trade sector adds trade margins and domestic cost components to the costs of imports to make up the consumer prices of imported goods.

Assuming a representative profit-maximizing firm facing an isoelastic, downward-sloping demand curve and a Cobb–Douglas production function in labor and imports of intermediate inputs, we have

$$p_t^* = m_0 + \psi_1 ulc_t + \psi_2 uic_t, \tag{1}$$

where lower-case letters indicate natural logarithms and $p_t^*$, $m_0$, $ulc_t$, and $uic_t$ are the steady-state value of the price level, the constant mark-up, the unit labor costs, and the unit import costs of intermediates, respectively.[1] The price equation in (1) is homogeneous of degree one in the factor prices (i.e., $\psi_2 = (1 - \psi_1)$).

The various forward-looking models proposed in the literature on the new Keynesian Phillips curve share essentially the same semi-structural form, but differ with respect to the specific underlying pricing behavior, including the forward-looking linear quadratic adjustment cost model of Rotemberg (1982), the models of staggered contracts developed by Taylor (1979, 1980), and the sticky price model of Calvo (1983). We build on equation (2.9) of Sbordone (2002), and we specify our forward-looking models as

$$\Delta p_t = \delta E_t \Delta p_{t+1} - \lambda(p_t - p_t^*), \tag{2}$$

where $\Delta p_t = p_t - p_{t-1}$ is current inflation and $E_t \Delta p_{t+1}$ is expected inflation one period ahead, conditional on information available at time $t$. Now,

[1] A full derivation of equation (1) is shown in Appendix A. In the literature on the closed-economy new Keynesian Phillips curve, it is common to assume that producers face isoelastic demand curves, so that the mark-up is a constant (e.g., Galí and Gertler, 1999; Galí *et al.*, 2001).

inserting equation (1) into equation (2) yields a forward-looking model of inflation dynamics

$$\Delta p_t = \delta E_t \Delta p_{t+1} - \lambda eqcm_t + \psi_0, \tag{3}$$

where

$$eqcm_t = p_t - \psi_1 ulc_t - \psi_2 uic_t \tag{4}$$

and $\psi_0 = \lambda m_0$. If our specification of price formation in a small open economy, as shown in equation (1), is supported by the data, then the equilibrium correction term in equation (4) with the homogeneity restriction imposed should be a stationary variable. This is a testable implication that we return to in Section VI. Inspired by Galí and Gertler (1999) among others, we specify the so-called hybrid version of equation (3) as

$$\Delta p_t = \gamma_f E_t \Delta p_{t+1} + \gamma_b \Delta p_{t-1} - \lambda eqcm_t + \psi_0, \tag{5}$$

which allows for some firms to be backward-looking according to a "rule of thumb" hypothesis where past inflation drives current inflation. The semi-structural parameter spaces $0 \leq \gamma_f, \gamma_b \leq 1$, and $0 \leq \lambda$ are required in order to provide an admissible economic interpretation of equation (5). The hybrid forward-looking model reduces to its non-hybrid version when $\gamma_b = 0$. We note that equation (5) can be reparametrized as

$$\Delta p_t = \varphi_1 E_t \Delta p_{t+1} + \varphi_2 \Delta p_{t-1} + \varphi_3 \Delta ulc_t + \varphi_4 \Delta uic_t - \varphi_5 eqcm_{t-1} + \varphi_6, \tag{6}$$

where $\varphi_1 = \gamma_f/(1+\lambda)$, $\varphi_2 = \gamma_b/(1+\lambda)$, $\varphi_3 = \lambda\psi_1/(1+\lambda)$, $\varphi_4 = \lambda\psi_2/(1+\lambda)$, $\varphi_5 = \lambda/(1+\lambda)$, and $\varphi_6 = \psi_0/(1+\lambda)$. Hence, equation (6) becomes a backward-looking Phillips curve written in the usual equilibrium correction form when $\gamma_f = 0$.

As pointed out by Mavroeidis *et al.* (2014), one can, in principle, allow for any number of lagged inflation terms in the model if the objective is to nest traditional Phillips curves. In addition to the backward-looking rule of thumb, lagged inflation terms could be motivated by staggered relative wage contracts and the indexation of prices to past inflation (e.g., Fuhrer and Moore, 1995; Christiano *et al.*, 2005). We return to these issues in the discussion of our empirical findings in Section VI.

## III. Data

The empirical analysis is based on quarterly, seasonally unadjusted data that span the period 1982Q1–2011Q4, from which data taken from the periods 1982Q1–2005Q4 and 2006Q1–2011Q4 are used for estimation and out-of-sample forecasting, respectively. Mavroeidis (2004) concludes that the estimates of forward-looking models are less reliable when the sample covers periods in which inflation has been under effective policy control. The starting point of our estimation period is thus motivated by the fact that the 1970s and the early 1980s were characterized by massive governmental price controls in Norway (see Bowitz and Cappelen, 2001). If the expectation term in the forward-looking model relationship is the most important factor determining the correlation between exchange rate movements and inflation, we would expect the relationship to depend closely on the monetary policy regime. We explored this hypothesis by ending the estimation period in 2001Q1 rather than in 2005Q4, as monetary policy in Norway changed fundamentally from exchange rate targeting to inflation targeting in late March 2001; see Boug *et al.* (2006) for details. It turns out, however, that the results from the cointegration analysis in the next section are virtually the same for the two ending points. By extending the estimation sample by 24 quarters for out-of-sample forecasting, we shed light on any change there might be in the link between exchange rate movements and domestic inflation following the financial crisis in 2008 and onwards. The out-of-sample forecasting ends in 2011Q4 because available data on marginal costs for the years 2012 and 2013 are only preliminary figures from the national accounts.

We measure quarterly inflation by the official consumer price index (CPI) rather than by the gross domestic product (GDP) deflator often used in the new Keynesian Phillips curve literature. Prices set by agents in the economy are based on gross output and not on value added. Deflators based on value added are typically residuals in the national accounts, particularly those that follow the recommended principle of double-deflating, in which different deflators are used for gross output and material inputs. Hence, the GDP deflator is less related to micro price-setting behavior than the CPI. However, there might be a problem with using the CPI if indirect taxes change in a systematic way, although this also affects the GDP deflator. As noted below, we adjust for one episode of indirect tax changes in the sample period. In line with Batini *et al.* (2005), we employ the deflator for total imports as a proxy for unit import costs, whereas total labor costs relative to value added in the private mainland economy serve as a proxy for unit labor costs. The details can be found in Appendix D. Figure 1 shows the log of the CPI $p_t$, the log of unit import costs $uic_t$, and the log of unit labor costs $ulc_t$, together with the inflation rates $\Delta p_t$ over the sample period.

During the estimation period, consumer price inflation shows rather large changes in the quarters 1986Q3, 1996Q1, 2001Q3, 2003Q1, and 2003Q2. These changes are associated with a 12 percent devaluation of the Norwegian currency in May 1986, a reduction in indirect tax rates during the first quarter of 1996, a reduction in the VAT rate on food from 24 percent to 12 percent in July 2001, and a substantial increase and decrease in electricity

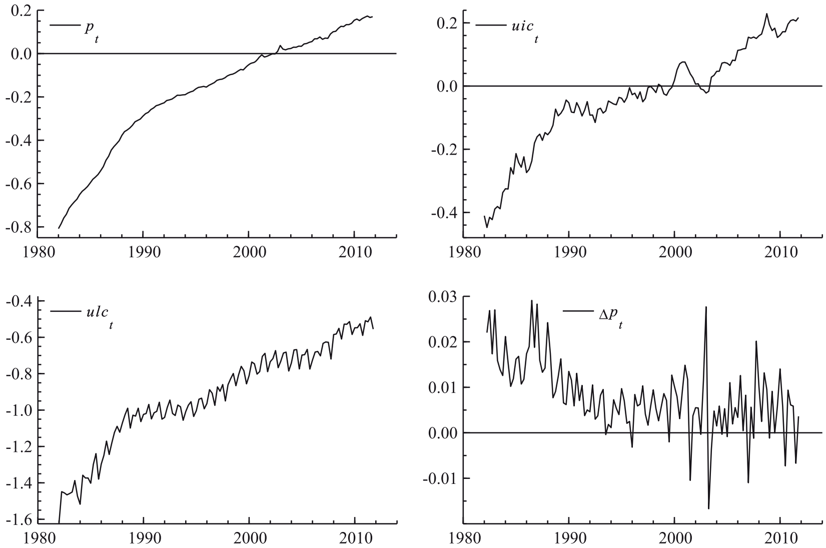


*Fig. 1.* Time series for $p_t$, $uic_t$, $ulc_t$, and $\Delta p_t$

prices during the first and second quarters of 2003, respectively. The fact that inflation increased considerably in the third quarter and not in the second quarter of 1986 is because of delayed pass-through from exchange-rate changes to import prices and consumer prices (see Boug *et al.*, 2013). Fluctuations in electricity prices are, to a large extent, related to natural causes (e.g., temperature) and not much to immediate economic phenomena, as electricity is mainly based on hydroelectric power in Norway. We control for the above-mentioned episodes in the empirical analysis, using impulse dummies labeled $D86Q3$, $D96Q1$, $D01Q3$, $D03Q1$, and $D03Q2$. Consumer price inflation also shows some huge fluctuations during the forecasting period, especially in the quarter 2007Q1 and in the years 2008 and 2009, which can likely be attributed to the substantial fall in electricity prices and the large movements in the exchange rate during the recent financial crisis, respectively. We further note that the time series exhibit a clear upward trend, but with no apparent mean-reverting property, suggesting that $p_t$, $uic_t$, and $ulc_t$ are non-stationary $I(1)$ series. Therefore, a reduced-rank vector autoregressive (VAR) model is a candidate as an empirical model.

## IV. Cointegration Analysis

We adopt the cointegration rank test suggested by Johansen (1995, p. 167) to find an empirical counterpart of equation (4). The point of departure of

Table 1. *Tests for cointegration rank*

| | $\lambda_i$ | $\lambda_{\text{trace}}$ | $\lambda^a_{\text{trace}}$ |
|---|---|---|---|
| $r = 0$ | 0.262 | 47.22 (0.016)* | 42.65 (0.052) |
| $r \leq 1$ | 0.136 | 18.95 (0.290) | 17.12 (0.414) |
| $r \leq 2$ | 0.056 | 5.36 (0.555) | 4.84 (0.626) |

*Notes:* $r$ denotes the cointegration rank, $\lambda_i$ are the eigenvalues from the reduced-rank regression (see Johansen, 1995), and $\lambda_{\text{trace}}$ and $\lambda^a_{\text{trace}}$ are the trace statistics without and with degrees of freedom adjustments, respectively. The $p$-values in parentheses, which are reported in OxMetrics, are based on the approximations to the asymptotic distributions derived by Doornik (1998). It should be noted that inclusion of impulse dummies in the VAR affects the asymptotic distribution of the reduced-rank test statistics. Thus, the critical values are only indicative. The asterisk * denotes rejection of the null hypothesis at the 5 percent significance level.

the $I(1)$ analysis and the tests that follow is a $p$-dimensional VAR of order $k$ written as

$$X_t = A_1 X_{t-1} + \cdots + A_k X_{t-k} + \Phi_0 D_t + \Phi_1 + \Phi_2 t + \varepsilon_t, \tag{7}$$

where $X_t = (p_t, ulc_t, uic_t)'$, $D_t$ includes seasonal dummies labeled $SD_{it}$ $(i = 1, 2, 3)$ and the impulse dummies $D86Q3$, $D96Q1$, $D01Q3$, $D03Q1$, and $D03Q2$ as described above, $t$ is a linear deterministic trend, and $\varepsilon_{k+1}, \ldots, \varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) covariance matrix $\Omega$. The initial observations of $X_1, \ldots, X_k$ are kept fixed. We follow common practice and restrict the linear trend to lie within the cointegrating space, whereas the deterministic components $D_t$ and $\Phi_1$ are kept unrestricted in equation (7). If $X_t$ is $I(1)$, then the presence of cointegration implies $0 < r < p$, where $r$ denotes the rank or the number of cointegrating vectors of the impact matrix $\Pi = A_1 + \cdots + A_k - I$. The null hypothesis of $r$ cointegrating vectors can be formulated as $H_0$: $\Pi = \alpha\beta'$, where $\alpha$ and $\beta$ are $p \times r$ matrices, $\beta' X_t$ comprises $r$ cointegrating $I(0)$ linear combinations, and $\alpha$ contains the adjustment coefficients.

We find that $k = 3$ is the appropriate choice of lag length to arrive at a model with no serious misspecification in the residuals.[2] Table 1 shows the findings from applying the cointegration rank test to the data based on the VAR of order three.

We observe that the rank should be set to unity at the 5 percent significance level (albeit the $\lambda^a_{\text{trace}}$ statistics is a borderline case), indicating the existence of one cointegration relationship between consumer prices, unit labor costs, and unit import costs. It might be worth pointing out that starting with a VAR with an unrestricted constant only yields the same result,

[2] The preferred VAR includes an additional impulse dummy labeled $D84Q1$ to mop up a relatively large residual in 1984Q1 in the $ulc_t$ equation. Without $D84Q1$, the $ulc_t$ equation suffers from severe residual autoregressive heteroskedasticity. The cointegration analysis below is not significantly affected by any of the impulse dummies included in the preferred VAR.

$r = 1$. The $p$-values, not adjusted for degrees of freedom (dof), corresponding to $r = 0$, $r \leq 1$, and $r \leq 2$ are 0.00, 0.12, and 0.09, respectively.

The null hypothesis that the linear trend, $\Phi_2 = 0$, can be eliminated from the VAR, assuming the rank to be unity, is not rejected by a likelihood ratio test. The $p$-value is 0.388 based on a $\chi^2$ approximation with one degree of freedom. The corresponding maximized value of the 2 log likelihood, which is used in Section V, is 2536.72. Imposing a further restriction of homogeneity between $p_t$, $ulc_t$, and $uic_t$ entails a reduction in the value of the 2 log likelihood of 0.0097, which corresponds to a $p$-value of 0.921 using the same $\chi^2$ approximation. We obtain the following empirical counterpart of equation (4) when the restrictions of homogeneity between $p_t$, $ulc_t$, and $uic_t$ and no linear trend in $\beta$ are imposed:

$$\widehat{eqcm}_{1,t} = p_t - 0.641 ulc_t - 0.359 uic_t. \tag{8}$$

The issue of joint weak exogeneity of $ulc_t$ and $uic_t$ is more debatable. Here, the $-2$ log likelihood ratio value is 7.909. The $p$-value based on approximating the null distribution with a $\chi^2$ distribution with two degrees of freedom is 0.02. However, investigating weak exogeneity more closely, using both parametric and non-parametric bootstrap methods, reveals that the asymptotic approximations are not accurate in our case. A bootstrap of the likelihood ratio test, as in Omtzigt and Fachin (2007), using the estimated values of the VAR coefficients and not imposing weak exogeneity and resampling the residuals, yields a $p$-value of 0.515. The outcome of a non-parametric bootstrap is similar. We conclude that the cointegration vector enters the $\Delta p_t$ equation only. Imposing, in addition, weak exogeneity of both $ulc_t$ and $uic_t$ yields

$$\widehat{eqcm}_{2,t} = p_t - 0.604 ulc_t - 0.396 uic_t. \tag{9}$$

We see that $\widehat{eqcm}_{1,t}$ and $\widehat{eqcm}_{2,t}$ are quite similar, which provides further evidence that the restriction of weak exogeneity of both $ulc_t$ and $uic_t$ can be justified. Figure 2 depicts time series for $p_t$ and $\hat{p}_t$ based on equation (8) over the sample period.

It is evident that $\hat{p}_t$ matches $p_t$ rather closely, both in-sample and out-of-sample. Hence, we interpret equation (8) as a long-run consumer price equation that corresponds well with the theory of mark-up pricing, and takes into account that, for a small open economy such as the Norwegian economy, features of an open economy such as import prices are expected to matter. The estimates in equation (8) are in line with previous findings based on Norwegian data (e.g., Bårdsen *et al.*, 2005, p. 182). Nearly four decades ago, Aukrust (1977, p. 123) pointed out that the total direct effect on consumer prices that can be expected, under Norwegian conditions, from a proportionate increase in all import prices can be set at 0.33 percent.

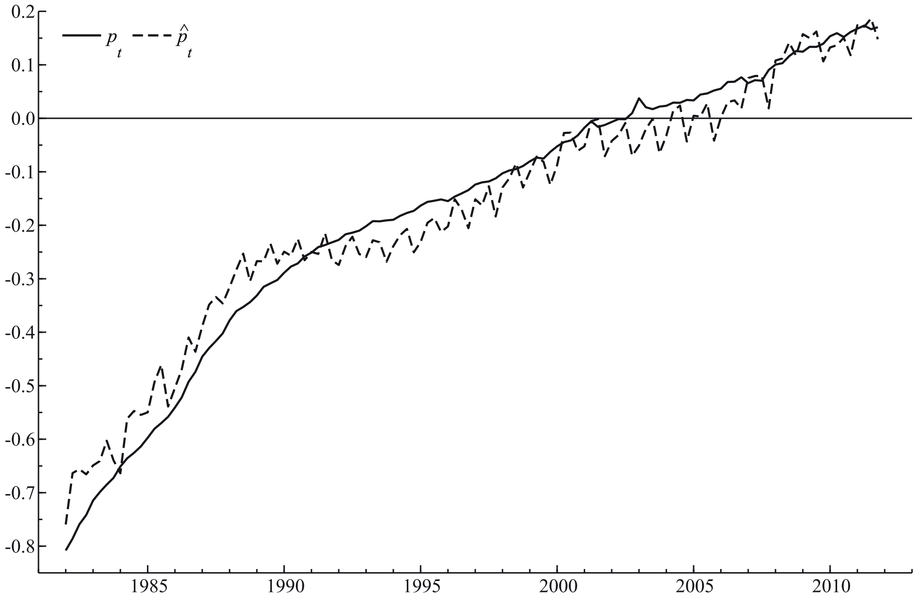


*Fig. 2.* Time series for $p_t$ and $\hat{p}_t$ based on equation (8)

Hence, equation (8) is also in line with the Scandinavian model of inflation (cf. Lindbeck, 1979).

## V. Tests of Forward-Looking Models

An important econometric issue when testing the forward-looking model concerns whether the model is specified in its exact or inexact form by introducing a stochastic error term $u_t$. In general, the absence of an unobserved disturbance term ($u_t = 0$) can be a restrictive and non-trivial assumption, as there are several justifications for why such a term could be included in the model (e.g., Sbordone, 2005). To shed light on the importance of the disturbance term in our empirical case, we evaluate both versions of the model in this paper. However, as demonstrated by Boug *et al.* (2010) in the case of the new Keynesian Phillips curve within a bivariate VAR, the exact version is algebraically less involved and produces much simpler rational expectations restrictions than what follow from the inexact version under the assumption that $u_t$ is a sequence of innovations (i.e., $E_t(u_{t+1}) = 0$). Hence, the numerical treatment of the exact model using likelihood-based methods is also much simpler than the inexact model. When a trivariate VAR is the underlying model, as is the case in the present study, the numerical treatment of the inexact model is even more complicated to handle using likelihood-based methods. As a consequence, we employ the likelihood-based testing procedures suggested by Johansen and Swensen (1999, 2008) for the exact model and a test based on a minimum distance approach, along the lines of Sbordone (2002) and Magnusson and Mavroeidis (2010), for the inexact model.

It is worth noting that, by considering the exact and inexact models as submodels of a reduced-rank VAR, we avoid cases like those described by Beyer and Farmer (2007). In a single equation framework, they demonstrated that exact and inexact models exist that have the same likelihood and are hence empirically indistinguishable. As we shall see, the restrictions from an exact model and an inexact model on the coefficients of the VAR are quite different, and thus also the likelihoods of the two submodels.

### *The Exact Version*

The basic idea behind the procedure suggested by Johansen and Swensen (1999, 2008) is to start with a well-specified VAR model and, using the likelihood criterion, to test the implications of the hybrid forward-looking model on the coefficients of the VAR. As explained in the previous section, a reduced-rank VAR model with three lags and a constant, but with no linear term, passed a test for homogeneity. Therefore, we explore this model further.

Expressing equation (5) on level form, taking the homogeneity restriction into account, yields

$$\begin{aligned}\gamma_f E_t[p_{t+1}] - (1+\gamma_f)p_t + (\gamma_b+1)p_{t-1} - \gamma_b p_{t-2}\\ - \lambda(p_t - \psi_1 ulc_t - \psi_2 uic_t) + \psi_0 = 0,\end{aligned}$$

or

$$\gamma_f E_t[p_{t+1}] - (1+\gamma_f)p_t + (\gamma_b+1)p_{t-1} - \gamma_b p_{t-2} - \lambda eqcm_t + \psi_0 = 0,$$

where $\psi_1 + \psi_2 = 1$. With $c_1 = (\gamma_f, 0, 0)'$, $c_0 = (-1-\gamma_f-\lambda, \lambda\psi_1, \lambda\psi_2)'$, $c_{-1} = (\gamma_b+1, 0, 0)'$, and $c_{-2} = (-\gamma_b, 0, 0)'$, this can be written as

$$c_1' E_t[X_{t+1}] + c_0' X_t + c_{-1}' X_{t-1} + c_{-2}' X_{t-2} + \psi_0 = 0.$$

The fitted VAR model contains six impulse dummies and three seasonal dummies in addition to a constant. Denote the coefficients of the constant term as $\Phi_1$ and the coefficients of the rest of the deterministic terms as $\Phi_0$, so that the deterministic part of the VAR model can be written as $\Phi_0 D_t + \Phi_1$. The rational expectation hypothesis will imply restrictions also on these coefficients. Because we are not focusing on the properties that the dummies capture, we simply drop those restrictions, which, as explained in Boug *et al.* (2006), amounts to formulating the model as

$$c_1' E_t[X_{t+1} - \Phi_0 D_{t+1}] + c_0' X_t + c_{-1}' X_{t-1} + c_{-2}' X_{t-2} + \psi_0 = 0. \tag{10}$$

Table 2. *Likelihood ratio tests of the exact forward-looking model: nested models*

| Model | $2 \log L$ | $-2 \log LR$ | dof | $p$-value |
|---|---|---|---|---|
| CVAR without homogeneity restriction | $2536.72^{a}$ | | | |
| CVAR with homogeneity restriction | $2536.71^{a}$ | 0.01 | 1 | 0.92 |
| Exact hybrid model | $2535.15^{b}$ | 1.56 | 4 | 0.82 |
| Exact non-hybrid model | $2529.64^{b}$ | 5.60 | 1 | 0.02 |

*Notes:* Maximal values of the likelihood ($a$) without and ($b$) with the rational expectations restrictions imposed.

Using equation (7) with $\Phi_2 = 0$ to obtain an expression for $E_t[X_{t+1} - \Phi_0 D_{t+1}]$ and inserting it into equation (10) implies that the following restrictions must be satisfied

$$c_1'\Pi = -(c_1' + c_2' + c_3' + c_4'),$$
$$c_1'A_2 = -c_{-1}, \quad c_1'A_3 = -c_{-2}, \quad c_1'\Phi_1 = -\psi_0, \tag{11}$$

where $\Pi = A_1 + A_2 + A_3 - I$. For fixed values of the semi-structural parameters $\gamma_f$, $\gamma_b$, $\lambda$, and $\psi = \psi_1 = (1 - \psi_2)$, the concentrated likelihood, $L_{c1}(\gamma_f, \gamma_b, \lambda, \psi)$, can be computed using the methods in Johansen and Swensen (1999).

However, in this particular case, the restrictions have a form that makes further simplifications possible. As explained in Appendix B, we can use the methods of Johansen and Swensen (2008) to concentrate out the parameters $\gamma_f$, $\gamma_b$, and $\lambda$, so that the concentrated or profile likelihood, $L_{c2}(\psi) = L_{c1}(\gamma_f(\psi), \ \gamma_b(\psi), \lambda(\psi), \psi)$, only depends on $\psi = \psi_1 = (1 - \psi_2)$ in the case of homogeneity. Hence, it is possible to compute for each value of $\psi$ the value of the likelihood when the restrictions implied by equation (11) are satisfied. Thus, by computing the likelihood repeatedly for many values of $\psi$, it is possible to determine the maximal value of the likelihood.[3]

Using ordinary likelihood ratio tests, we can now test the four nested models: reduced-rank without homogeneity restriction; reduced-rank with homogeneity restriction; exact hybrid model; and exact non-hybrid model. The results are summarized in Table 2.

Using the usual top-down procedure, the test of the hybrid model is not rejected, whereas the non-hybrid model with $\gamma_b = 0$ is. Hence, it is a clear advantage to include the extra inflation lag in the expectation restrictions in the non-hybrid model. These impressions are also evident from the plots of the concentrated likelihoods shown in Figure 3.

[3] The procedure `optim` in the statistical package R – see http://www.r-project.org/ and R Development Core team (2006) – is used to carry out tests and estimation of the forward-looking models. The R codes used for the procedures in this section and the bootstrap procedure in the previous section are available at http://folk.uio.no/~swensen/isoe/isoe.html.

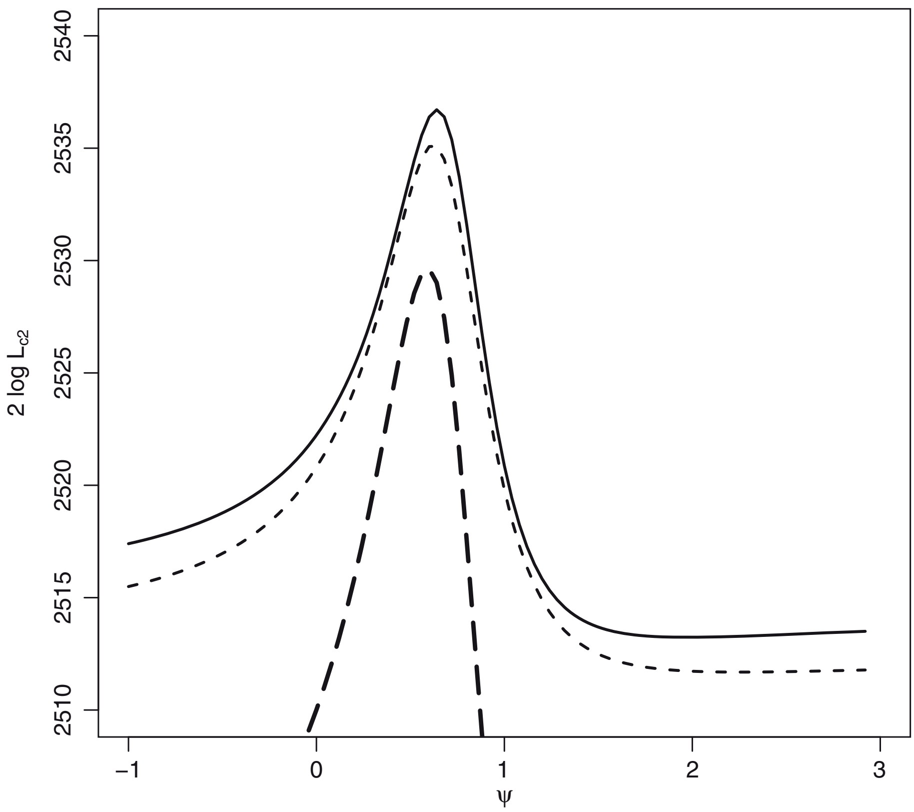


*Fig. 3.* Concentrated likelihood functions $2\log L_{c2}$ as functions of $\psi$ for the CVAR with homogeneity restriction (solid line), the exact hybrid model (short-dashed line) and the exact non-hybrid model (long-dashed line)

Table 3. *Some parameter estimates of the exact hybrid forward-looking model*

| $\psi$ | $\gamma_f$ | $\gamma_b$ | $\lambda$ |
|---|---|---|---|
| 0.80 | 3.56 | −0.74 | −0.14 |
| 0.62 | 5.16 | −0.75 | −0.27 |
| 0.40 | 5.84 | −0.87 | −0.23 |
| 0.10 | 4.83 | −0.92 | −0.11 |

*Notes:* The estimates of $\gamma_f$, $\gamma_b$, and $\lambda$ are computed for reasonable values of $\psi$.

The curve corresponding to the model imposing only the homogeneity restriction reaches a maximum at $\hat{\psi} = 0.641$. This is close to the maximum likelihood estimate of 0.621 in the hybrid model. To investigate this further, we have computed the restricted maximum likelihood estimates of the other semi-structural parameters $\gamma_f$, $\gamma_b$, and $\lambda$ for some reasonable fixed values of $\psi$, in addition to the maximum likelihood estimate $\hat{\psi} = 0.621$. The results are shown in Table 3. As can be seen, all sums of $\gamma_f$ and $\gamma_b$ are far

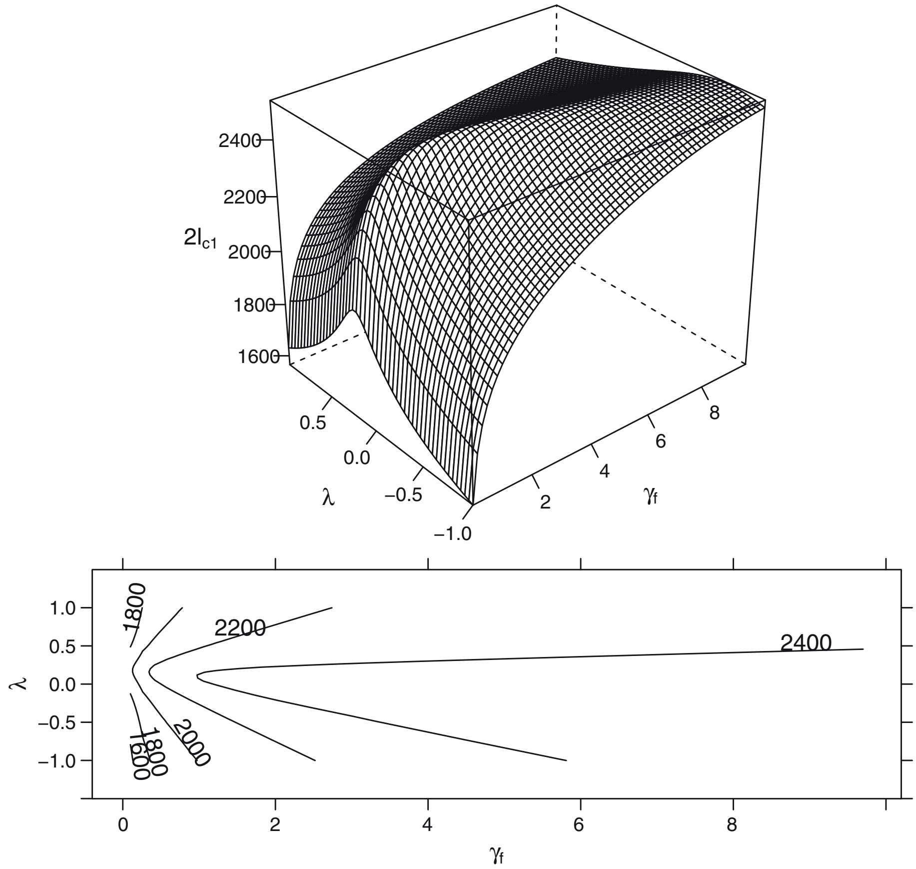


*Fig. 4.* Surface and contour plots of concentrated likelihood function $2l_{c1} = 2 \log L_{c1}$ as a function of $\gamma_f$ and $\lambda$ for the exact hybrid model with $\psi = 0.621$ and $\gamma_b = -0.745$; the maximal value is located at the point (5.16, −0.27)

from 1. This is not surprising. Fitting a hybrid model with the additional restriction $\gamma_f + \gamma_b = 1$ yields a maximal value of $2 \log L$ equal to 2489.10, corresponding to a value of 46.05 of $-2 \log LR$. The degree of freedom is 1, so the restriction is clearly rejected.

Additional evidence is provided by the plots in Figure 4 of the concentrated likelihood surface $2 \log L_{c1}(\gamma_f, \gamma_b, \lambda, \psi)$ as a function of $\gamma_f$ and $\lambda$ with $\psi$ and $\gamma_b$ fixed at 0.621 and −0.745, respectively, which are the maximum likelihood estimates of these parameters. A rather striking feature is the smooth ridge in the $\gamma_f$ direction, which indicates that this parameter is not well identified.

The economically meaningful values of the semi-structural parameters are $0 \leq \gamma_f, \gamma_b, \psi \leq 1$ and $\lambda \geq 0$. It is evident from Figure 4 that the unrestricted maximum likelihood estimates will be outside this region. By defining

$$\gamma_f(\theta_f) = \frac{\exp(\theta_f)}{1+\exp(\theta_f)}, \qquad \gamma_b(\theta_b) = \frac{exp(\theta_b)}{1+\exp(\theta_b)},$$

$$\psi(\theta) = \frac{\exp(\theta)}{1+\exp(\theta)}, \qquad \text{and} \qquad \lambda(\theta_l) = \exp(\theta_l),$$

and maximizing $L_{c1}(\gamma_f, \gamma_b, \lambda, \psi)$ with respect to $\theta_f$, $\theta_b$, $\theta$, and $\theta_l$, the restriction that the parameters have a meaningful economic interpretation can be imposed. The maximal value of $2\log L_{c1}$ is then equal to 2437.66, corresponding to the estimates $\hat{\gamma_f} = 1.0$, $\hat{\gamma_b} = 3.3E-06$, $\hat{\lambda} = 0.016$, and $\hat{\psi} = 0.70$, which are on the border of the permissible region. Therefore, the likelihood ratio test for the null hypothesis that they belong to this region has a non-standard asymptotic distribution, which is a convex combination of $\chi^2$ distributions with different degrees of freedom. In this case, the critical values are smaller than the critical values computed from a $\chi^2_4$ distribution. Because the difference between the maximal values of $2\log L$ is so large, $2535.15 - 2437.66 = 97.49$, and therefore exceeds all relevant critical values using a $\chi^2_4$ distribution, the likelihood ratio test also rejects the null hypothesis that $0 \leq \gamma_f, \gamma_b, \psi \leq 1$, and $\lambda \geq 0$.

Thus, we conclude that the hybrid model with no restrictions imposed is not rejected by likelihood ratio tests. However, when only economically meaningful parameters are allowed for, the model is not supported by the data. In addition, $\gamma_f$ seems to be poorly identified, as indicated by the rather smooth ridge of the likelihood surface plot.

### *The Inexact Version*

Another, perhaps more economically appealing, formulation of the restrictions (10) from the hybrid forward-looking model is

$$c_1' E_t[X_{t+1} - \Phi_0 D_{t+1}] + c_0' X_t + c_{-1}' X_{t-1} + c_{-2}' X_{t-2} + \psi_0 = u_t, \qquad (12)$$

where the error term $u_t$ is a sequence of innovations in the VAR model (i.e., $E_t[u_{t+1}] = 0$, and $c_1, c_0, c_{-1}$, and $c_{-2}$ are as defined earlier). The likelihood estimation of such a model must be handled using methods similar to those used by Boug *et al.* (2010) for bivariate systems (see also Kurmann, 2007; Fanelli, 2008).

The reduced-rank VAR model can be written on level form as

$$X_t = A_1 X_{t-1} + A_2 X_{t-2} + A_3 X_{t-3} + \Phi_0 D_t + \Phi_1 + \epsilon_t. \qquad (13)$$

Rewriting equation (12) at time $t+1$, using iterated expectations and inserting one-step ahead forecasts from the VAR, the restrictions on the coefficients implied by the hybrid model now take the form

$$c_1'(A_1^2 + A_2) + c_0'A_1 + c_{-1}' = 0 \tag{14}$$
$$c_1'(A_1A_2 + A_3) + c_0'A_2 + c_{-2}' = 0$$
$$c_1'(A_1A_3) + c_0'A_3 = 0$$
$$c_1'\Phi_1 + (c_1'A_1 + c_0')(\Phi_0 D_{t+1} + \Phi_1) + \psi_0 = 0. \tag{15}$$

The model (13) with reduced rank equal to unity and homogeneity imposed contains $18 + 3 + 1 = 22$ autoregressive parameters in addition to the coefficients of the deterministic terms. There are no more than nine restrictions on the coefficients of the VAR in equation (14), so using the reversed engineering approach of Kurmann (2007), expressing the likelihood in terms of the parameters of the inflation equation, and the semi-structural parameters $\gamma_f$, $\gamma_b$, and $\lambda$, we end up with at least 16 freely varying parameters in addition to those from equation (15). Finding the maximum likelihood estimates in such situations represents a computational problem that is beyond the scope of the present paper. Therefore, we explore some alternative approaches.

Equations (14) and (15) express necessary conditions for both the exact case, where $Var(u_t) = 0$, and the inexact case, where $Var(u_t) > 0$. By inspecting the equations, we can see that if equations (14) and (15) are fulfilled, the restriction that $c_1'A_1 + c_0' = 0$ is equivalent to the exact restrictions given in equation (11) – see also Swensen (2014). We can then use a minimum distance method along the lines of Sbordone (2002) and Magnusson and Mavroeidis (2010) in a related setting – see also Newey and McFadden (1994) for a general exposition – to derive estimators for $\gamma_f$ and $\lambda$ occurring in $c_1$ and $c_0$, and to test the restriction. The details are provided in Appendix C. The test statistic equals 0.11, and the degree of freedom is 1, so the corresponding $p$-value is 0.74. This might not be surprising as the likelihood ratio test for the exact hybrid model, without imposing economic admissible parameters, was not rejected either.

To investigate further the strong indication of an identification problem, which is evident in the exact version, we computed joint confidence regions of the parameters $\gamma_f$, $\lambda$, and $\gamma_b$ using an approach reminiscent of how the Anderson–Rubin statistic (Anderson and Rubin, 1949) is employed in similar situations. Consider the regression

$$\Delta p_t - \gamma_{f0}\Delta p_{t+1} + \lambda_0 \widehat{eqcm}_{1,t} - \gamma_{b0}\Delta p_{t-1} = \zeta_1' Z_{1,t} + \zeta_2' Z_{2,t} + error, \tag{16}$$

where the coefficients of the endogenous variables, $\gamma_{f0}$, $\lambda_0$, and $\gamma_{b0}$, have specified values, $Z_{1,t} = (1, SD_{1t}, SD_{2t}, SD_{3t})'$ are exogenous variables, and $Z_{2,t}$ are the instruments from the reduced-rank VAR model: $Z_{2,t} = (\Delta p_{t-2}, \Delta uic_{t-1}, \Delta uic_{t-2}, \Delta ulc_{t-1}, \Delta ulc_{t-2}, \widehat{eqcm}_{1,t-1})'$. The

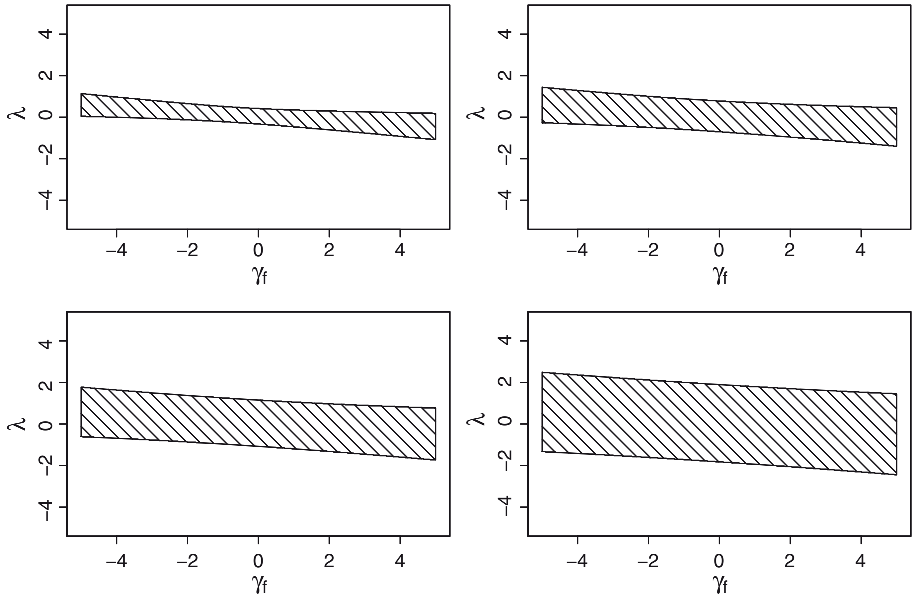


*Fig. 5.* Estimated confidence region for $\gamma_f$, $\lambda$, and $\gamma_b$ with confidence level 0.999; plots of $(\gamma_f, \lambda)$ where $\gamma_b$ is maximized over intervals $(-2, 2)$, $(-4, 4)$, $(-6, 6)$, and $(-10, 10)$

particular form of equation (16) arises by replacing the conditional expectation $E_t[\Delta p_{t+1}]$ in equation (5) with $\Delta p_{t+1}$, a common practice when estimating models of this form (see, e.g., Hansen *et al.*, 1996; Bårdsen *et al.*, 2005). The errors will have a first-order moving average structure. To take this dependency into account, we use the residuals from an ordinary least-squares (OLS) regression of equation (16) to estimate the correlation structure, and then we use a Wald statistic instead of the usual $F$-statistic. Figure 5 shows projected regions consisting of the parameters $(\gamma_f, \lambda, \gamma_b)'$ for which a test of $\zeta_2 = 0$ with the level 0.001 is not rejected. In the figure, the values $\gamma_b$ that are maximized above belong to the intervals $(-2, 2)$, $(-4, 4)$, $(-6, 6)$, and $(-10, 10)$.

The increasing size of the projected regions indicates that the three-dimensional confidence region for $(\gamma_f, \lambda, \gamma_b)'$ is unbounded, which signifies that there is an identification problem (e.g., Dufour, 2003). Thus, the fact that the exact and inexact models can share several features is confirmed by more than the non-rejection of the test.

## VI. A Competing Backward-Looking Model

So far, the formal tests clearly indicate that the forward-looking model is at odds with the data. By themselves, however, these tests are not sufficient evidence that inflation expectations do not matter. Expectations might still matter for inflation dynamics if past information is relevant through its implications for the expectations of future inflation. Remember that, so far, we have only considered expectations as conditional mathematical expectations. In this section, we compare and contrast estimates of a reduced form of equation (5) with a backward-looking Phillips curve as a competing model of inflation dynamics. To this end, we first estimate a backward-looking Phillips curve based on equation (6) with $\gamma_f = 0$ and the same information set used in the testing of the forward-looking models. Based on the estimated backward-looking Phillips curve, we infer one-step ahead inflation expectations, $E_t \Delta p_{t+1}$, substitute into equation (5), and solve for inflation, $\Delta p_t$, to obtain a reduced-form model. We then assess the fit of this reduced-form model by asking whether there are any economically reasonable values of the semi-structural parameters, $\gamma_f$, $\gamma_b$, and $\lambda$, which make the model consistent with the data in-sample. Finally, we evaluate the fit of the two models by means of a forecasting competition out-of-sample.

We rely on a general-to-specific modeling strategy in the estimation of the backward-looking model for $\Delta p_t$ using the autometrics procedure available in OxMetrics (see Doornik and Hendry, 2009). Our point of departure is a general conditional model for $\Delta p_t$ with $\Delta p_{t-1}$, $\Delta p_{t-2}$, $\Delta uic_t$, $\Delta uic_{t-1}$, $\Delta uic_{t-2}$, $\Delta ulc_t$, $\Delta ulc_{t-1}$, $\Delta ulc_{t-2}$, $\widehat{eqcm}_{2,t-1}$, 1, $SD_{1t}$, $SD_{2t}$, $SD_{3t}$, $D84Q1$, $D86Q3$, $D96Q1$, $D01Q3$, $D03Q1$, and $D03Q2$ as regressors. This general model is fully in accordance with the fitted reduced-rank VAR, both in terms of the number of lags and the weak exogeneity status of $uic_t$ and $ulc_t$. Autometrics picks the following specific model in our case, together with diagnostic tests and standard errors in parentheses:[4]

$$
\begin{aligned}
\Delta p_t = {} & \underset{(0.061)}{0.150}\Delta p_{t-1} + \underset{(0.064)}{0.132}\Delta p_{t-2} - \underset{(0.011)}{0.059}\widehat{eqcm}_{2,t-1} + \underset{(0.005)}{0.028} \\
& + \underset{(0.0011)}{0.0058} SD_{1t} + \underset{(0.0012)}{0.0036} SD_{2t} - \underset{(0.0011)}{0.0036} SD_{3t} + \underset{(0.003)}{0.020} D86Q3 \\
& - \underset{(0.003)}{0.011} D96Q1 - \underset{(0.003)}{0.015} D01Q3 + \underset{(0.002)}{0.023} D03Q1Q2 \\
& OLS,\ T = 93\ (1982Q4 - 2005Q4), \qquad \hat{\sigma} = 0.0033 \\
& AR_{1-5}\colon F(5,77) = 1.76\ [0.13], \qquad ARCH_{1-4}\colon F(4,85) = 1.72\ [0.15], \\
& NORM\colon \chi^2(2) = 1.21\ [0.55], \qquad HET\colon F(11,78) = 1.58\ [0.12].
\end{aligned} \tag{17}
$$

[4] $AR_{1-5}$ is a test for up to fifth-order residual autocorrelation; $ARCH_{1-4}$ is a test for up to fourth-order autoregressive conditional heteroskedasticity in the residuals; $NORM$ is a joint test for residual normality (no skewness and excess kurtosis); and $HET$ is a test for residual heteroskedasticity (see Doornik and Hendry, 2009). The numbers in square brackets are $p$-values. The dummy variable $D03Q1Q2$ combines the two dummy variables $D03Q1$ and $D03Q2$ and takes the value 1 in $2003Q1$ and $-1$ in $2003Q2$.

Several features of Norwegian inflation dynamics stand out from equation (17). First, the economic variables entering the model are all highly significant. Consumer price inflation in Norway seems to be rather persistent, as represented by the significant autoregressive coefficients of $\Delta p_{t-1}$ and $\Delta p_{t-2}$. The $\widehat{eqcm}_{2,t-1}$ appears with a $t$-value of $-5.36$, hence adding force to the results obtained from the cointegration analysis. Secondly, the sign of the impulse dummies corresponds well with the expected effects of the associated economic events described above. Thirdly, there are no significant contemporaneous short-run effects on inflation from unit import costs and unit labor costs in equation (17). No contemporaneous short-run effects combined with the small magnitude of the estimated loading coefficient ($-0.059$) imply very slow consumer price adjustment in the face of shocks in unit import costs and unit labor costs.

As stated by Fuhrer (2006) among others, lagged inflation is not simply a second-order add-on to the model, but it is important when accounting for persistence in inflation. It is not commonplace to require an explicit economic interpretation of parameters in a VAR model, except those necessary for identification, but we think this is possible in our case. The Norwegian CPI includes housing rents with a weight of around 0.17. Most of the contracts in the housing market include an index clause that allows the owner to adjust the rent based on the observed increase in the CPI. This usually takes place in January based on the CPI in December. During the rest of the year, rents are not adjusted in line with inflation. In many cases, rents are not adjusted every year, but remain nominally constant as long as the contract lasts. However, the total index for rents will increase during the year as old contracts expire and new contracts are signed. In addition, the standard contract has a clause that states that the rent can be renegotiated every third year to bring it into line with current market prices. The CPI for imputed rents for households living in their own house or flat is based on the rent equivalence principle, such that imputed rent follows observed rent with some sampling modifications. The relevance to modeling aggregate CPI in Norway is the acknowledgement that the housing market is quite different from standard product markets and that nominal price rigidity and long lags are present and observable in the microdata on which the CPI is based.

Empirical evidence of constancy of equation (17) can be assessed from recursive test statistics (see Doornik and Hendry, 2009). Neither one-step residuals with two estimated equation standard errors nor a sequence of break point Chow tests at the 1 percent significance level indicate non-constancy. All recursive estimates vary little, especially relative to their estimated uncertainty. The fact that no significant structural breaks are detected around the date of the shift in monetary policy regime from exchange rate targeting to inflation targeting (late March 2001) points to equation (17) not being subject to the Lucas critique.

To investigate whether incorporating forward-looking features into the model can lead to any improvement, we write the part of equation (17) not involving the impulse dummies as $\Delta p_t = a\Delta p_{t-1} + b\Delta p_{t-2} + c\widehat{eqcm}_{2,t-1} + d + seasonals$. We let that part be past information relevant to one-step ahead inflation expectations in equation (5) to obtain the following reduced-form forward-looking model for $\Delta p_t$:

$$\begin{aligned}\Delta p_t = \beta_0 + \beta_1\Delta p_{t-1} - \beta_2(\Delta ulc_t - \Delta uic_t)\\ -\beta_3(\Delta uic_t - \widehat{eqcm}_{2,t-1}) + seasonals,\end{aligned} \tag{18}$$

where

$$\begin{aligned}\beta_0 &= (\psi_0 + \gamma_f d)\varpi,\\ \beta_1 &= (\gamma_f b + \gamma_b)\varpi,\\ \beta_2 &= (\gamma_f c - \lambda)\psi_1\varpi,\\ \beta_3 &= (\gamma_f c - \lambda)\varpi,\\ \varpi &= (1 - a\gamma_f - c\gamma_f + \lambda)^{-1}.\end{aligned} \tag{19}$$

We see that equation (18) consists of the three composite parameters, $\beta_1$, $\beta_2$, and $\beta_3$, for determination of the three semi-structural parameters, $\gamma_f$, $\gamma_b$, and $\lambda$, which are thereby exactly identified. Because both $uic_t$ and $ulc_t$ are weakly exogenous, we can generate the variables $(\Delta ulc_t - \Delta uic_t)$ and $(\Delta uic_t - \widehat{eqcm}_{2,t-1})$, and we can estimate equation (18) by means of OLS. By also including the impulse dummies from equation (17) as additional regressors, we obtain the following OLS estimate of equation (18):

$$\begin{aligned}\Delta p_t = \underset{(0.067)}{0.126}\Delta p_{t-1} + \underset{(0.009)}{0.032}(\Delta ulc_t - \Delta uic_t) + \underset{(0.009)}{0.070}(\Delta uic_t - \widehat{eqcm}_{2,t-1})\\ + \underset{(0.004)}{0.035} + \underset{(0.0015)}{0.0029}SD_{1t} + \underset{(0.0015)}{0.0007}SD_{2t} - \underset{(0.0014)}{0.0061}SD_{3t}\\ + \underset{(0.004)}{0.017}D86Q3 - \underset{(0.004)}{0.010}D96Q1 - \underset{(0.004)}{0.013}D01Q3 + \underset{(0.003)}{0.022}D03Q1Q2\\ OLS, T = 93\ (1982Q4 - 2005Q4), \qquad \hat{\sigma} = 0.0035\\ AR_{1-5}\colon F(5,77) = 1.77\ [0.13], \qquad ARCH_{1-4}\colon F(4,85) = 1.19\ [0.32],\\ NORM\colon \chi^2(2) = 4.14\ [0.13], \qquad HET\colon F(11,78) = 1.68\ [0.09].\end{aligned} \tag{20}$$

Like the estimated backward-looking Phillips curve, the estimated reduced-form forward-looking model is well specified, judging by the diagnostic tests. However, $\hat{\sigma}$ increases somewhat, and calculations using equation (19) reveal that $\hat{\gamma}_f = 6.84$, $\hat{\gamma}_b = -0.908$, and $\hat{\lambda} = -0.408$, all of which are outside the regions containing sensible economic interpretations. These parameter estimates correspond to the earlier findings in Section V (Table 3) and might not be surprising, as both the hypotheses of weak exogeneity of $uic_t$ and $ulc_t$ and zero restrictions on the autoregressive parameters from the VAR are not rejected.

Although equation (20) has a slightly poorer fit than equation (17) and the parameter estimates are difficult to interpret economically, we compare the out-of-sample forecasting performance of the two competing models of inflation to shed light on their robustness with respect to relatively large movements in the exchange rate during the recent financial crisis. Taylor (2000) argues that the extent to which a firm matches exchange rate movements by changing its own price depends on how persistent the movements are expected to be. For a retail firm that adds services to its imports of goods, a depreciation of the exchange rate will raise the costs of the imports valued in domestic currency. According to Taylor (2000), if the depreciation is viewed as temporary, the retail firm will pass through less of the depreciation to its own price. In any case, if the price-setting behavior changed significantly following the financial crisis, we should expect instabilities in the estimated $\Delta p_t$ equations, as indicated, for example, by poor out-of-sample forecasting ability.

To assess the forecasting performance of equations (20) and (17), we employ 24 quarters (2006Q1–2011Q4) of out-of-sample observations, including the period of the financial crisis. According to the theoretical model in Section II, both the forward-looking and the backward-looking models embody the price level. This feature is supported by the data, as is evident from the cointegration analysis and from Figure 2. An inflation-targeting central bank is interested in forecasting inflation over a certain horizon (i.e., two to four years), but might not be very concerned about forecasts of quarterly inflation a few quarters ahead. Thus, to evaluate the forecasting performance of the two competing models of inflation, we focus on the medium-term (*ex ante*) dynamic forecasts for the consumer price level.

A preliminary investigation of the two models reveals that a majority of the actual values of $p_t$ stay within their corresponding confidence intervals over the forecasting period. However, the actual value of $p_t$ is close to being outside the confidence intervals in the first quarter of 2007. The actual value of $p_t$ in 2007Q1 and the values thereafter are likely to be influenced by the huge and transitory fall in electricity prices during 2007Q1. Consequently, the dynamic forecasts over-predict the actual values of $p_t$ thereafter, irrespective of whether equation (20) or equation (17) is used as the underlying forecasting model.

To take a closer look at this hypothesis, we re-estimated the two models over the period 1982Q4–2007Q1 with an impulse dummy in 2007Q1 as a separate regressor, controlling for the substantial fall in electricity prices

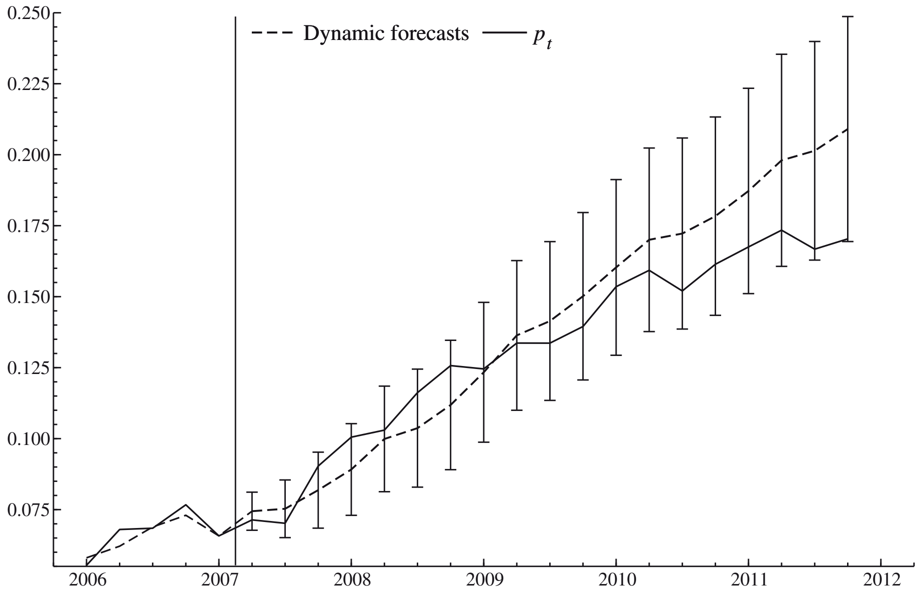


*Fig. 6.* Backward-looking model: actual values and dynamic forecasts of $p_t \pm$ 2STD

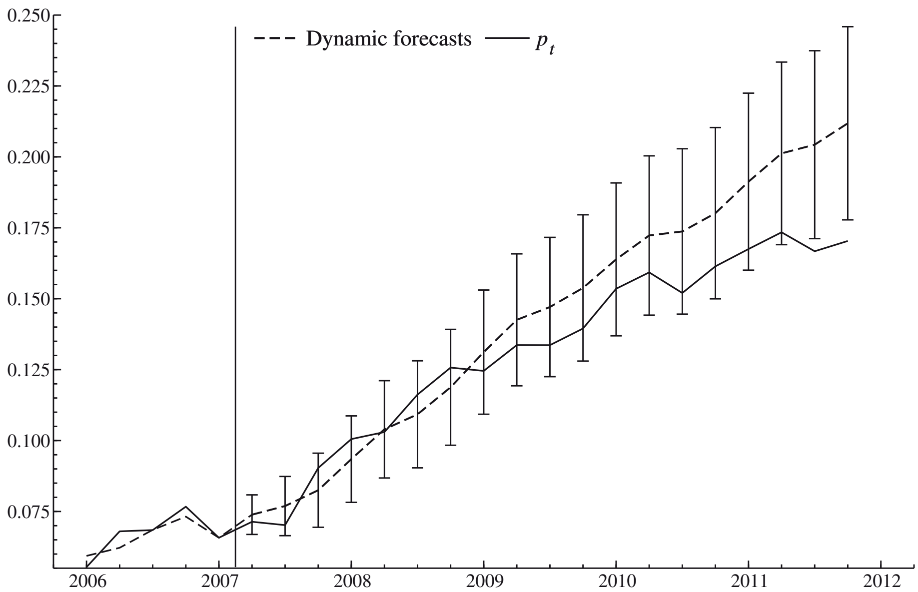


*Fig. 7.* Forward-looking model: actual values and dynamic forecasts of $p_t \pm$ 2STD

during that quarter. Hence, 19 observations are now available for forecasting. The two re-estimated models are virtually unchanged from equations (20) and (17) with respect to both parameter estimates and diagnostics. Figures 6 and 7 depict actual values of $p_t$ together with dynamic forecasts, adding bands of 95 percent confidence intervals to each forecast, when the re-estimated models are used for forecasting.

We observe that the out-of-sample forecasting ability of each of the two competing models of inflation is reasonably good, despite relatively large exchange rate movements in the wake of the financial crisis. That said, the forecasting performance of the backward-looking model is somewhat better than the forward-looking model because the RMSE is 0.0167 and 0.0183, respectively. Moreover, the average yearly forecasting failure within the relevant three-year horizon for the Norwegian central bank is about 0.23 percentage points with the former model and about 0.35 percentage points with the latter model. Because no significant forecasting failures are evident with the backward-looking model, we can argue, in light of Taylor (2000), that the exchange rate movements during the financial crisis were perceived as transitory rather than permanent shocks, such that firms found it rational not to alter their pricing behavior. We conclude that there might not be any value added of having forward-looking expectations once the model accounts for all relevant backward-looking terms, as is the case with equation (17).

Our results as regards forecasting can be compared to the results in King and Watson (2012). As mentioned in the introduction, they find that the Smets and Wouters (2007) forward-looking model of the US economy predicted inflation poorly after the financial crisis in 2008. Del Negro *et al.* (2015) argue that, when a standard dynamic stochastic general equilibrium (DSGE) model is extended by including financial frictions and the forecasts are conditional on the increased financial stress that occurred in autumn 2008, the forward-looking pricing equation forecasts inflation well after the financial crisis. Consequently, they accept that earlier inflation models did not perform well post-sample and that previous DSGE models did not include parameters that were stable over time. That said, their modified model has one interesting feature in that lagged inflation becomes more important than is standard in the literature. This supports the argument by Fuhrer (2006) that lagged inflation is not a minor feature of the forward-looking model. Rather, it *is* the model.

## VII. Conclusions

In this paper, using Norwegian data, we have evaluated the empirical performance of forward-looking models for inflation dynamics in a small open economy. Our forward-looking models relate current inflation to expected future inflation and the difference between the actual price and the steady-state value of levels as a theory-consistent forcing variable. The steady-state value is specified as a mark-up over marginal costs, which, in turn, are determined by costs of both labor and imported intermediate goods in line with the open economy literature. The models thus include variables, both levels and differences, that require caution with respect to both the time series properties and possible cointegrated nature of the variables involved. Such econometric issues have typically been ignored in related studies on data from open economies, and they raise questions about the validity of the existing statistical inference.

First, using reduced-rank regressions, we established a cointegrating vector between consumer prices and marginal costs. Thus, the theory-driven price equation underlying the models of inflation dynamics is well supported by the data. Secondly, using likelihood-based testing procedures, we found that the exact version of the forward-looking model is at odds with the data. That is, the rational expectations hypothesis is not rejected statistically, but when only economically meaningful parameters are allowed for, the model is not supported by the data. In addition, some of the parameters in the exact version of the model are not well identified. We also found that the inexact version is not rejected statistically using a test based on a minimum distance method. However, confidence regions obtained by inverting a test reminiscent of the Anderson–Rubin statistic reveal an identification problem in this model as well.

Finally, we established a well-specified dynamic backward-looking model, which, in addition to the theory-consistent forcing variable, includes backward-looking terms only. The backward-looking model of inflation dynamics outperforms a reduced-form forward-looking model, it is reasonably stable in a sample containing a major monetary policy regime shift from exchange rate targeting to inflation targeting, and it forecasts well post-sample, during and after the recent financial crisis. All these findings provide strong evidence in favor of the backward-looking model, and the Lucas critique does not seem to be important in our empirical context. Our findings are in line with Fuhrer (2006), among others, who points out that lagged inflation is not a second-order add-on to the optimizing model, it *is* the model.

## Appendix A

We show here how the steady-state value of the price level in equation (1) is derived. As mentioned in the text, a representative firm is assumed to face a Cobb–Douglas production function of the form

$$Y_t = A L_t^{\alpha} I_t^{\beta}, \tag{A1}$$

where $L_t$ and $I_t$ denote labor and imports of intermediate inputs, respectively. Variable costs in production are given by the sum of labor costs and import costs

$$C_t = W_t L_t + P I_t I_t, \tag{A2}$$

where $W_t$ is wage per hour and $PI_t$ is the price of imports. Minimizing variable costs given the production function leads to

$$W_t L_t = (\alpha/\beta) PI_t I_t. \tag{A3}$$

Solving for $I_t$ in equation (A3) and inserting it into equation (A1) gives

$$L_t = L_0 (W_t/PI_t)^{-\beta/(\alpha+\beta)} Y_t^{1/(\alpha+\beta)}. \tag{A4}$$

Inserting equation (A4) into equation (A3) and solving for $I_t$ gives a similar expression for $I_t$:

$$I_t = I_0 (W_t/PI_t)^{\alpha/(\alpha+\beta)} Y_t^{1/(\alpha+\beta)}. \tag{A5}$$

Inserting equations (A4) and (A5) into equation (A2) gives the cost function, and it is straightforward to show that marginal costs $MC_t$ become

$$MC_t = C_0 W_t^{\alpha/(\alpha+\beta)} PI_t^{\beta/(\alpha+\beta)} Y_t^{1/(\alpha+\beta)-1}, \tag{A6}$$

which is homogeneous of degree one in factor prices. Using the production function (A1), we can express marginal costs as a function of two terms

$$MC_t = C_0 (W_t L_t / Y_t)^{\alpha/(\alpha+\beta)} (PI_t I_t / Y_t)^{\beta/(\alpha+\beta)}. \tag{A7}$$

Using lowercase letters to indicate logs, we have

$$mc_t = c_0 + \psi_1 ulc_t + \psi_2 uic_t, \tag{A8}$$

where $ulc_t$ is labor unit costs, $uic_t$ is import unit costs, $\psi_1 = \alpha/(\alpha+\beta)$, and $\psi_2 = \beta/(\alpha+\beta)$. As explained in the text, a representative profit-maximizing firm facing an isoelastic demand function will set prices as a mark-up on marginal costs. Thus, in log form, we have $p_t^* = m_0 + mc_t$, where $m_0$ is the constant mark-up. Using the expression for marginal costs in equation (A8), we have

$$p_t^* = m_0 + \psi_1 ulc_t + \psi_2 uic_t, \tag{A9}$$

which is the steady-state value of the price level given in equation (1).

## Appendix B

Equation (10), which describes the estimated hybrid forward-looking model, can be reformulated as

$$\begin{aligned} c_1' E_t[\Delta X_{t+1} - \Phi_0 D_{t+1}] + (c_1' + c_0' + c_{-1}' + c_{-2}') X_t \\ -(c_{-1}' + c_2') \Delta X_t - c_2' \Delta X_{t-1} + \psi_0 = 0. \end{aligned} \tag{B1}$$

If $e_1 = (1, 0, 0)'$ denotes the vector having the first element equal to 1 and zero otherwise, $c_1 = \gamma_f e_1$, $c_{-1} + c_{-2} = e_1$ and $c_{-2} = -\gamma_b e_1$. Furthermore, if we define $d = -(c_1' + c_0' + c_{-1}' + c_{-2}')/\lambda = (1, -\psi_1, -(1 - \psi_1))$, the model (B1) can be written as

$$\begin{aligned}
e_1' E_t[\Delta X_{t+1} - \Phi_0 D_{t+1}] &= \left(\frac{\lambda}{\gamma_f}\right) d' X_{t-1} + \left(\frac{1}{\gamma_f}\right) e_1' \Delta X_t \\
&\quad - \left(\frac{\gamma_b}{\gamma_f}\right) e_1' \Delta X_{t-1} - \psi_0 \\
&= \tau d' X_{t-1} + \tau_1 e_1' \Delta X_t + \tau_2 e_1' \Delta X_{t-1} + \mu_0 \\
&= 0,
\end{aligned}$$

where the parameters $\tau = \lambda/\gamma_f$, $\tau_1 = 1/\gamma_f$, $\tau_2 = -\gamma_b/\gamma_f$, and $\mu_0 = -\psi_0/\gamma_f$ vary freely. Thus, for fixed $\psi = \psi_1$, this has exactly the form treated in Johansen and Swensen (2008).

The marginal part of the model now takes the form

$$\begin{aligned}
\Delta p_t = \tau(p_{t-1} - \psi ulc_{t-1} - (1 - \psi) uic_{t-1}) + \tau_1 \Delta p_{t-1} + \tau_2 \Delta p_{t-2} \\
+ \phi' D_t + \epsilon_{1t},
\end{aligned}$$

where $\phi = (\phi_1, \ldots, \phi_{10})$ are the coefficients of the seasonal dummies, the other six dummies, and the constant. Note that there are four restrictions involved, as the coefficients of $\Delta ulc_{t-1}$, $\Delta uic_{t-1}$, $\Delta ulc_{t-2}$, and $\Delta uic_{t-2}$ are constrained to zero. The conditional part involves an unrestricted regression of $(\Delta ulc_t, \Delta uic_t)'$ on $\Delta p_t$, $(p_{t-1} - \psi ulc_{t-1} - (1 - \psi) uic_{t-1})$, $\Delta p_{t-1}$, $\Delta ulc_{t-1}$, $\Delta uic_{t-1}$, $\Delta p_{t-2}$, $\Delta ulc_{t-2}$, $\Delta uic_{t-2}$, 1, and the seasonal dummies and the impulse dummies.

Therefore, for fixed values of $\psi$, the numerical value of the likelihood can be evaluated. Hence, the maximum likelihood estimate of $\psi$ can be determined. Once this is given, the estimates for $\tau$, $\tau_1$, and $\tau_2$ can be found by OLS from the marginal part with the estimate of $\psi$, and the semi-structural parameters $\gamma_f$, $\gamma_b$, and $\lambda$ computed.

## Appendix C

In the following, we present some details about the minimum distance approach used in the test of the restrictions $c_1' A_1 + c_0' = 0$. It is convenient to rewrite the restrictions as $B(\nu)\theta + h = 0$, where $B(\nu)$ is a $3 \times 2$ matrix depending on constants and the parameters $\nu$ of the autoregressive model, $\theta = (\gamma_f, \lambda)'$, and $h$ is a vector where the elements are known constants.

In the alternative equilibrium correction parametrization, the VAR model can be written as

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \Phi_0 D_t + \Phi_1 + \epsilon_t,$$

where $\Pi = A_1 + A_2 + A_3 - I = \alpha\beta'$, $\Gamma_1 = -A_2 - A_3$, and $\Gamma_2 = -A_3$. Thus, $A_1 = \Pi + I - A_1 - A_2 = \Pi + I + \Gamma_1$. In particular, due to the super-consistency, the parameters in the cointegration vector (i.e., $\beta = (1, -\psi_1, -\psi_2)'$), where $\psi_1 + \psi_2 = 1$, can be considered as known. In this case, $\nu$ corresponds to $\alpha = \{\alpha_i\}_{i=1}^3$, $\Gamma_1 = \{\gamma_{1,ij}\}_{i,j=1}^3$, $\Gamma_2 = \{\gamma_{2,ij}\}_{i,j=1}^3$, and

$$B(\nu)\theta + h = \begin{pmatrix} \alpha_1 + \gamma_{1,11} & -1 \\ -\alpha_1\psi_1 + \gamma_{1,12} & \psi_1 \\ -\alpha_2\psi_2 + \gamma_{1,13} & \psi_2 \end{pmatrix} \begin{pmatrix} \gamma_f \\ \lambda \end{pmatrix} + \begin{pmatrix} -1 \\ 0 \\ 0 \end{pmatrix}.$$

If $\hat{\nu}$ is the usual estimator for $\nu$, as in Johansen (1995), then the minimum distance estimator for $\theta(\nu)$ can be expressed as $\hat{\theta}(\hat{\nu}) = -[B(\hat{\nu})'B(\hat{\nu})]^{-1}B(\hat{\nu})'h$.

From theorem 13.5 in Johansen (1995), it follows that $\sqrt{T}(\hat{\nu} - \nu)$ is approximately multivariate Gaussian. Under the restriction that $B(\nu)\theta + h = 0$, we can write

$$\begin{aligned}\sqrt{T}(B(\hat{\nu})\hat{\theta}(\hat{\nu}) + h) &= \sqrt{T}(B(\hat{\nu})\hat{\theta}(\hat{\nu}) + h - B(\nu)\theta - h) \\ &= \sqrt{T}[(B(\hat{\nu}) - B(\nu))\hat{\theta}(\hat{\nu}) + B(\nu)(\hat{\theta}(\hat{\nu}) - \theta)].\end{aligned}$$

From the $\delta$-method, it follows that $\sqrt{T}(\hat{\theta}(\hat{\nu}) - \theta)$ has the same asymptotic distribution as $\sqrt{T}(\partial\hat{\theta}/\partial\nu)|_{\nu=\hat{\nu}}(\hat{\nu} - \nu)$. Thus, both terms in $\sqrt{T}(B(\hat{\nu})\hat{\theta}(\hat{\nu}) + h)$ can be expressed by $\sqrt{T}(\hat{\nu} - \nu)$, and the sum is therefore also asymptotically Gaussian. If the covariance matrix is estimated by plugging in the estimates for the unknown parameters, a test statistic for the null hypothesis $c_1'A_1 + c_0' = 0$ can be found. The asymptotic distribution is $\chi^2$ with one degree of freedom.

## Appendix D

$P$ denotes the official consumer price index (2002 = 1). (Source: Statistics Norway.)

*UIC* denotes unit import costs proxied by the implicit deflator of total imports (2002 = 1). (Source: Statistics Norway, the Quarterly National Accounts.)

*ULC* denotes unit labor costs proxied by $YWP/QP$, where $YWP$ and $QP$ are total labor costs and value added in the private mainland economy, respectively. (Source: Statistics Norway, the Quarterly National Accounts.)

$D84Q1$ is an impulse dummy used to account for a large residual in the $ulc_t$ equation of the VAR. Equals unity in the first quarter of 1984, and zero otherwise.

$D86Q3$ is an impulse dummy used to control for the 12 percent devaluation of the Norwegian currency in May 1986. Equals unity in the third quarter of 1986, and zero otherwise.

$D96Q1$ is an impulse dummy used to control for the reduction in indirect tax rates during the first quarter of 1996. Equals unity in the first quarter of 1996, and zero otherwise.

$D01Q3$ is an impulse dummy used to control for the drop in the VAT rate on food from 24 percent to 12 percent in July 2001. Equals unity in the third quarter of 2001, and zero otherwise.

$D03Q1$ is an impulse dummy used to control for the large increase in electricity prices during the first quarter of 2003. Equals unity in the first quarter of 2003, and zero otherwise.

$D03Q2$ is an impulse dummy used to control for the large decrease in electricity prices during the second quarter of 2003. Equals unity in the second quarter of 2003, and zero otherwise.

$D07Q1$ is an impulse dummy used to control for the large decrease in electricity prices during the first quarter of 2007. Equals unity in the first quarter of 2007, and zero otherwise.

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# 4 Chapter 4

## Expectations and regime robustness in price formation: evidence from vector autoregressive models and recursive methods

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ORIGINAL PAPER

# Expectations and regime robustness in price formation: evidence from vector autoregressive models and recursive methods

**Pål Boug · Ådne Cappelen · Anders Rygh Swensen**

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**Abstract** The forward-looking linear quadratic adjustment cost (LQAC) model has received attention when modelling prices. Empirical evidence supporting the model seems, however, ambiguous. We find that the LQAC-model is severely at odds with price data for Norwegian machinery exports also when the pure forward-looking rule is augmented by additional lags of the targeted variable. A conditional equilibrium correction (EqCM) model explains the export price behaviour more accurately. Our findings may rule out a large class of expectations based models and not just the particular LQAC-model in the formation of export prices. We also demonstrate that the EqCM-model performs well post-sample despite that monetary policy in Norway has changed from a fixed to a floating exchange rate regime following a recent introduction of inflation targeting. This regime robustness shows that the Lucas critique lacks force empirically in our case.

**Keywords** Expectations · Export prices · LQAC-model · VAR model · EqCM-model · Lucas critique

**JEL classification** C51 · C52 · D84 · E31

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## 1 Introduction

While the usefulness of cointegration analysis for studying long run economic relationships is hardly controversial, the modelling of short run behaviour is more of an unsettled issue. Data based methods are often used to develop conditional equilibrium correction models (EqCM), but theory based methods are also widely applied in the literature. One theory-based model used to explain short run dynamics is the linear quadratic adjustment cost (LQAC) model under rational expectations, cf. Sargent (1978). Proponents of this approach often refer to the Lucas critique[1] as motivation for discarding conditional models and argue that LQAC-models may circumvent the critique if economic agents indeed are forward-looking. Estimation of such a model is however not compelling evidence in itself for or against the Lucas critique, even if the proposed LQAC-model is not rejected. One should also demonstrate the weaknesses of competing models and verify that those models suffer from the Lucas critique. Conversely, rejection of a particular LQAC-model does not necessarily preclude economic agents from acting on alternative expectation based models.

The formation of export prices is an area in which the LQAC-model under rational expectations has been studied by e.g. Cuthbertson (1986, 1990) and Price (1991, 1992). Cuthbertson (1990) and Price (1991, 1992) conclude that the empirical performance of the model using UK manufacturing export price data compares favourably with an EqCM-model. However, not much evidence is given in these studies about the constancy or otherwise of the conditional and/or marginal models. One may then in accordance with Hendry (1988), see also Ericsson and Hendry (1999), argue that the LQAC-models and the EqCM-models reported in these studies are hard to distinguish empirically. The present paper evaluates the empirical performance of both the LQAC-model and the EqCM-model using Norwegian machinery export price data. The Norwegian central bank followed a policy of exchange rate targeting in various forms for many years. Recently, monetary policy in Norway changed fundamentally to inflation targeting and hence floating exchange rates, which could in principle have caused the export price formation to *shift* in accordance with the Lucas critique.

Our point of departure is the standard assumption of imperfectly competitive markets, where the optimal export price is determined as a mark-up over marginal costs. We estimate marginal costs using a fully specified cost-function consistent with economic theory. The mark-up in turn is modelled as depending on relative prices between Norwegian and competing prices. To evaluate the LQAC-model we use the testing procedure suggested by Johansen and Swensen (1999). The idea behind the procedure is to formulate the restrictions implied by the forward-looking rule (i.e., the Euler equation) from the dynamic optimisation of the LQAC-model within a cointegrated vector autoregressive (VAR)

[1] Lucas (1976) argues that the parameters of econometric models depend crucially on agents' expectations and are unlikely to remain stable in the event of policy regime changes. See Favero and Hendry (1992) for a comprehensive review and interpretation of the Lucas critique.

model. The results show that there is overwhelming empirical evidence against the LQAC-model. We then follow Hendry (1988) and show by means of recursive methods that there exists a stable EqCM-model of export price behaviour (despite regime changes) along with an unstable marginal process for one of the conditioning variables in that model. Finally, we demonstrate that the estimated EqCM-model performs well post-sample in spite of the known regime change in monetary policy following the introduction of inflation targeting. Hence, the invariance property to monetary policy regimes shows that the EqCM-model is not subject to the Lucas critique.

The remainder of the paper is organised as follows: Sect. 2 outlines the economic background, while Sect. 3 describes the data used in the analyses. Sections 4, 5 and 6 presents the empirical findings from tests of cointegration, tests of the LQAC-model and the EqCM-model along with tests of the Lucas critique, respectively. Section 7 concludes.

## 2 The economic background

We assume that products aimed for exports are differentiated from similar goods produced in other countries. Producers are consequently assumed to face regular downward sloping demand curves. Profit maximisation then leads to the standard formula stating that the export price (PA) equals a mark-up (MU) times marginal costs (MC)

$$\mathrm{PA} = \mathrm{MU} \cdot \mathrm{MC}. \tag{1}$$

Our measure of marginal costs is based on a fully specified cost function in line with neoclassical producer behaviour assuming a Cobb–Douglas production function that relates gross output to capital being fixed and labour, materials and energy being variable factors.[2] Minimising costs with respect to variable inputs allows us to derive a *dual* cost function for variable factor costs as a function of gross output ($Y$), the capital stock ($K$), the input prices of labour ($W$), materials (PM) and energy (PE) and a time trend ($t$) reflecting total factor productivity. Marginal costs are then found by differentiating this function with respect to gross output. Thus, we specify MC as

$$\mathrm{MC} = f(Y, K, W, PE, PM, t, \lambda), \tag{2}$$

[2] The reason why we specify a Cobb–Douglas production function in this way is that our price variable is an actual price index (albeit an aggregation of a number of SITC-unit prices) not a value added deflator for a sector or the economy as a whole like Gali et al. (2001). The prices that agents in the economy determine are on gross output not value added. Thus the model counterpart to that is marginal costs based on total inputs not only capital and labour.

where $\lambda$ is a vector containing the parameters of $f(\cdot)$. We attain a (data based) marginal cost variable by utilising estimates of $\lambda$.[3] The rationale of using the cost function (2) is that MC is theoretically more consistent than the standard measure.

The mark-up in (1) is often assumed to be a constant by referring to one particular case in Dixit and Stiglitz (1977). If the commodities we study are good substitutes among themselves, but poor substitutes for other goods, and a number of other assumptions are invoked, so-called two-stage budgeting is valid. Moreover, if the number of goods in the industry is large (denoted by $n$) so that $1/n$ is small, Dixit and Stiglitz (1977) show that the individual price has little impact on the aggregate industry price. Hence, we may assume that the individual producer ignores the effect of his price setting on the aggregated industry price. In a less restrictive case, the mark-up is not constant anymore but will depend on all factors affecting demand for the particular commodity, cf. Eq. (32) in Dixit and Stiglitz (1977). We allow the mark-up to depend on relative prices in a way that can accommodate the view that for small open economies (or sectors in an economy) producers can be price takers on world markets. If we denote the competing price that exporters face for PF, we assume

$$\mathrm{MU} = m\left(\frac{\mathrm{PA}}{\mathrm{PF}}\right). \tag{3}$$

This specification of the mark-up allows for a general model to be tested, with a constant mark-up as a special case. From (1) and (3) we then obtain

$$\mathrm{PA} = g(\mathrm{MC}, \mathrm{PF}). \tag{4}$$

The function $g(\cdot)$ is homogenous of degree one in MC and PF. Although the price equation is derived from a theory of monopolistic competition, it also contains the main alternative as a special case, namely that of "the law of one price" or perfect competition for homogenous goods. In the latter case the export price is equal to the price of the competitors, so that PA = PF.[4]

We may interpret (4) as a *static* model describing a price target for exporters. Hence, it will serve as the starting point for the cointegration analysis in Sect. 4.

[3] Applying Shephard's lemma to the *dual* cost function related to variable production costs, we may show that the optimal levels of variable factors of production under cost-minimisation are functions of the same parameters of $\lambda$ as in Eq. (2). Thus, the estimates of $\lambda$ are deduced from EqCM-models of variable factors of production using quarterly data. For reasons of space we do not report all the estimation results of the EqCM-models here. The estimates of $\lambda$ are as follows: $\alpha = 1, \alpha_K = 0.39, \alpha_L = 0.34, \alpha_M = 0.64, \alpha_E = 0.02$ and $\rho = 0.33$. Noticeably, $\alpha$ and $\alpha_K$ are the scale elasticities with respect to variable factors of production and the capital stock, respectively, while $\alpha_L, \alpha_M$ and $\alpha_E$ represent factor input elasticities associated with labour, materials and energy inputs. The estimate $\rho$ represents the semi-elasticity (quarterly) with respect to total factor productivity. All estimates of $\lambda$ make economic sense and are significant at conventional levels.

[4] To see this, let $\mathrm{MU} = m_0(\mathrm{PA}/\mathrm{PF})^{m_1}$. Then, using (1), we have $\mathrm{PA} = m_0^{1/(1-m_1)} \cdot \mathrm{PF}^{-m_1/(1-m_1)} \cdot \mathrm{MC}^{1/(1-m_1)}$ and PA = PF if $m_1$ approaches infinity. Strictly speaking, we can only identify proportionality as PA and PF are indices and not actual prices. See Sect. 3 for details.

It is however likely that deviations from the target in response to shocks are rational in an environment with adjustment costs. In a set-up involving rational expectations, the forward-looking LQAC-model generally attempts to reconcile costs incurred by not hitting the target on the one hand and costs resulting from changing the prices on the other hand. We assume that the export prices are chosen in order to minimise the intertemporal quadratic loss function

$$E_t\left[\sum_{j=0}^{\infty}\beta^j\left[\theta\left(\mathrm{pa}_{t+j}-\mathrm{pa}^*_{t+j}\right)^2+\left(\mathrm{pa}_{t+j}-\mathrm{pa}_{t+j-1}\right)^2\right]\right], \tag{5}$$

where $E_t$ denotes the conditional expectation given the information contained in the information set at time $t$, $\mathrm{pa}^*_t$ represents the target and lower case letters indicate logs, i.e., $\mathrm{pa}_{t+j}=\log(\mathrm{PA}_{t+j})$. The parameter $\beta$ represents the discount rate, and thus must satisfy the restriction that $0<\beta<1$. The parameter $\theta$ is the relative weight of the deviation from the optimal price (first term in (5)) compared to the cost of changing the price (second term in (5)). Exporters are assumed to determine a sequence of export prices so as to minimise the expected present discounted value of all future squared deviations from the target. However, since changes in the export price will be penalised as well, immediate adjustment towards the target will be non-optimal unless $\theta$ is large.

As is well known, see e.g. Sargent (1978), Nickell (1985) and Engsted and Haldrup (1994, 1997), the first order condition for this minimisation problem is an Euler equation of the form

$$\Delta\mathrm{pa}_t=\beta E_t[\Delta\mathrm{pa}_{t+1}]-\theta(\mathrm{pa}_t-\mathrm{pa}^*_t). \tag{6}$$

Our price equation has some similarities with recent advances in macroeconomic modelling of price setting behaviour, see e.g. Roberts (1995), Gali et al. (2001), Bårdsen et al. (2004) and Henry and Pagan (2004). Roberts (1995) shows that several New Keynesian models with rational expectations have a common representation, including the LQAC-model and the models of staggered contracts developed by Taylor (1979, 1980) and Calvo (1983). Gali et al. (2001) obtain the following new New Keynesian Phillips curve (cf. their equation (10)): $\pi_t=\beta E_t(\pi_{t+1})+\lambda\mathrm{mc}_t$, where $\pi$ is inflation and mc is the log deviation of marginal costs from its steady state value. The inflation terms in our equation (6) are identical to those in Gali et al. (2001). Their marginal cost variable is based on a Cobb–Douglas production function although only with labour as variable input as they use highly aggregated data. They assume isoelastic demand, which implies that the mark-up is a constant while we test (and reject!) that assumption. Using the same assumptions as Gali et al. (2001) our term $(\mathrm{pa}_t-\mathrm{pa}^*_t)$ may be expressed (disregarding a constant) as $(\mathrm{mc}_t-\mathrm{mc}^*_t)$ where the starred variable is the steady state value of marginal costs. Thus, our Eq. (6) corresponds to Eq. (10) in Gali et al. (2001) by assuming isoelastic demand and by substituting out the level (equilibrium correction) term.

The target $\mathrm{pa}_t^*$ in (6) is assumed to be generated by

$$\mathrm{pa}_t^* = \gamma_1 \mathrm{pf}_t + \gamma_2 \mathrm{mc}_t \tag{7}$$

to facilitate comparisons with related export price studies. We note that (7) is a logarithmic transformation of (4), such that $\gamma_1$ and $\gamma_2$ may be interpreted as the partial effect of the competing price and the marginal costs, respectively, on the export price. The homogeneity restriction implies that $\gamma_1 + \gamma_2 = 1$. To simplify matters, we follow the common practice of *exact* linear rational expectations models in the spirit of Hansen and Sargent (1991) and do not add a disturbance term on the right hand side of (7) when evaluating the empirical performance of the LQAC-model.

## 3 The data[5]

The empirical analyses are conducted using quarterly, seasonally unadjusted data on machinery and equipment export prices (PA) over the period 1978:1–1998:4. The actual estimation periods are shorter due to loss of observations as a result of lags and differences. We extend the sample, however, by twenty four quarters in order to conduct *out-of-sample* forecasting over the period 1999:1–2004:4. The chosen aggregate of products accounts for approximately 25% of total Norwegian manufacturing exports. Our aggregate of relatively homogenous products is motivated so as to reduce the risk of instability in estimated parameters caused by aggregation of heterogeneous products. Thus, we may claim that any instability findings in this paper regarding estimated export price equations are less likely to be due to the choice of the dependent variable. Another reason why we have chosen to study sector prices and not aggregate prices such as the consumer price index, which is typically the case in most recent studies of the New Keynesian Phillips curve, is that two-stage budgeting as in the original Dixit and Stiglitz (1977) paper is clearly more likely to be a good approximation at the sector level than at the macro level.

We employ the Norwegian import price (PF) of machinery and equipment as a proxy for the world market price of machinery and equipment that Norwegian exporters face. The advantage of using this proxy is that the definition and aggregation of PF is identical to that of PA. Also, given the very small size of the Norwegian economy, it seems reasonable to assume that PF is not much influenced by Norwegian markets, but reflects world market prices. The data for PA and PF are indices with 1996 as base year and are expressed in Norwegian currency. As explained in Sect. 2, a data based marginal cost variable (MC) serves as a proxy for production costs. Since the variable MC is a function of $\lambda$, the question arises as to whether the subsequent analyses suffer from the "generated regressor" problem, cf. Pagan (1984). To shed light on the robustness

[5] The data source is quarterly national accounts from Statistics Norway.

of the results of the LQAC-model, we thus also use a more traditional way of measuring marginal costs, namely that of variable unit costs (VC).

Figure 1 displays the export price, the competing price and the marginal costs, together with the two ratios $(\mathrm{pa}-\mathrm{pf})_t$ and $(\mathrm{pa}-\mathrm{mc})_t$ over the sample period. We observe that the two price series as well as the marginal costs exhibit a clear upward trend, suggesting $\mathrm{pa}_t, \mathrm{pf}_t$ and $\mathrm{mc}_t$ to be I(1). Therefore, a reduced rank VAR is a candidate as an empirical model. An alternative may be a stationary VAR around a breaking linear trend (with a break point around 1990). We also notice that Norwegian export prices have increased somewhat relative to both competing prices and marginal costs over the entire period. The movements in the ratios over the business cycle offer a somewhat more complex picture. Overall, the data set seems to suggest that both competing prices and marginal costs are important candidates for explaining the Norwegian export price of machinery. It is not obvious, however, from Fig. 1 whether the data series have the property of forming a cointegrating vector(s) that is consistent with the economic model.

## 4 Cointegration analysis

In this section, we utilise the procedure suggested by Johansen (1991) to fit a reduced rank VAR to the data. The cointegration analysis commenced from a fifth order VAR in $X_t = (\mathrm{pa}_t, \mathrm{pf}_t, \mathrm{mc}_t)'$ augmented with an unrestricted intercept and three centred seasonal dummies (labelled $\mathrm{CS}_{1t}, \mathrm{CS}_{2t}$ and $\mathrm{CS}_{3t}$).[6] Residual misspecification tests show that a fourth order VAR produces residuals with statistically acceptable properties. Also, recursive methods indicate that the fourth order VAR is constant over the sample, cf. Boug et al. (2005).[7]

Table 1 contains results from applying the procedure suggested by Johansen (1991) to determine the rank of the VAR. The *P*-values of the maximum eigenvalue ($\lambda_{\max}$) and trace statistics ($\lambda_{\text{trace}}$) reject the null of no cointegration at the 10 and 5% levels, respectively, but the null of at most one cointegration vector

[6] We follow existing studies and simplify matters by assuming that the export market can be distinguished from the home market. This assumption may be doubtful if Norwegian manufacturers set the export price and the domestic price simultaneously. We investigated this possibility on our data set expanded with data on the domestic price of machinery. Using Johansen's procedure, we found some evidence of two significant cointegrating vectors among the four variables. A *priori*, one of the vectors should most naturally be interpreted as an export price equation, whereas the other vector should be interpreted as a domestic price schedule. The results from testing economically meaningful identifying restrictions on the two vectors were, however, not very encouraging. We thus decided to leave this issue on the agenda for future research. Furthermore, we note that a cointegration model with an *unrestricted* constant imposes the restriction that the linear combination eliminating the stochastic trends also eliminates the deterministic linear trends. This may be an acceptable hypothesis in our case as a *restricted* trend is *in*significant at conventional levels in the reduced rank VAR reported below.

[7] It should be noted that the preferred VAR includes two additional impulse dummies that account for "outliers" in the $\mathrm{pa}_t$-equation in 1979:1 and 1979:2. These dummies probably capture effects of the governmental wage and price controls in 1979, cf. Bowitz and Cappelen (2001). However, these dummies do not significantly affect the conclusions from the cointegration analysis, see Boug et al. (2005).

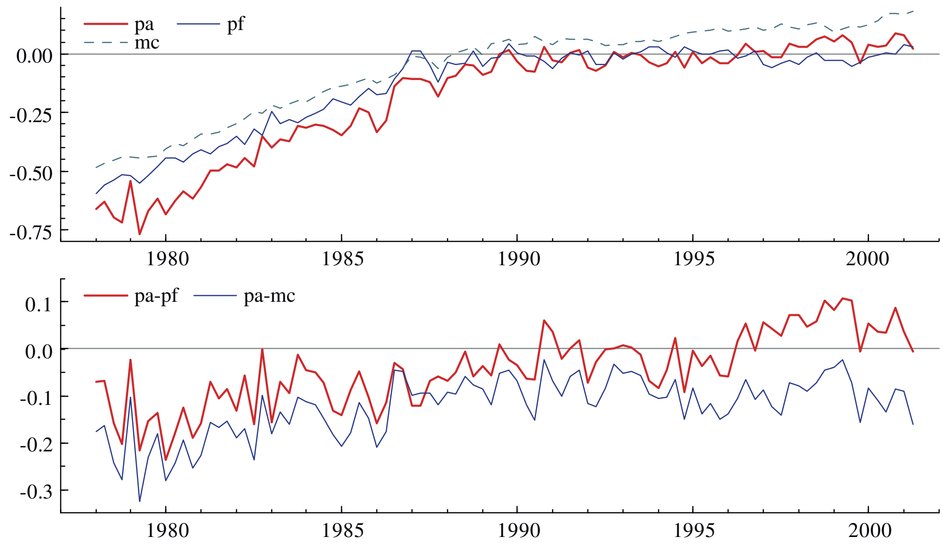


**Fig. 1** The log of export prices (pa), competing prices (pf) and marginal costs (mc)

**Table 1** Johansen's cointegration tests

Information set: (pa, pf, mc)
Eigenvalues: 0.220, 0.128, 0.040
Log likelihood value: 931.5

| | Statistics | | | | | |
|---|---|---|---|---|---|---|
| Hypothesis | $\lambda_{\text{max}}$ | $\lambda^{\text{a}}_{\text{max}}$ | 95% | $\lambda_{\text{trace}}$ | $\lambda^{\text{a}}_{\text{trace}}$ | 95% |
| $r = 0$ | 19.90 | 16.92 | 21.00 | 34.15* | 29.03 | 29.7 |
| $r \leq 1$ | 11.00 | 9.35 | 14.10 | 14.25 | 12.11 | 15.4 |
| $r \leq 2$ | 3.25 | 2.76 | 3.80 | 3.25 | 2.76 | 3.8 |

Estimate of the *unrestricted* cointegrating vector[a]

$$\text{pa} = \hat{\alpha}_0 + \underset{(0.252)}{0.525}\,\text{pf} + \underset{(0.256)}{0.585}\,\text{mc}$$

| Weak exogeneity tests[b] | pa | pf | mc |
|---|---|---|---|
| $\chi^2(1)$ | 7.409** | 2.396 | 3.269 |
| *P*-value | (0.007) | (0.122) | (0.071) |

$r$ denotes the cointegration rank, i.e., the number of cointegrating vectors. The $\lambda_{\text{max}}$ and $\lambda_{\text{trace}}$ statistics are the maximum eigenvalue and trace statistics, whereas $\lambda^{\text{a}}_{\text{max}}$ and $\lambda^{\text{a}}_{\text{trace}}$ are the corresponding statistics with a degrees-of-freedom-adjustment, cf. Reimers (1992). The 95 per cent quantiles are taken from Table 1 in Osterwald-Lenum (1992). It should be noted that the inclusion of impulse dummies in the VAR affects the asymptotic distribution of the reduced rank test statistics and therefore the critical values given in Osterwald-Lenum (1992) are only indicative. However, the bias induced by such deterministic components is supposed to be only minor as only two impulse dummies are included in our case. Asterisk $*$ and ** denote rejection of the null hypothesis at the 5% and 1% significance levels

[a] The figures in parentheses are standard errors

[b] The tests, which are asymptotically distributed as $\chi^2(1)$ under the null, are calculated under the assumption that $r = 1$

is not rejected by any of the statistics. Although the small sample adjusted tests do not indicate cointegration[8], recursive estimation of the eigenvalues supports our conclusion that there is a single cointegrating vector between pa, pf and mc, cf. Boug et al. (2005). The estimate of the *unrestricted* cointegrating vector (normalised on pa) is interpretable as an export price equation. The estimates of the export price elasticities (i.e., $\gamma_1$ and $\gamma_2$) are economically reasonable and statistically significant. Besides, the results of the weak exogeneity tests [see Johansen and Juselius (1990)] imply that the cointegrating vector enters the $\mathrm{pa}_t$-equation only, albeit the $\mathrm{mc}_t$-equation may be a borderline case. A possible interpretation of the latter result is that production costs increase due to, say, increased wages in the labour market. The wage increase in turn may lead to a depreciation of the Norwegian currency (due to a loss in the competitiveness) and a subsequent increase in the export price. Another channel from which the marginal costs may not be weakly exogenous is through the competing price. That is, the competing price affects the price index for material inputs that enters our marginal cost function, cf. Eq. (2).

We also notice that the sum of the unrestricted estimates of the export price elasticities (i.e., $\gamma_1+\gamma_2$) is not far from unity, as predicted by theory. To complete the cointegration analysis, we thus tested for, and could not reject, the existence of homogeneity between pa, pf and mc. Imposing the homogeneity restriction gives $\chi^2(1)$=0.487 (with a *P*-value of 0.485) and the following *restricted* estimate of the cointegrating vector (normalised on pa)

$$\mathrm{pa} = \hat{\alpha}_0 + \underset{(0.381)}{0.645}\mathrm{pf} + 0.355\mathrm{mc}, \tag{8}$$

where the standard error is in parenthesis. A sequence of $\chi^2(1)$ test statistics also confirms the validity of the homogeneity restriction for any sample size between 1985 and 1998, cf. Boug et al. (2005). The restrictions of long run homogeneity and weak exogeneity of pf and mc, both individually and jointly, were however rejected by the data. The *t*-value of 1.693 on the coefficient of pf is considered statistically significant because, for a one-sided test, its *P*-value is 0.047. Similarly, the recursively estimated parameter of pf in (8) is reasonably stable. To sum up, we interpret (8) as an export price equation that corresponds well with standard economic theory in terms of homogeneity and the fact that for a small open economy like Norway, foreign prices are expected to matter much, i.e., market power is low.[9]

## 5 Tests of the LQAC-model

We now proceed to test the empirical performance of (6) assuming a reduced rank VAR and long run homogeneity between pa, pf and mc (i.e., $\gamma=\gamma_1=1-\gamma_2$)

[8] Albeit $\lambda^{a}_{\mathrm{trace}}$ (with a degrees-of-freedom-adjustment) is a borderline case at the 5% level.

[9] Cuthbertson (1990) finds equal weights of the cost and competing price elasticity when homogeneity is imposed in the study of the UK manufacturing export price.

**Table 2** Tests of the LQAC-model using equations (6) and (7)

| | $\gamma = 0.2$ | | | $\gamma = 0.6$ | | | $\gamma = 0.8$ | | |
|---|---|---|---|---|---|---|---|---|---|
| $\beta$ | $\theta = 0.05$ | $\theta = 0.15$ | $\theta = 0.25$ | $\theta = 0.05$ | $\theta = 0.15$ | $\theta = 0.25$ | $\theta = 0.05$ | $\theta = 0.15$ | $\theta = 0.25$ |
| 0.90 | 856.6 | 853.2 | 849.7 | 857.4 | 854.1 | 850.5 | 857.8 | 854.4 | 850.8 |
| | (149.8) | (156.5) | (163.6) | (148.1) | (154.8) | (162.0) | (147.5) | (154.2) | (161.5) |
| 0.95 | 859.2 | 855.9 | 852.4 | 860.1 | 856.8 | 853.3 | 860.4 | 857.1 | 853.5 |
| | (144.6) | (151.2) | (158.1) | (142.9) | (149.5) | (156.5) | (142.2) | (148.8) | (156.0) |
| 0.99 | 861.2 | 857.9 | 854.5 | 862.0 | 858.8 | 855.3 | 862.3 | 859.1 | 855.6 |
| | (140.7) | (147.2) | (154.1) | (139.0) | (145.5) | (152.4) | (138.3) | (144.9) | (151.9) |

*Notes*: Figures are maximum values of the log likelihood of the cointegrated VAR *with* expectations restrictions imposed [log $L(\beta,\theta,\gamma)$], calculated by means of the procedure suggested by Johansen and Swensen (1999) and the reduced rank VAR reported in Sect. 4. Numbers in parentheses are $-2$ log likelihood ratio statistics computed by comparing the values in this table with the value 931.5, which is the maximum value of the log likelihood of the cointegrated VAR *without* the expectations restrictions imposed

in accordance with the evidence presented above. To do so, we make use of the procedure suggested by Johansen and Swensen (1999). The basic idea behind this method is to formulate the restrictions implied by the LQAC-model on a form so that, when imposed, the maximum value of the likelihood of the VAR model can be found. Comparing this value to the maximum value of the unrestricted likelihood yields a likelihood ratio test. That is, one has to work out the maximum likelihood estimator of the cointegrated VAR, both with and without the rational expectation restrictions imposed, in order to construct the likelihood ratio test. The approach of Johansen and Swensen (1999) requires that the restrictions contain no unknown parameters. In (6) and (7) there are unknown parameters $\beta, \theta$ and $\gamma$. Fixing parameters, we can proceed to find the maximum value of the likelihood $L(\beta, \theta, \gamma)$.

In principle, the maximum likelihood estimates for $\beta, \theta$ and $\gamma$ can be found by using an appropriate numerical optimiser. In our case it is sensible just to proceed with a grid search over the set of relevant parameter values for at least three reasons. First, since $0 < \beta < 1$ by definition, this constraint has to be incorporated in the optimisation. Second, only parameter values where $0.9 \leqslant \beta < 1, 0 \leqslant \theta \leqslant 1$ and $0 \leqslant \gamma \leqslant 1$ can be given an economic interpretation. Hence, estimates outside these parameter regions have to be discarded. Third, as we shall see below, knowledge of the actual shape of the likelihood surface contains valuable information. In the Appendix, we explain how (6) together with the VAR from the previous section may be formulated so that the methods in Johansen and Swensen (1999) can be applied to find $L(\beta,\theta,\gamma)$ with no rational expectations restrictions imposed on the deterministic part of the VAR (i.e., the constant, the seasonals and the impulse dummies).

Table 2 presents a sample of these calculations. The maximum value of the log likelihood varies between 849.7 and 862.3 with the parameter combination $\beta = 0.99$, $\theta = 0.05$ and $\gamma = 0.8$ producing the largest value. Compared to the value 931.5 from Table 1, which is the maximum value of the log likelihood of the cointegrated VAR without the rational expectation restrictions imposed, this yields a –2 log likelihood ratio statistic between 138.3 and 163.6. Thus, regardless of whether the parameter values $\beta, \theta$ and $\gamma$ are considered as fixed, corresponding to $\chi^2$ with 12 degrees of freedom, or estimated, corresponding to $\chi^2$ with 9 degrees of freedom, there is *overwhelming empirical evidence against* the LQAC-model (6).[10] We emphasise that the reported maximum likelihood ratio estimates are obtained with boundary values of the grid. In fact, the log likelihood is monotone in each of the parameters $\beta$, $\theta$ and $\gamma$. Thus, there may exist values outside the grid where the model is *not* rejected. As previously noted, such values would not match parameters of the model having reasonable economic interpretation.

It is possible to consider the distribution of the test statistic also for parameter values on the boundary of the parameter set. The asymptotic distribution of the likelihood ratio statistic can then typically be expressed as the difference between the suprema of two stochastic processes, see e.g. Andrews (2001). To find the critical values simulations are often needed. We have not pursued this line, since the numerical values of the likelihood ratio statistic are so large that it is unlikely that more accurate critical values will change the overall picture.

Also other evidence supports this claim. As pointed out by Campbell and Shiller (1987) among others, rational expectations models may have economic content even though they are rejected by formal statistical tests. In Engsted and Haldrup (1994), see also Engsted (2002, p. 343), it is explained that if the LQAC-model is correct, then

$$\Delta \mathrm{pa}_t - (\lambda - 1)\left(\mathrm{pa}_{t-1} - \mathrm{pa}^*_{t-1}\right) - (1-\lambda)\Delta \mathrm{pa}^*_t = (1-\lambda)\sum_{j=1}^{\infty} (\lambda\beta)^j E_t\left[\Delta \mathrm{pa}^*_{t+j}\right], \tag{9}$$

where $\lambda$ is the solution in the interval (0, 1) of the *characteristic* equation implied by (6). Hence, plotting the movements of the left hand side (LHS) of (9) together with the movements of the estimate of the right hand side (RHS) of (9) based on the *unrestricted* reduced rank VAR model, provides an informal method for evaluating the extent to which the LQAC-model fits the observed data. Figure 2 displays the case when the parameter values $\beta = 0.99$ and $\theta = 0.05$ are used. The picture clearly confirms the conclusion from the formal statistical tests.[11] In Sect. 8, we report some robustness analyses and show that assuming a stationary VAR around a breaking linear trend make no improvement in describing the data. Likewise, we demonstrate in the spirit of Gali and Gertler (1999) among others that a hybrid version of (6) allowing for additional lags of the targeted variable does not rescue the forward-looking model.

[10] Fanelli (2002) has recently proposed a test for the LQAC-model in an integrated VAR. We expect even stronger rejection of (6) using Fanelli's test as that test is based on additional restrictions to those following from the Euler equation.

[11] The discrepancies between LHS and RHS of (9) are even larger using other values of $\beta$ and $\theta$.

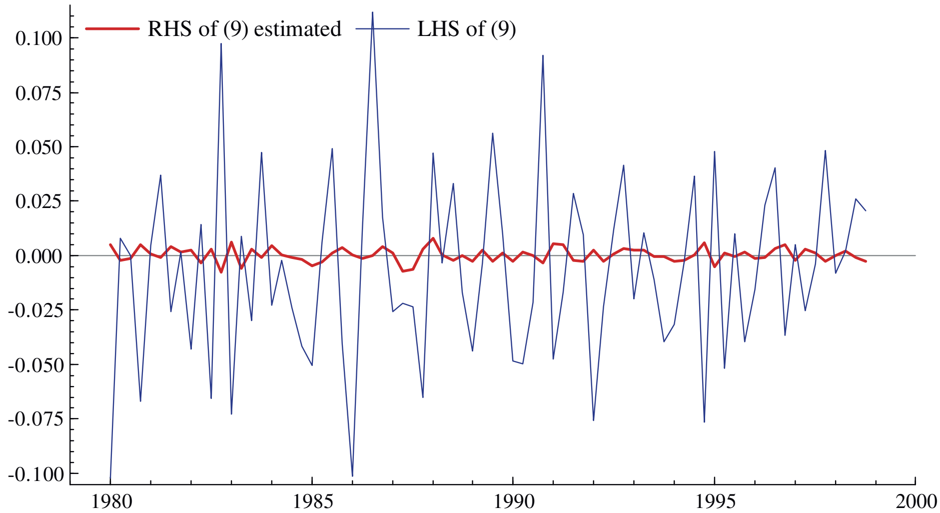


**Fig. 2** The fit of the LQAC-model

## 6 The EqCM-model and the Lucas critique

As noted in the introduction, rejection of the LQAC-model does not preclude Norwegian exporters from acting on alternative expectations based models in the formation of prices. To establish evidence for or against such a hypothesis, this section develops an EqCM-model of export prices and examines the relevance of the Lucas critique based on Hendry (1988). Suppose that agents are forward-looking, so that the value of pa depends on expected future values of pa* (the target value of pa). If we assume that firms use current and previous values of pa* to forecast future values of pa*, then in a backward-looking equilibrium-correction model for pa, the coefficients on pa* will be a combination of structural and expectational terms. If the estimated parameters for pa* are unstable, then the equilibrium-correction model for pa will also be unstable. Consequently, if the model for pa* is unstable but the equilibrium correction model for pa is stable, then pa must be purely backward looking. There are a number of published studies that use this methodology to test the Lucas critique, see e.g. Ericsson and Irons (1995).

The Lucas critique is likely to be relevant in the export price formation in Norway. The main channel through which exchange rate policies operate is the competing price pf. Although all prices and costs are measured in domestic prices, contracts are often made in foreign currency. Unexpected exchange rate shocks will change the profitability compared to the negotiated contract. Thus, exchange rate policy is something that producers should take into account in their export price formation. Although the prime policy for a long time was that of a fixed nominal exchange rate to a basket of European currencies and later to the ECU, the exchange rate has changed sometimes quite substantially in our sample period as a result of shocks to the economy.

During 1982 and 1984 there were two devaluations each year by around 3 and 2%, respectively, and in May 1986 the Norwegian currency devaluated by as much as 12%. In addition, there was a short period of a floating exchange rate regime after the turmoil in the European financial markets in December 1992, which led to a depreciation of approximately 4%.[12] After our sample period, monetary policy has changed more fundamentally as the exchange rate now is floating following the introduction of inflation targeting in late March 2001.[13]

Even if the Lucas critique may be theoretically important, it might not be so empirically, as argued by Leeper (1995). We shall turn to this issue shortly after having established an EqCM-model of export prices. Applying the same information set as in the cointegration analysis and using deviations from (8) as an EqCM, we proceed from the following model

$$\begin{aligned}\Delta \mathrm{pa}_t = \beta_0 + \sum_{i=1}^{4} \beta_{1i}\Delta \mathrm{pa}_{t-i} + \sum_{i=0}^{4} \beta_{2i}\Delta \mathrm{pf}_{t-i} + \sum_{i=0}^{4} \beta_{3i}\Delta \mathrm{mc}_{t-i} \\ + \delta[\mathrm{pa} - 0.645\mathrm{pf} - 0.355\mathrm{mc}]_{t-1} + \mathrm{dummy} + e_t,\end{aligned} \tag{10}$$

where $\Delta$ is the difference operator, the expression in $[\cdot]$ is the EqCM, dummy represents the impulse dummies (i.e., $D79(1)$ and $D79(2)$, cf. footnote 7 in Sect. 4) and the seasonals from the VAR and $e_t$ is the error term, assumed to be white noise. A simplified export price model that approximates the data well is presented in (11). It may be worth noting that (11) is derived from a single equation analysis rather than a system one. Following Boswijk and Urbain (1997), one may apply single equation analysis with the long run relationship(s) estimated and deduced from a VAR model in cases where the conditioning variables are error correcting but weakly exogenous for the short run parameters. Since the evidence of mc being weakly exogenous for the cointegration parameters is not particularly strong (see Table 1), initial estimations and subsequent simplifications of (10) excluding insignificant variables were conducted with and without instruments for $\Delta \mathrm{mc}_t$. This variable was, however, far from being significant in any of the regressions considered, and (11) thus omits contemporaneous effects from $\mathrm{mc}_t$, cf. Boug et al. (2005).

$$\begin{aligned}\Delta \mathrm{pa}_t = \text{const.} - \underset{(0.087)}{0.360}(\mathrm{pa}_{t-1} - \mathrm{pa}_{t-3}) - \underset{(0.090)}{0.292}\Delta \mathrm{pa}_{t-3} - \underset{(0.081)}{0.124}\Delta \mathrm{pa}_{t-4} \\ + \underset{(0.120)}{0.233}\Delta \mathrm{pf}_t + \underset{(0.262)}{0.731}\Delta \mathrm{mc}_{t-1} - \underset{(0.070)}{0.214}[\mathrm{pa} - 0.65\mathrm{pf} - 0.35\mathrm{mc}]_{t-1} \\ - \underset{(0.009)}{0.025}\mathrm{CS}_{1t} + \underset{(0.038)}{0.195}D79^{(2)}\end{aligned} \tag{11}$$

[12] See Bowitz and Cappelen (2001) for details of the Norwegian exchange rate policy over the last decades.

[13] Several Norwegian economists argue that the regime change occurred early in 1999. In any case, the policy change took place after 1998:4; the last observation used in our estimations.

$$\text{Method: OLS}T = 79(1979:2 - 1998:4)R^2 = 0.601\ \sigma = 3.57\% DW = 2.10$$
$$\text{AR}_{1-5}: F(5,65) = 1.145\ \text{ARCH}_{1-4}: F(4,62) = 0.843\ \text{NORM}: \chi^2(2) = 1.054$$
$$\text{HET}_1: F(14,55) = 1.663\ \text{HET}_2: F(35,34) = 1.301\ \text{RESET}: F(1,69) = 2.303.$$

The imposed term $-0.360(\text{pa}_{t-1} - \text{pa}_{t-3})$ is statistically accepted by the data as the estimated coefficients of $\Delta\text{pa}_{t-1}$ and $\Delta\text{pa}_{t-2}$ turned out to be very close. Below the EqCM-model we report several test statistics.[14] None of the diagnostics are significant at the 1% level. The EqCM appears in the model with a $t$-value of $-3.06$, hence adding force to the results obtained from the cointegration analysis. Empirical evidence of constancy of the EqCM-model may be judged from recursive test statistics shown in Fig. 3. Neither the one-step residuals with $\pm 2$ estimated equation standard errors (denoted $\pm 2$ SE *(t)* in the graph) nor the sequence of break point Chow tests at the 5% significance level reject constancy. Similarly, aside from a minor fluctuation in the early 1980s in some of the estimates, all recursive estimates vary little, especially relative to their estimated uncertainty, cf. Boug et al. (2005). Noticeably, (11) implies rejection of *dynamic* homogeneity and consequently a negative relationship between inflation and the mark-up, a finding which is in line with previous studies based on European and American data, see e.g. Bénabou (1992), Blanchard and Muet (1993), Bowitz and Cappelen (2001) and Banerjee and Russell (2004).

We now consider empirical evidence of nonconstancy of the marginal process for $\Delta\text{pf}_t$. Noticeably, no marginal model for $\Delta\text{mc}_t$ needs to be considered because it enters (11) with one lag. Starting with a fifth-order autoregressive process for $\Delta\text{pf}_t$ and simplifying, we obtain the following specification (with standard errors in parentheses)

$$\Delta\text{pf}_t = \text{const.} - \underset{(0.106)}{0.244}\Delta\text{pf}_{t-1} - \underset{(0.009)}{0.021}\text{CS}_{2t} \tag{12}$$

$$\text{Method: OLS } T = 75(1978:3 - 1998:4)\ R^2 = 0.118\ \sigma = 3.33\% DW = 2.05$$
$$\text{AR}_{1-5}: F(5,74) = 0.654\ \text{ARCH}_{1-4}: F(4,71) = 2.145\ \text{NORM}: \chi^2(2) = 2.058$$
$$\text{HET}_1: F(3,75) = 2.989\ \text{HET}_2: F(4,74) = 2.255\ \text{RESET}: F(1,78) = 0.057.$$

Sequentially estimated equation standard errors increase by about 40% during the sample. In essence, a marked worsening of fit, as also detected by the significant break point Chow statistics at the 5% level shown in Fig. 4.

[14] Estimated standard errors are in parentheses, and $T$, $R^2$, $\sigma$ and DW are the number of observations, the squared multiple correlation coefficient, the residual standard error and the Durbin–Watson statistic, respectively. Further, $\text{AR}_{1-5}$ is the Harvey (1981) test for fifth-order residual autocorrelation, $\text{ARCH}_{1-4}$ is the Engle (1982) test for 4th order autoregressive conditional heteroskedasticity in the residuals, NORM is the normality test described in Doornik and Hansen (1994), $\text{HET}_1$ is the White (1980) test for residual heteroskedasticity, $\text{HET}_2$ tests whether the squared residuals depend on the levels, squares and cross products of the regressors [cf. White (1980)] and RESET tests for functional form misspecification [cf. Ramsey (1969)].

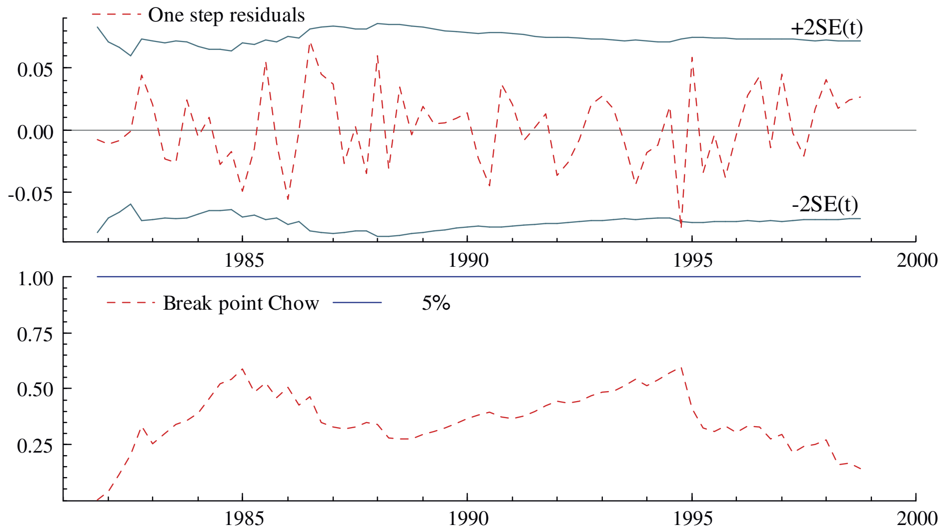


**Fig. 3** Recursive test statistics of (11)

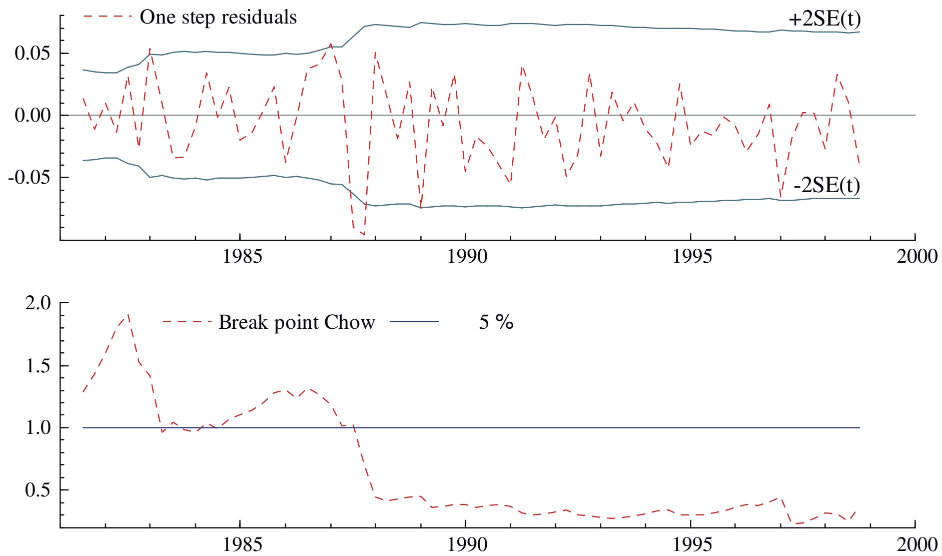


**Fig. 4** Recursive test statistics of (12)

The instabilities in $\Delta pf_t$ probably relate to the mentioned exchange rate changes during the 1980s.

Following Hendry (1988), the nonconstancy of the marginal model for $\Delta pf_t$ together with the constancy of the conditional model for $\Delta pa_t$ has two important implications. First, the Lucas critique is unlikely to apply to (11).[15] This implication holds even if the $\Delta pf_t$-equation ignores *some* information relevant

[15] According to Engle and Hendry (1993), one may also evaluate the Lucas critique by making the marginal model of $\Delta pf_t$ constant through dummy variables and testing the significance of such dummies in the conditional model of $\Delta pa_t$. The significance of such dummies in (11) would then be evidence that the Lucas critique *is* empirically relevant. In our case the dummies $D87(3)$ and

to the process generating competing prices, cf. Hendry (1988). The combination of empirical evidence is thus inconsistent with the hypothesis that Norwegian exporters act on expectations based models. Second, our findings imply that we cannot reject that $\Delta \mathrm{pf}_t$ is weakly exogenous for the short run parameters in (11)[16], parameters which thereby are consistently estimated by OLS.

The evidence in Lindé (2001) that tests based on Hendry (1988) are not capable of detecting the significance of the Lucas critique in small samples is not relevant here. Lindé (2001) shows that despite *unstable* parameter estimates of the conditional and the marginal model in two particular examples, tests of the kind used here is not able to detect the instabilities. Consequently, unwarranted conclusions about the relevance of the Lucas critique follow in his case. In our case, however, the situation is quite different. We show that it is not difficult to establish a price equation with *stable* parameter estimates that make good sense from an economic point of view. This stability is established despite the fact that there have been some changes in monetary policy during the sample period.

We have demonstrated that (11) exhibits historical constancy *in-sample*. Now, we study the *out-of-sample* forecasting performance of the estimated export price equation to shed further light on its robustness with respect to monetary policy regime shifts. If the export price behaviour has changed significantly following the introduction of inflation targeting, we should expect instabilities in the estimated $\Delta \mathrm{pa}_t$ -equation as, for example, indicated by poor *out-of-sample* forecasting ability. To assess the forecasting performance of the estimated export price equation, we employ twenty four quarters (1999:1–2004:4) of *out-of-sample* observations, including the period after the formal change in monetary policy regime (2001:2). Figure 5a–c depicts actual values of $\mathrm{pa}_t$ together with dynamic forecasts, four-step-ahead forecasts and one-step-ahead forecasts of $\mathrm{pa}_t$, respectively, adding bands of 95% confidence intervals to each forecast in the forecasting period.[17] Except 1999:4 in Fig. 5c, which is a borderline case, the actual values of $\mathrm{pa}_t$ stay clearly within their corresponding confidence intervals over the forecasting period. Also, a Chow-test statistic of parameter constancy between the sample and the forecasting periods, cf. Hendry and Doornik (2001, p. 241), is far from being significant with $F[24, 69]$=0.798 and the corresponding $P$-value of 0.727. Thus, the *out-of-sample* forecasting ability of (11) is quite good

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$D$87(4), which make (12) stable, are *in*significant when added both individually and jointly to (11). Also, the dummy $D$79(2) is far from being significant when included in (12).

16 Adding the predicted counterpart to $\Delta \mathrm{pf}_t$ from the VAR in (11) yields a $P$-value of 0.902, which provides further evidence that $\Delta \mathrm{pf}_t$ is weakly exogenous for the short run parameters in (11).

17 The forecast for period $s$ is $\hat{y}_s = x_s' \hat{\beta}_t$, where $x_s$ is the observed value of $x$ for period $s$, $\hat{\beta}_t$ is estimated from the first $t$ observations of data and $s > t$. In our case $s$ spans the period 1999:1–2004:4, while $t$ covers the period 1978:1–1998:4. Although forecasts for the *level* $\mathrm{pa}_t$ could be derived from (11), the forecasts in Fig. 5a–c (notice the different scales in the figures) were obtained directly by re-expressing the dependent variable as $\mathrm{pa}_t - \mathrm{pa}_{t-1}$ and estimating rather than imposing the unit coefficient on $\mathrm{pa}_{t-1}$, cf. Hendry and Ericsson (1991, p. 852).

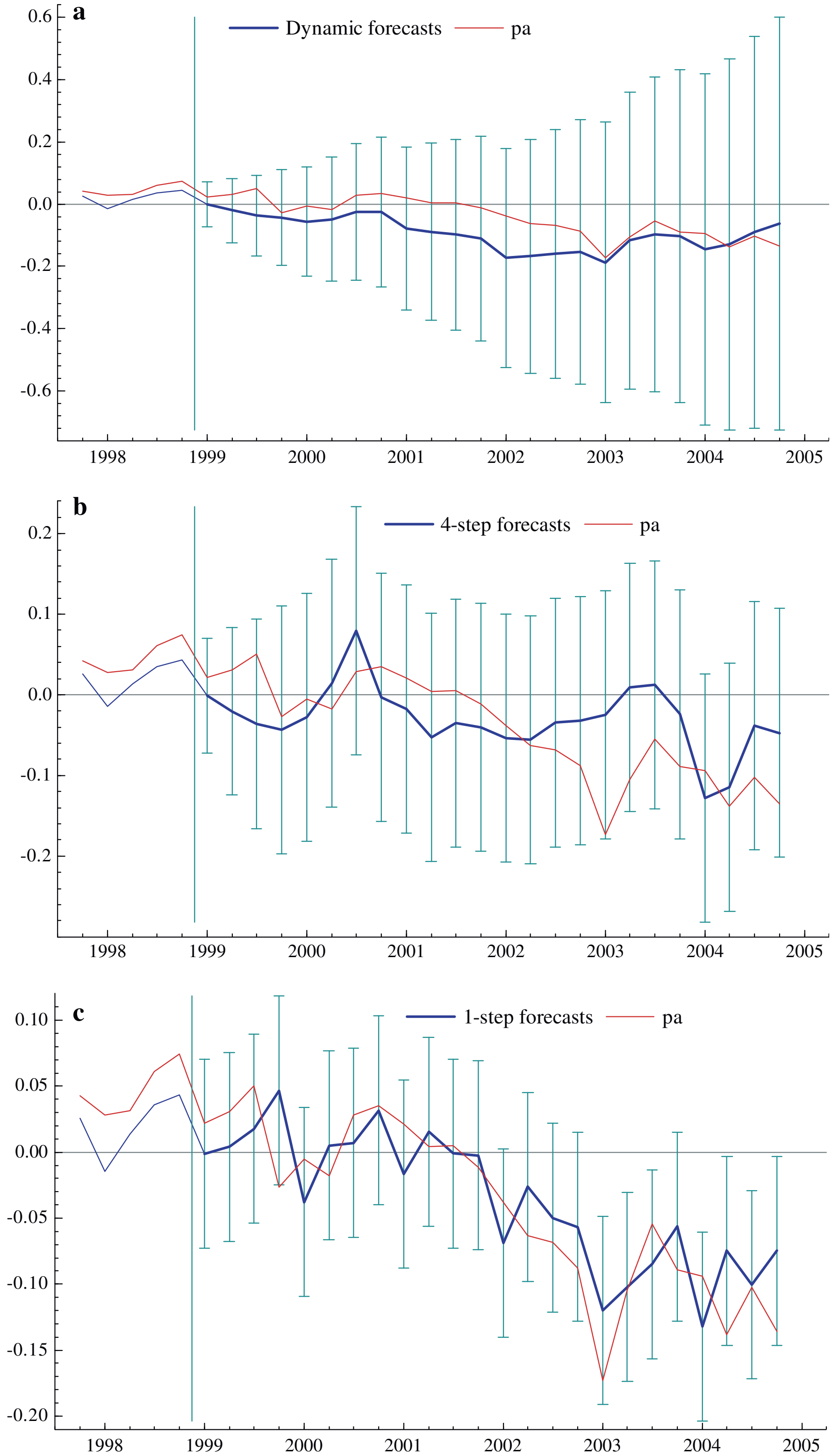


**Fig. 5** **a** Actual values of $pa_t$ and dynamic forecasts with 95% bands. **b** Actual values of $pa_t$ and four-step forecasts with 95% bands. **c** Actual values of $pa_t$ and one-step forecasts with 95% bands

despite a major regime change in monetary policy.[18] The regime robustness is inconsistent with the Lucas-critique being quantitatively important in our case.

A possible objection to the preferred price equation concerns the economic interpretation of the rather complicated dynamics inherent in the model that explicitly is not justified by economic theory. However, given the rejection of the LQAC-model, we have let the dynamics be determined by data. The dynamics implied by (11) is shown in Fig. 6.[19] For mc there is a rapid adjustment towards the long run response of 0.35. Within two years most of the dynamic adjustment is complete.[20] The short run overshooting can be interpreted as an effect of contracts containing some elements of indexation based on costs.[21] In the longer run these contracts are renegotiated and producers take into account the effect of foreign competition when setting the prices. Regarding the competing price (pf), the estimated impact elasticity of 0.23 is considerably smaller than its long run counterpart of 0.65. Accordingly, Norwegian firms seem to smooth the export prices with respect to changes in the competing prices. The apparent slow adjustment of pa to shifts in pf may reflect that the Norwegian exchange rate was fixed within certain bands for the most part of the sample period. If it is costly to change the export price, it may be wise to adjust slowly to changes in the competing price that are caused by temporary and *small* fluctuations in the exchange rate.

## 7 Conclusions

In this paper we have undertaken several tests to evaluate the empirical performance of both the forward-looking LQAC-model and the EqCM-model using data for the Norwegian export price of machinery. The optimal price is determined as a mark-up over marginal costs and marginal costs are based on a fully specified cost function in line with theory. Using multivariate cointegration methods we established a single long run relationship for the export price with the competing price and marginal costs as its long run determinants. Next, we tested the LQAC-model's ability to explain short run dynamics in the export price using two plausible, underlying VAR models. The results, however, were not supportive of the LQAC-model. An augmented model that allows for additional lags of the targeted variable does not do a better job of describing the data. By way of contrast, the data seems to be more consistent with

[18] The estimated model (11) is also (as indicated by recursive methods) constant when the sample is extended to include the period 1999:1-2004:4.

[19] To make it explicit, the dynamics reflect the responses of the export price inherent in (11) to a one per cent impulse in the marginal costs and competing price, respectively. For instance, the immediate impulse of pa to a one per cent increase in pf is 0.23% as seen by the estimated coefficient of $\Delta \mathrm{pf}_t$.

[20] Interestingly, Fuhrer (1997) reports dynamics of an aggregated backward-looking inflation model for USA that are characterised by cyclical behaviour lasting for more than 20 years.

[21] In fact, we know that such contracts do exist for machinery as Statistics Norway often are asked by agents to provide cost and price indices used for indexation purposes.

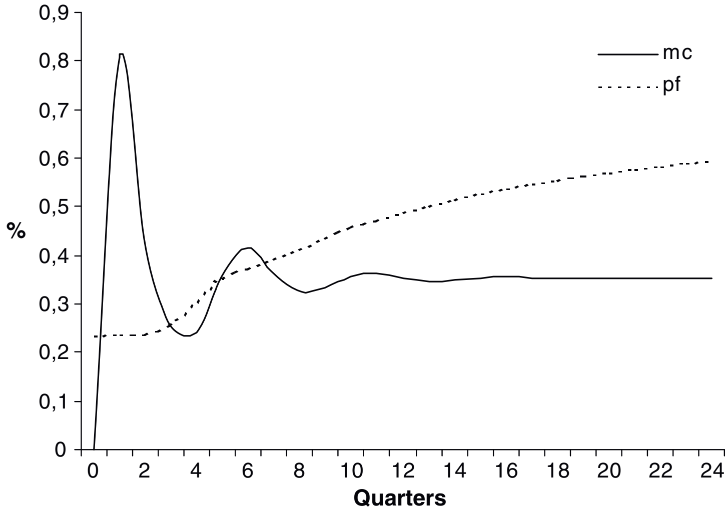


**Fig. 6** Dynamics implied by (11). Effect on pa (in per cent) of a 1% change in pf and mc (ceteris paribus)

an EqCM-model. The EqCM-model is found reasonably stable despite policy regime shifts and other economic shocks over the past decades, supporting the conclusion that the Lucas critique lacks force in our case. Likewise, the fact that the EqCM-model forecasts well post-sample and during a major change in the policy regime, certainly is a strong evidence in favour of this model.

Given the *clear* rejection of the LQAC-model, we feel compelled to accept the EqCM-model even though the latter has the dynamics determined by data rather than by well-defined theories.[22] At the same time, we emphasise the aspect that expectations here mean conditional expectations in the model the LQAC-analyses are based upon. Naturally, the question then arises as to which information set is the appropriate one to use in the testing of the model. We have shown with reasonable assumptions about the information set contained in the variables of the VAR models that the forward-looking hypothesis is overwhelmingly rejected.

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[22] This is not particularly controversial. Olivier Blanchard among others sees no alternative to using ad hoc dynamics until more adequate theory is developed, cf. Sims (2002, p. 59).

## 8 Appendix

### 8.1 Details about the procedure suggested by Johansen and Swensen (1999)

As noted in the text, the set of restrictions implied by the LQAC-model under rational expectations needs to be formulated on a specific form so that the methods in Johansen and Swensen (1999) can be applied. Letting $X_t = (\mathrm{pa}_t, \mathrm{pf}_t, \mathrm{mc}_t)'$ we may write the set of restrictions in (6) in compact form as

$$c_1' E_t[X_{t+1}] + c_0' X_t + c_{-1}' X_{t-1} = 0, \tag{13}$$

where $c_1' = (\beta, 0, 0), c_0' = -(1 + \beta + \theta, -\theta\gamma_1, -\theta\gamma_2)$ and $c_{-1}' = (1, 0, 0)$. We observe that (13) for fixed values of $\beta$ , $\theta$ and $\gamma = \gamma_1 = 1 - \gamma_2$ contains *linear* restrictions involving the conditional expected value of the observations *one-step-ahead* and the present and lagged observed values. In other words, the restrictions satisfying the Euler equation entail restrictions on the long run relationships as well as some restrictions on the short run parameters. When the information set is specified the conditional expectation $E_t[X_{t+1}]$ can be worked out and the restrictions stated explicitly.

Recall that a VAR of the form

$$X_t = A_1 X_{t-1} + \cdots + A_4 X_{t-4} + \Phi D_t + \varepsilon_t, t = 1979:1, \ldots, 1998:4, \tag{14}$$

with reduced rank equal to one was shown to have satisfactory statistical properties. Here $D_t$ represents the constant, the centred seasonals and the impulse dummies, as explained in the text. Since the model contains dummy variables a slightly different formulation of (13) makes more sense. We first notice that

$$E_t[X_{t+1}] = \Phi D_{t+1} + A_1 X_t + \cdots + A_4 X_{t-3}. \tag{15}$$

It seems reasonable to consider restrictions involving the part of the forecasts of future values that do not involve the deterministic part, and test restrictions of the form

$$c_1' E_t[X_{t+1} - \Phi D_{t+1}] = -c_0' X_t - c_{-1}' X_{t-1}. \tag{16}$$

Inserting the previous expression for $E_t[X_{t+1}]$, (16) becomes

$$c_1' A_1 X_t + \cdots + c_1' A_4 X_{t-3} = -c_0' X_t - c_{-1}' X_{t-1}, \tag{17}$$

so there are no restrictions on the coefficients of the deterministic part of the VAR. Hence, the restrictions satisfied by the Euler equation and deduced from the LQAC-model take the following form in terms of (14):

$$\begin{aligned} c_1' A_1 &= -c_0', \\ c_1' A_2 &= -c_{-1}', \\ c_1' A_3 &= c_1' A_4 = 0. \end{aligned} \tag{18}$$

We can then use the same procedure as described in Johansen and Swensen (1999) with the exception that the deterministic terms are included without restrictions when the marginal part of the model is estimated. Since the method is based upon likelihood ratio tests we have to find the maximum likelihood estimators of the coefficients of the VAR under the imposed restrictions. Referring to the notation in Johansen and Swensen (1999), $b' = c_1' = (\beta, 0, 0)$ in our case, so we may set the $3 \times 2$ matrix $a$ equal to the matrix $(e_2, e_3)$, where $e_2$ and $e_3$ are the $3 \times 1$ unit vectors having one as the second and third element, respectively, and else consisting of zeros. Since the rank of the matrix $c_1'$, $q$, and the reduced rank of the VAR model, $r$, both are equal to unity in our case, we first determine the conditional model by regressing $a'\Delta X_t = (\Delta \mathrm{pf}_t, \Delta \mathrm{mc}_t)'$ on $b'\Delta X_t - c_{-1}'\Delta X_{t-1}, d'X_{t-1}, \Delta X_{t-1}, \ldots, \Delta X_{t-3}$, the constant, the seasonals and the impulse dummies where $d' = -(c_1 + c_0 + c_{-1})' = \theta[1, -\gamma, -(1-\gamma)]'$. Notice that $b'\Delta X_t - c_{-1}'\Delta X_{t-1} = (\beta \Delta \mathrm{pa}_t - \Delta \mathrm{pa}_{t-1}, 0, 0)$ and $d'X_{t-1} = \theta[\mathrm{pa}_{t-1} - \gamma \mathrm{pf}_{t-1} - (1-\gamma)\mathrm{mc}_{t-1}]$. To find the contribution from the marginal model we then, in accordance with the modifications above, regress $b'\Delta X_t - d'X_{t-1} - c_{-1}'\Delta X_{t-1} = \beta \Delta \mathrm{pa}_t - \Delta \mathrm{pa}_{t-1} - \theta \mathrm{pa}_{t-1} + \theta\gamma \mathrm{pf}_{t-1} + \theta(1-\gamma)\mathrm{mc}_{t-1}$ on the constant, the seasonals and the impulse dummies. Let the residuals from the first regression be $R_{1t}$ and let the mean sum of the matrices $R_{1t}R_{1t}$' be $S_{11}$. Compute the product of the determinant of $S_{11}$ and the mean sum of squares from the last regression. Finally, the value of the maximised likelihood for fixed values of $(\beta, \theta, \gamma)$ is this product divided by $b$'$b$ det$(a'a) = \beta^2$. A program in GAUSS, available on our web site: http:/www.ssb.no/forskning/ansatte/, was written in order to carry out the Johansen and Swensen (1999) procedure.

## 8.2 Robustness analyses

As we have seen, the LQAC-model (6) does not fit the data. To explore the robustness of this result we have considered two modifications. The first is to fit a stationary VAR around a breaking linear trend, with a break point around 1990, i.e., a model of the form

$$X_t = B_1 X_{t-1} + \cdots + B_k X_{t-k} + \phi D_t + a_0 W_{1t} + a_1 W_{2t} + e_t. \tag{19}$$

where $W_{1t}$ defines a linear, deterministic time trend and $W_{2t}$ is a linear spline function, i.e., $W_{2t} = 0$ if $t \leq b$ and $W_{2t} = t - b$ if $b<t$ (where $b$ is the assumed

break point in the time trend). A specification search for determining the lag length $k$ suggests $k = 4$, as for the reduced rank VAR case. Similarly to the reduced rank VAR, we do not consider restrictions implied by the rational expectation hypothesis that include the deterministic part of the model. Using the same algebraic steps as explained above, the set of restrictions we test in terms of (19) takes the following form:

$$\begin{aligned} c_1' B_1 &= -c_0', \\ c_1' B_2 &= -c_{-1}', \\ c_1' B_3 &= c_1' B_4 = 0. \end{aligned} \tag{20}$$

Since the estimation of (19) does not involve a reduced rank VAR regression, we test the set of restrictions directly by means of a standard Wald test, its statistic being asymptotically distributed as $\chi^2(\cdot)$ with degrees of freedom given in parenthesis. In the most restricted model with a break point in the time trend assumed in 1990(1), we used $\beta = 0.95, \theta = 0.15$ and $\gamma = 0.5$ as one plausible combination of parameter values. The restrictions are overwhelmingly rejected also in this case, with a Wald test statistic of $\chi^2(12) = 378.3$ and a $P$-value of essentially zero. Performing the same exercise with a less restricted model, in which $\beta$ is considered as the only fixed parameter, yields a Wald test statistic corresponding to 9 degrees of freedom of 90.3 for $\beta = 0.99$, 97.4 for $\beta = 0.95$ and 108.7 for $\beta = 0.90$. Once again, we find strong rejection of the restrictions. Altering the break point in the time trend between 1989(1) and 1991(1) produces very little variation in the results. To sum up, the results of the LQAC-model (6) are robust to the modification in (19) of the underlying VAR. This conclusion is not sensitive to the choice of using variable unit costs (vc) instead of mc.

The second modification exploits the fact that (6) and its restrictions have some similarities with an approach used in the New Keynesian Phillips curve models of inflation. In the literature on sticky prices due to quadratic costs of changing nominal prices, a number of studies on aggregate data have found that simple models similar to (6) fit the data poorly, see Fuhrer (1997) and Estrella and Fuhrer (2002). One frequently used approach to modify the strict version of the forward-looking price equation is to assume two types of firms in the economy, cf. Gali and Gertler (1999). A fraction $\omega$ of all firms are backward-looking and use a simple rule of thumb for forecasting inflation, while the rest $(1 - \omega)$ are forward-looking in line with (6). Like the forward-looking firms the backward-looking firms behave in accordance with optimal rules in steady state (pa=pa*), but expect the future inflation to be a weighted average of inflation in the two previous periods. Thus, inspired by Gali and Gertler (1999), we modify (6) and specify the following hybrid model:

$$\Delta \mathrm{pa}_t = (1 - \omega)\beta E_t[\Delta \mathrm{pa}_{t+1}] + \omega\rho \left[(1 - \eta)\Delta \mathrm{pa}_{t-1} + \eta \Delta \mathrm{pa}_{t-2}\right] - \theta(\mathrm{pa}_t - \mathrm{pa}_t^*), \tag{21}$$

**Table 3** Tests of the LQAC-model using equations (21) and (7), $\beta = 0.99, \theta = 0.05$ and $\gamma = 0.8$

| | $\eta = 0.0$ | | $\eta = 0.5$ | | $\eta = 1.0$ | |
|---|---|---|---|---|---|---|
| $\omega$ | $\rho = 0.5$ | $\rho = 1.0$ | $\rho = 0.5$ | $\rho = 1.0$ | $\rho = 0.5$ | $\rho = 1.0$ |
| 0.0 | 862.3 | 862.3 | 862.3 | 862.3 | 862.3 | 862.3 |
| | (138.3) | (138.3) | (138.3) | (138.3) | (138.3) | (138.3) |
| 0.4 | 785.9 | 783.5 | 786.7 | 785.5 | 787.1 | 785.8 |
| | (291.2) | (296.0) | (289.5) | (291.9) | (288.8) | (291.4) |
| 0.8 | 658.8 | 656.1 | 659.5 | 658.2 | 659.6 | 657.7 |
| | (545.4) | (550.8) | (543.9) | (546.6) | (543.8) | (547.6) |

*Notes:* Figures are maximum values of the log likelihood of the cointegrated VAR *with* expectations restrictions imposed [log $L(\beta, \theta, \gamma, \omega, \rho, \eta)$], calculated by means of the procedure suggested by Johansen and Swensen (1999) and the reduced rank VAR reported in Sect. 4. Numbers in parentheses are $-2\log$ likelihood ratio statistics computed by comparing the values in this table with the value 931.5, which is the maximum value of the log likelihood of the cointegrated VAR *without* the expectations restrictions imposed

where the parameter $\eta$ represents the weight of inflation in the two previous periods and $\rho$ is the coefficient of a univariate regression of inflation on lagged inflation. In principle, more lags of inflation may be included in (21). However, as discussed in Sect. 4, a VAR model of order 4 fits the data well. Therefore, two previous values of inflation are the maximum possible we can deal with empirically. This means that (21) is the most general forward-looking rule allowed, as long as only linear combinations of observed inflation and a forecast of next period inflation are considered.

The formulation (21) nests several interesting special cases. If the backward-looking firms expect next period inflation to be equal previous period inflation only and not some average of inflation in the two previous periods, then $\eta = 0$. If all firms are backward-looking, then $\omega = 1$. The pure forward-looking model (6) remains if $\omega = 0$. Hence, (21) is a more flexible, but ad hoc way, of modelling prices than (6) as it allows for both forward-looking and backward-looking firms. Some investigations reveals that we obtain the largest values of the likelihood with $\beta = 0.99, \theta = 0.05$ and $\gamma = 0.8$ and $\omega = 0.0, 0.2, \ldots, 0.8$, $\rho = 0.5, 1.0$ and $\eta = 0.0, 0.25, \ldots, 1.0$ as a grid when evaluating the alternative model (21) together with (7). Table 3 presents the main results. A key feature of the results, in correspondence to those in Table 2, is that the parameter combination $\beta = 0.99, \theta = 0.05, \gamma = 0.8$ and $\omega = 0.0$, regardless of different values of the parameters $\rho$ and $\eta$, produces the largest value of the likelihood, namely the value 862.3. Thus, no improvements of the fit of the pure forward-looking model are made once lagged inflation is allowed for and the overall conclusion from the analysis of (6) applies also in this case.

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# 5 Chapter 5

## Exchange rate pass-through in a small open economy: the importance of the distribution sector

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RESEARCH ARTICLE

# Exchange Rate Pass-through in a Small Open Economy: the Importance of the Distribution Sector

**Pål Boug · Ådne Cappelen · Torbjørn Eika**

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**Abstract** The degree of exchange rate pass-through to domestic goods prices has important implications for monetary policy in small open economies with floating exchange rates. Evidence indicates that pass-through is faster to import prices than to consumer prices. Price setting behaviour in the distribution sector is suggested as one important explanation. If distribution costs and trade margins are important price components of imported consumer goods, adjustment of import prices and consumer prices to exchange rate movements may differ. We present evidence on these issues for Norway by estimating a cointegrated VAR model for the pricing behaviour in the distribution sector, paying particular attention to exchange rate channels likely to operate through trade margins. Embedding this model into a large scale macroeconometric model of the Norwegian economy, which inter alia includes the pricing-to-market hypothesis and price-wage and wage-wage spirals between industries, we find exchange rate pass-through to be quite rapid to import prices and fairly slow to consumer prices. We show the importance of the pricing behaviour in the distribution sector in that trade margins act as cushions to exchange rate fluctuations, thereby delaying pass-through significantly to consumer prices. A forecasting exercise demonstrates that exchange rate pass-through to trade margins has not changed in the wake of the financial crises and the switch to inflation targeting. We also find significant inflationary effects of exchange rate changes even in the short run, an insight important for inflation targeting central banks.

**Keywords** Exchange rate pass-through · Pricing behaviour · The distribution sector · Econometric modelling and macroeconomic analysis

**JEL classifications** C51 · C52 · E31 · F31

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## 1 Introduction

Much of the literature on the new open economy macroeconomics is based on models that feature rational expectations, optimizing agents and imperfect competition in markets for goods and possibly also labour. Small new Keynesian open economy models with several or all of these ingredients are popular when analysing analytically exchange rate pass-through and effects of monetary policy, see e.g. Svensson (2000) and Galí and Monacelli (2005) among others. However, it is often necessary to introduce ad hoc based micro-behaviour in order for these models to reproduce essential empirical aspects of real world data, typically by introducing backward-looking behaviour in the new Keynesian Phillips curve. The new open economy literature is based on the assumption of monopolistically competitive pricing behaviour, whereas the standard assumption of the "old" open economy models is price-taking behaviour in international markets, cf. Aukrust (1977) and Lindbeck (1979).[1] The pricing-to-market hypothesis introduced by Krugman (1987) and others, based on the assumptions of imperfect competition, nominal rigidities and market segmentation, is now the standard workhorse of the new open economy literature, see e.g. Atkeson and Burstein (2008) and Bugamelli and Tedeschi (2008).

In open economy models the degree of exchange rate pass-through − the responsiveness of import prices to changes in the nominal exchange rate − plays a vital role. Studies in the new open economy literature typically draw a distinction between producer currency pricing (PCP) and local currency pricing (LCP) when analysing exchange rate pass-through to domestic prices, see e.g. Galí and Monacelli (2005) and Devereux and Engel (2003). According to PCP, prices on internationally traded goods are set in the currency of the producer (exporter). If PCP holds, producers do not change their prices frequently, whereas consumers (and importers) face prices that vary one-for-one with nominal exchange rate changes (due to full pass-through). In this framework, changes in the nominal exchange rate are passed on to the terms of trade and consumers demand for home relative to foreign produced goods. LCP, on the other hand, is a price setting strategy where prices are set in the currency of the consumer, with no (or limited) pass-through of nominal exchange rate changes to import prices, at least in the short run. Thus, there may be only small effects from exchange rate changes to producer costs (to the extent that production is based on imported materials) as well as to consumer prices (to the extent that consumption is based directly on imported goods and services).

Some investigators use evidence of limited exchange rate pass-through to consumer prices as a justification for models with local currency pricing, see for instance Betts and Devereux (1996), Engel (2000) and Engel and Rogers (2001). Goldberg and Knetter (1997) emphasise this evidence as consumer prices often are found to be less affected by changes in exchanges rates than export prices. According to Engel and Rogers (2001), possible explanations and failure of the law of one price are tariff and non-tariff barriers to trade, transportation costs and non-traded inputs such as marketing and other distribution services that are part of final goods prices, but not to

[1] See Rogoff (1996) and Goldberg and Knetter (1997) for surveys about the evidence of systematic failure of the law of one price to hold for internationally traded goods. Persistent deviations from long run purchasing power parity are also found by e.g. Engel (2000) and Chen and Rogoff (2003).

the same extent part of prices of imported or exported goods. Obstfeld and Rogoff (2000) argue that correlations between changes in terms of trade and exchange rates for a large sample of countries are consistent with models of producer currency pricing. However, they argue that local currency pricing is relevant for retail prices, while prices on imported goods faced by retailers react to fluctuations in exchange rates as these prices are based on producer currency pricing. Burstein et al. (2003) consider models in which the pass-through to consumer prices is lower than to import prices as a result of local distribution costs in the wholesale and retail trade sector. Taylor (2000) puts forth the view that the extent to which a firm matches exchange rate movements by changing its own price depends on how persistent the movements are expected to be. For a retail firm that adds services to its imports, a depreciation of the exchange rate raises the costs of the imports evaluated in domestic currency. If the depreciation is viewed as temporary, the retail firm passes through less of the depreciation to its own price. Hence, less persistent exchange rate movements lead to smaller exchange rate pass-through to consumer prices.

In this paper, we present empirical evidence on exchange rate pass-through for the Norwegian economy by estimating a cointegrated VAR model for trade margins in the distribution sector. The degree and speed of exchange rate pass-through to retailers' trade margins are important for inflation dynamics as trade margins make up close to 30 % of the official consumer price index. We assume monopolistically competitive pricing behaviour when modelling prices, but do not consider forward-looking behaviour as this hypothesis is found to be clearly at odds with Norwegian data, see e.g. Bjørnstad and Nymoen (2008), Bårdsen et al. (2005, p. 145) and Boug et al. (2006). The cointegrated VAR model for trade margins is then analysed within an existing large scale macroeconometric model of the Norwegian economy assuming a 10 % depreciation of the exchange rate on a permanent basis.[2] By using the macroeconometric model, which inter alia includes the pricing-to-market hypothesis and price-wage and wage-wage spirals between industries, we are able to examine exchange rate pass-through to import prices, production costs, mark-ups and consumer prices for a large number of commodities and industries. Unlike studies in the new open economy literature, which typically are based on partial analyses of aggregated single-equation models, we thus take account of numerous channels through which the exchange rate is likely to operate in a small, open economy like the Norwegian.

Model simulations show that exchange rate pass-through is quite rapid to import prices and fairly slow to consumer prices in the Norwegian economy. We demonstrate that pass-through to consumer prices is not complete even within a ten-year horizon, a finding which may support the LCP hypothesis. The importance of the distribution sector is clearly apparent as trade margins act as cushions to exchange rate fluctuations in the short run, thus limiting the extent of exchange rate pass-through to consumer prices. If domestic inputs to the distribution sector are quantitatively important, then tradable goods sold to consumers include national value added (retail services) that may explain why there is incomplete pass-through. Likewise, imports

[2] A full description of the large scale macroeconometric model is beyond the scope of this paper. We refer to Bowitz and Cappelen (2001), Boug et al. (2006), Boug and Fagereng (2010), Benedictow and Boug (2012), Hungnes (2011), Jansen (2012) and Boug et al. (2013) for descriptions of main parts of the model.

as intermediate goods that together with domestic inputs produce final goods sold to consumers may also contribute to limited pass-through of exchange rate movements to consumer prices. We also present evidence that the exchange rate pass-through in the retailers' price setting has not changed significantly following the exchange rate volatility during the financial crisis and the shift in monetary policy to inflation targeting in 2001.

The rest of the paper is organised as follows: Section 2 outlines the main channels of exchange rate pass-through inherent in the macroeconometric model. Section 3 presents the estimated dynamic model of the pricing behaviour in the distribution sector. Section 4 reports empirical findings of exchange rate pass-through in the Norwegian economy based on simulations on the macroeconometric model. Section 5 concludes.

## 2 Channels of Exchange Rate Pass-Through

The theoretical set up of prices is generally based on imperfectly competitive markets characterised by differentiated products. However, the econometric specification of price equations also includes the possibility of price taking behaviour as a special case to accommodate the traditional assumption for small open economies. The Norwegian national accounts system, which is conceptually and statistically our main database, operates with three prices on each product depending on origin and market destination. A domestically produced good is either delivered to the domestic market (home goods) or abroad (exported goods) with potentially different prices. On the domestic market the price of a product produced in Norway may also be different from the import price of the "similar" product.

Figure 1 shows the main channels of exchange rate pass-through on prices and costs included in the macroeconometric model. First, there is a pricing-to-market link between exchange rates, marginal costs and import prices, which in turn affect consumer prices due to imported final consumer goods. Second, import prices affect export and domestic prices through the price setting behaviour of domestic firms in markets with imperfect competition. Third, import prices affect prices of intermediate goods and services used in the production of consumer goods. Domestic producers of consumer goods are also affected in their price setting by import competition. Fourth, import shares (denoted by *IS* in Fig. 1) and changes in these shares influence the degree of exchange rate pass-through on consumer prices and prices on material inputs. Finally, there is exchange rate pass-through on production costs through wage formation which is strongly affected by producer prices or profitability in addition to the unemployment rate and labour productivity.[3] These five exchange rate pass-through channels have partly immediate impacts as well as lagged effects on prices and costs. Consequently, there are considerable lags in the pass-through from changes in exchange rates to consumer prices. We further see from Fig. 1 that trade margins in the distribution sector, which inter alia are determined by world market prices, exchange rates, domestic prices and marginal costs, also affect consumer prices.

[3] Figure 1 does not show the effects of factor prices and real wages on unemployment and productivity, effects which are present in the macroeconometric model.

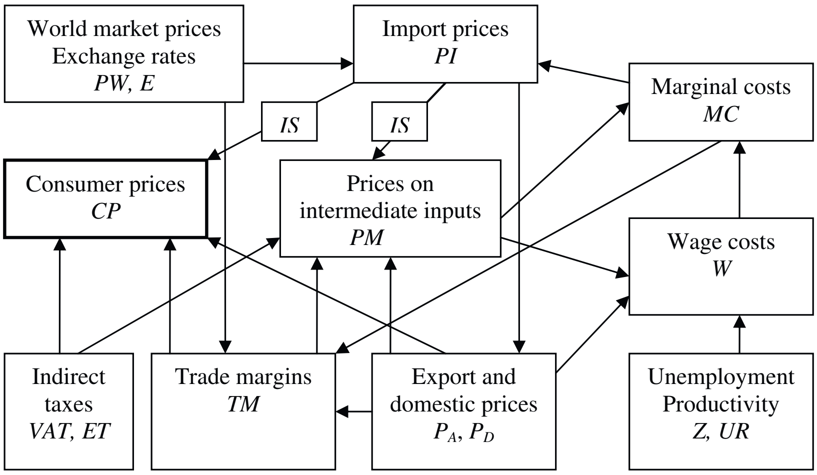


**Fig. 1** Main exchange rate pass-through channels

Below, we examine the main exchange rate pass-through channels closer by means of their representative equations included in the macroeconometric model.

## 2.1 Exchange Rates and Import Prices

World market prices in foreign currency times the exchange rate is often considered the main determinants of import prices in domestic currencies, at least in a small open economy. A number of empirical studies have found less than complete pass-through of exchange rate changes to prices of competitive imports, see for instance Goldberg and Knetter (1997), Campa and Goldberg (1999) and Campa and Mínguez (2006). When assuming imperfectly competitive markets symmetry considerations imply that foreign producers may potentially take their market position into account when setting export prices to Norway. The cointegrated VAR modelling of import prices of manufactures is therefore based on the pricing-to-market hypothesis advanced by Krugman (1987). In a simplified form we may write the import price equation of manufactures ($m$) as

$$PI_m = g(PW_m \cdot E, MC_m), \tag{1}$$

where $PW_m$ is the aggregate foreign export price (in foreign currency) of manufactures, $E$ is the import-weighted nominal exchange rate and $MC_m$ is marginal costs in manufacturing production representing the pricing-to-market hypothesis. The function $g(\cdot)$ is homogenous of degree one in prices and marginal costs enter with an elasticity of 0.35 in the long run, which is close to the corresponding estimate in Naug and Nymoen (1996). The pricing-to-market effects imply that the exchange rate is (partially), but not fully passed on to import prices in domestic currency in the long run due to imperfect competition. In the short run, the effect of the exchange rate is even smaller according to the estimated dynamics of (1). However, as domestic prices and costs also are influenced by changes in the exchange rate the reduced form of the macroeconometric model implies full pass-through in the very long run. For non-competitive imports where no similar domestic production exists, we assume the law of one price or the traditional small open economy assumption to be valid.

Benedictow and Boug (2012) discuss the cointegrated VAR modelling of import prices of manufactures in more detail.

## 2.2 Competing Prices, Mark-ups and Product Prices

With imperfect competition producers in each industry $j$ face regular downward sloping demand curves both domestically and abroad. Profit maximisation then leads to the formula stating that the destination specific price $l$ of product $i$ ($P_{li}$) equals a mark-up ($MU_{li}$) times marginal costs ($MC_j$)

$$P_{li} = MU_{li} \cdot MC_j, \ l = D, A, i = 1, \ldots, 44 \text{ and } j = 1, \ldots, 24, \tag{2}$$

where $P_D$ and $P_A$ are the prices on the domestic market ($D$) and the export market ($A$), respectively, $MU_{li}$ are the product specific mark-ups in the domestic and export markets and $MC_j$ is the industry specific marginal costs.[4] Typically in the new Keynesian literature, see e.g. Galí et al. (2001), producers are assumed to face isoelastic demand so the mark-up is a constant. This is the case with CES-utility functions in Dixit and Stiglitz (1977). When commodities within each industry are close substitutes, but poor substitutes for goods in other industries, so-called two-stage budgeting is valid. Moreover, if the number of goods in the industry is large, Dixit and Stiglitz (1977) show that individual producer prices have little impact on the aggregate industry price. Hence, we may assume that the individual producer ignores the effect of his price setting on the aggregate price. In general, the mark-up is not constant, but depends on all factors affecting demand for the particular commodity, see Eq. (32) in Dixit and Stiglitz (1977). In an open economy framework, the a priori assumption of all goods and services being perfect substitutes is clearly unreasonable. We therefore allow the mark-up to depend on relative prices in a way that also accommodates the possibility that producers may be price takers on world markets. Specifically, we let in accordance with Bowitz and Cappelen (2001) and Bjørnstad and Nymoen (2008) the mark-up be determined by

$$MU_{li} = m_{0li}(P_{li}/PI_i)^{m_{1li}}, \ l = D, A \text{ and } i = 1, \ldots, 44, \tag{3}$$

where $PI_i$ denotes the competing import price of product $i$ and $m_{0li} > 0$ and $m_{1li} \leq 0$ reflect conditions on the demand side of the product markets. With $m_{1li} < 0$ an increase in the competing price allows the producer to increase the mark-up over marginal costs. The more negative $m_{1li}$ becomes the more closely pricing decisions resemble price taking behaviour, which can be seen by inserting (3) into (2) and solving for $P_{li}$

$$P_{li} = m_{0li}^{1/(1-m_{1li})} \cdot PI_i^{-m_{1li}/(1-m_{1li})} \cdot MC_j^{1/(1-m_{1li})}, \ l = D, A, i = 1, \ldots, 44 \text{ and } j = 1, \ldots, 24. \tag{4}$$

We see from (4) that $P_{li}$ is homogenous of degree one in $PI_i$ and $MC_j$ and that $P_{li} = PI_i$ if $m_{1li}$ approaches infinity. The estimated price models in most cases imply that

[4] The input-output structure of the macroeconometric model contains more goods than industries so there is joint production in some industries. For instance, the industry Refineries produces petrol, gasoline and a composite good in addition. For these goods the mark up $MU_{li}$ can vary both between goods and destination (home or abroad). However, there is only one observable marginal cost which therefore is indexed by industry not commodity in (2).

mark-ups on domestic markets are independent of the import price while the specification in (3) receives much support by data. The latter result indicates that formation of export prices is closer to the small open economy assumption than prices on domestic markets. Accordingly, exchange rate pass-through to domestic prices is mainly related to production costs, while exchange rate pass-through to export prices is both related to marginal costs and mark-ups. Export prices of crude oil and natural gas in USD are assumed to be determined on the world markets and the exchange rate pass-through immediate with no domestic cost component, i.e., we assume oil companies are price takers.

## 2.3 Import Prices and Domestic Production Costs

For each industry, we specify a Cobb-Douglas production function with labour and materials as variable factors and capital as quasi fixed. Then, it follows that $MC_j$ and variable unit costs ($VUC_j$) are proportional. Thus, we replace $MC_j$ with $VUC_j$, which is defined as

$$VUC_j = \left(PM_j \cdot M_j + W_j \cdot LW_j\right)/X_j, \quad j = 1, \ldots, 24, \tag{5}$$

where $M_j$ is material inputs or intermediate inputs defined as a simple Leontief-aggregate of the 44 commodities included in the model, $PM_j$ is the dual price index of $M_j$, $W_j$ is wage costs per hour, $LW_j$ is hours worked and $X_j$ is gross production, see Hungnes (2011) for details on how $LW_j$ and $M_j$ are modelled in the macroeconometric model. The input price index $PM_j$ by industry is determined by summing over all goods

$$\begin{aligned} PM_j = \textstyle\sum_i \alpha_{ij}\Big[\left(1 + VAT_{ij}\right) \cdot \Big((1 - IS_i) \cdot P_{Di} + IS_i \cdot PI_i + \psi_{ij}ET_{ij}\Big)\Big] + \alpha_{Dj} \cdot TM, \\ i = 1, \ldots, 44 \text{ and } j = 1, \ldots, 24, \end{aligned} \tag{6}$$

where the $\alpha_{ij}$'s are input-output coefficients, $VAT_{ij}$ are value added taxes, $ET_{ij}$ are excise taxes, $IS_i$ are import shares and $TM$ denotes the national accounts price index for trade margins in the distribution sector. Because inputs of imported materials are important for total material costs (many large values of $a_{ij}$ $IS_i$ in (6)), changes in exchange rates − when passed through to prices in local currency − will affect domestic prices and hence $PM_j$ and $VUC_j$ significantly.

## 2.4 Import Prices and Import Penetration

The size of import shares determines the degree of import penetration from exchange rate changes both in consumption and production. For each commodity, assuming weak separability in demand between imported goods and home goods of the same variety (i.e., they are linked together using a CES aggregate), the import shares are functions of the relative domestic price to the import price

$$IS_i = l(P_{Di}/PI_i), \ i = 1, \ldots, 44. \tag{7}$$

For each consumption group $k$ we define a consumer price index ($CP_k$) similar to (6)

$$CP_k = \sum_i \alpha_{ik}[(1 + VAT_{ik}) \cdot ((1 - IS_i) \cdot P_{Di} + IS_i \cdot PI_i + \psi_{ik} ET_{ik})] + \alpha_{Dk} \cdot TM, \quad k = 1, \ldots, 14 \text{ and } i = 1, \ldots, 44. \tag{8}$$

We see that (8) links consumer prices to domestic prices, import prices, import shares, value added taxes, excise taxes and trade margins. The coefficient $\alpha_{Dk}$ represents the share of the trade margins (*TM*) in total consumer price for each consumption group in the base year. For some categories of consumption, say electricity and transportation, there is no trade margin at all so $\alpha_{Dk}$=0. For the CPI as a whole, the share of the trade margins is close to 0.3.[5] Thus, *TM* is of great importance for some consumer prices and thereby inflation. The direct (and partial) effect of an import price increase, say due to a depreciation of the exchange rate, on the CPI through the imported goods is estimated in the national accounts to 0.17 in 2006. As long as the trade margins are assumed constant in nominal terms, this pass-through effect takes place in the same quarter as import prices increase.

## 2.5 Product Prices, Production Costs and Wage Formation

The modelling of wages in the macroeconometric model is based on the symmetric Nash bargaining model following Nickell and Andrews (1983) and Hoel and Nymoen (1988). In manufacturing, wages are determined by profitability in that sector (which determines the "wage-corridor" in the long run version of the Scandinavian model of inflation, see Aukrust (1977)), while consumer prices as well as income taxes have no long run effects. Thus, the wage-curve relating real wages to unemployment thus includes the producer real wage not the consumer real wage. In private and government services wages are based on the alternative or "outside" wage depending on wages in other sectors and unemployment benefits, see Bowitz and Cappelen (2001) for a more detailed discussion. Consequently wage-wage spirals in non-manufacturing industries lead in the long run to profitability in manufacturing being the main nominal factor determining wages. The wage equation in manufacturing can be simplified as

$$W_m = Z_m \cdot PYF_m \cdot UR^{-\eta}, \tag{9}$$

where $Z_m$ is labour productivity in manufacturing defined as value added per hours worked $(X_m{-}M_m)/LW_m$, $PYF_m$ is the value added deflator at factor prices in manufacturing defined as $(P_{Dm} \cdot (X_m - A_m) + P_{Am} \cdot A_m - PM_m \cdot M_m)/X_m$ and *UR* is the unemployment rate determined as the difference between supply and demand of labour in the economy as a whole. Hence, exchange rate pass-through to wages mainly works through import prices to the extent that imported materials are important for total material costs and mark-ups and output prices in particular export prices

[5] The CPI is a weighted sum of all the $CP_k$. At an aggregate level these weights are determined using national accounts data. The weights are determined by a detailed consumer demand system in the macroeconometric model, cf. Boug et al. (2013).

that depend on competing prices in world markets. Further delays in the exchange rate pass-through process result from wage-wage spirals that pass wage impulses from manufacturing to more sheltered sector of the economy. Wages in sheltered sectors affect marginal costs and domestic prices, and thereby reflected in the CPI.

The modelling of wages based on the institutional set-up in Norway is one feature that distinguishes our macroeconometric model from mainstream econometric models in most countries where centralised wage bargaining has become less important. However, disregarding wage formation, the presentation of the main pass through channels in this section is quite general and not very specific to the Norwegian economy. The specific characteristics of the distribution sector in Norway and the importance of the trade margins for the overall pass-through of exchange rate changes to consumer prices are discussed in the next section.

## 3 The Distribution Sector

Our modelling of the pricing behaviour in the distribution sector differs somewhat from the general price Eq. (4). In the Norwegian national accounts, the domestic price in the distribution sector comprises the *trade margins* on distribution services from supplier to user. Thus, the national accounts make a clear distinction between services delivered and products traded, and it is the former that constitutes the production activity in the distribution sector. The consumer price thus consists of two components: the price on the services delivered (i.e., the trade margin) and the purchasing price (or costs) of the good sold (exclusive of the trade margin).

### 3.1 Theory

In line with (4), we assume trade margins to be proportional to marginal costs in the distribution sector. Here, we may think of two cases depending on the substance of the marginal costs in each particular wholesale and retail trade firm. First, some firms may set their trade margins as a *constant amount of money per unit traded commodity*, i.e., independent on the purchasing price per unit (exclusive of the trade margin). In this case, the marginal costs are only related to costs of production. We approximate these costs in accordance with (5) letting $VUC^d$ denote variable unit costs in the distribution sector. Second, some firms may set their trade margins as a *fixed percentage mark-up on the purchasing price* (exclusive of the trade margin). In this case, the marginal costs also depend on the purchasing price on goods sold, and not only on costs in production. We approximate these costs by constructing a price index of purchasing prices ($PP$) in the distribution sector

$$PP = \sum\nolimits_k \delta_k \Big[ \sum\nolimits_i \beta_{ik}(1 - IS_i) \cdot P_{Di} + \beta_{ik} \cdot IS_i \cdot PI_i \Big], \tag{10}$$

where $\delta_k$ is the volume share of demand category $k$ out of total trade, $\beta_{ik}$ is the input-output coefficient for total delivery of commodity $i$ to demand category $k$

in the base year and $IS_i$, $P_{Di}$ and $PI_i$ are as defined in the previous section.[6] This price index thus weighs together domestic and import prices on all commodities traded by the wholesale and retail trade sector. Both the import shares (through $IS_i$) and each demand category weight (through $\delta_k$) are time series. For simplicity, we assume constant $\beta_{ik}$ coefficients and ignore any variations in these coefficients relative to base year values.

The distribution sector in Norway has undergone significant structural changes over the past decades. Shopping centres have replaced a large part of small shops run by self-employed. This has resulted in lower trade margins over time. We capture this underlying structural change by including a ratio of hours-worked by self-employed ($H$) and total production or services delivered ($X^d$) in the price equation for the trade margins. We specify aggregate trade margins as

$$TM = PP^{\gamma} \cdot VUC^{d(1-\gamma)} \cdot \left(H/X^d\right)^{\phi}. \tag{11}$$

The coefficients $\gamma$, $(1-\gamma)$ and $\phi$ measure the degree of pass-through of changes in purchasing prices (through domestic and/or import prices), variable unit costs and the mentioned ratio of the two "trend"-variables, respectively. We interpret (11) as a long run relationship between $TM$, $PP$, $VUC^d$ and $H/X^d$ and will serve as the starting point for the cointegration analysis below. In the dynamic modelling, we introduce changes in the nominal exchange rate ($E$) as an additional explanatory variable and suggest that trade margins act as cushions to exchange rate fluctuations in the short and medium term, thereby mitigating the degree of exchange rate pass-through. We also open up for the nominal interest rates ($R$) to play a potential role in the short run dynamics of trade margins to account for *financial* costs associated with stock of goods. Details on data definitions and sources can be found in the Appendix.

## 3.2 Data

The econometric modelling of trade margins is based on quarterly, seasonally unadjusted data that span the period $1970Q1-2010Q3$, of which data from the period $1970Q1-1998Q4$ and $1999Q1-2010Q3$ are used for estimation and *out-of-sample* forecasting, respectively. The reasons for ending the estimation period in $1998Q4$ are as follows: Taylor (2000) among others argues that the degree of exchange rate pass-through to domestic prices depends on the monetary policy regime in force. During the 1970s Norway joined the European exchange rate agreement, the so-called "snake". However, the Norwegian currency (the krone) experienced significant revaluations and devaluations during the first decade of our sample period. When Norway left the "snake" at the end of the 1970s and established a currency basket, the krone still showed relatively high variability during the 1980s. Following a 12 %

[6] The main categories in (10) are Food, Beverages, Tobacco, Fuels for heating purposes, Purchase of and expenses on own transport vehicles, Purchase of other durable goods, Clothes and footwear, Health services and Investments in machines and transport vehicles. The $\beta_{ik}$ coefficients in (10) are scaled such that for each demand category $k$ they sum up to unity for those commodities included in $k$. Likewise, by definition, the $\delta_k$ coefficients also sum up to unity. Prices included in $PP$ are producer prices (exclusive of trade margins) based on sales to the domestic market only. The national accounts define a similar price index for sales on foreign markets and the export price deflator for the trade sector is equal to this index.

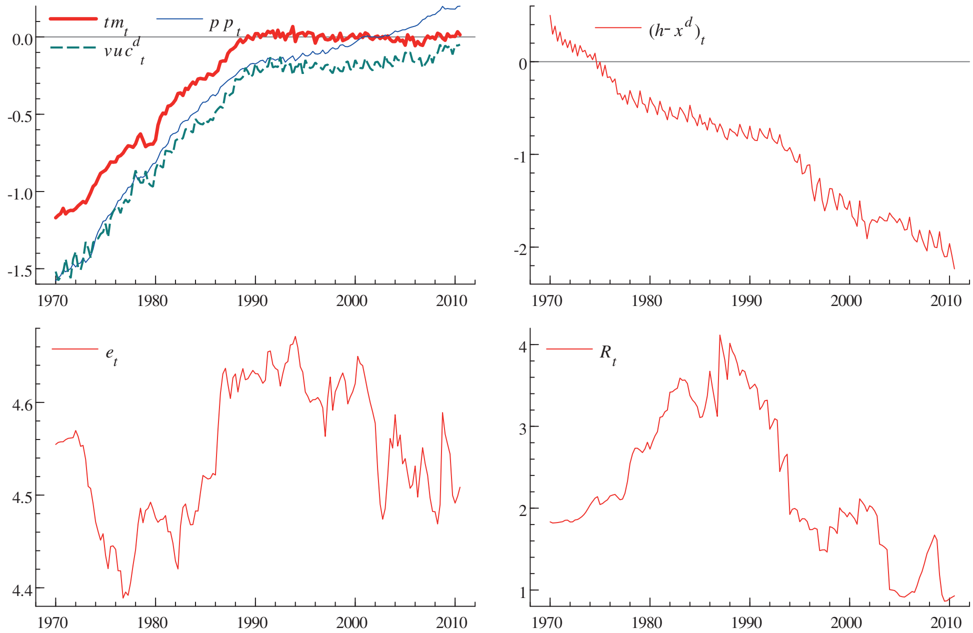


**Fig. 2** Time series for $tm_t$, $pp_t$, $vuc_t^d$, $(h-x^d)_t$, $e_t$ and $R_t$

devaluation of the krone in May 1986, a flexible interest rate policy was introduced with the explicit goal of supporting a policy of fixed exchange rates, and as of October 1990 against the ECU. After the turmoil following the speculative attacks against the krone by the end of 1992, Norway changed to a managed "floating" exchange rate regime whereby the exchange rate was allowed to float within defined

**Table 1** Augmented Dickey-Fuller tests

| Variable | *t*-ADF | 5 % critical value | Lags |
|---|---|---|---|
| $tm_t$ | −1.207 | −3.44 | 1 |
| $pp_t$ | −1.856 | −3.44 | 4 |
| $vuc_t^d$ | −1.384 | −3.44 | 5 |
| $(h-x^d)_t$ | −1.997 | −3.44 | 5 |
| $e_t$ | −1.613 | −3.44 | 4 |
| $R_t$ | −1.294 | −3.44 | 4 |
| $\Delta tm_t$ | −4.541 | −2.88 | 3 |
| $\Delta pp_t$ | −3.265 | −2.88 | 3 |
| $\Delta vuc_t^d$ | −4.035 | −2.88 | 4 |
| $\Delta(h-x^d)_t$ | −5.390 | −2.88 | 4 |
| $\Delta e_t$ | −7.100 | −2.88 | 4 |
| $\Delta R_t$ | −4.355 | −2.88 | 3 |

Sample period: 1971*Q*3 – 2010*Q*3 for variables in levels and 1971*Q*4 – 2010*Q*3 for variables in first differences. The regressions include a constant, a trend and seasonals and a constant and seasonals in the cases of ADF-test on variables in levels and variables in first differences, respectively. Initially, the regressions include five lags. Akaike´s information criterion is used in order to choose the optimal lag order.

upper and lower bands. However, Norway *formally* changed from exchange rate targeting in various forms to freely floating exchange rates following the introduction of inflation targeting in late March 2001. Several Norwegian economists argue that the regime change in fact occurred early in 1999. In any case, the monetary policy change took place after 1998$Q$4; the last observation used in our estimation. The *major* regime shift in monetary policy could in principle have caused the degree of exchange rate pass-through to alter in accordance with the Lucas critique. We shed light on this hypothesis by conducting an *out-of-sample* forecasting exercise with data covering the period 1999$Q$1−2010$Q$3. By ending the forecasting period in 2010$Q$3, we are also able to test the hypothesis that the distributers pricing behaviour has changed significantly during the financial crises in 2008 and 2009 with highly volatile exchange rates.

Figure 2 displays time series of trade margins ($tm_t$), purchasing prices ($pp_t$) and variable unit costs ($vuc_t^d$) together with the ratio $(h-x^d)_t$, the nominal exchange rate ($e_t$) and the nominal interest rate ($R_t$) over the entire sample period 1970$Q$1−2010$Q$3.[7] We observe that trade margins, purchasing prices and variable unit costs exhibit a clear upward trend. The underlying development in the market conditions described above is also evident in the data series as $(h-x^d)_t$ shows a clear downward trend, a trend which presumably has contributed to the observed price dampening since the late 1980s. The exchange rate series and the interest rate series, on the other hand, show some evidence of mean reversion (although slow) property. A clear reduction in trade margins through 1979 coincides with massive governmental price regulations during the second half of the 1970s, cf. Bowitz and Cappelen (2001).

Table 1 reports standard Augmented Dickey-Fuller tests. It is evident that $tm_t$, $pp_t$, $vuc_t^d$ and $(h-x^d)_t$ as well as $e_t$ and $R_t$ in our sample period are integrated of order one. Accordingly, all six variables should in principle be modelled as non-stationary $I(1)$ variables in the cointegration analysis below. However, we choose to carry out the cointegration analysis with $e_t$ and $R_t$ being non-modelled stationary variables in order to get a manageable and still statistically well specified underlying vector autoregressive (VAR) model. There is also supporting evidence for this approach in that the corresponding equilibrium correction term in the final dynamic model for trade margins seems to be stationary.

### 3.3 Cointegration Analysis

Because multiple cointegrating vectors among the variables in (11) may exist, we employ the Johansen (1995, p. 155) trace test for cointegration rank determination, both with and without small sample adjustments. In accordance with the Augmented Dickey Fuller tests, we may fit a six-dimensional VAR to the data and then test formally, after having determined the cointegration rank, the exogeneity status or otherwise of the exchange rate series and the interest rate series. However, as pointed out by Johansen (1995, p. 213), the power of the trace test decreases as the dimension of the underlying VAR increases. For this reason, and the fact that the number of parameters to be estimated in a six-dimensional VAR is very large relative to the number of observations in the available data set, it would be useful to impose weak exogeneity on both the

[7] All variables except the interest rates are in logarithms and denoted by lower case letters in what follows.

exchange rate series and the interest rate series. Hence, we rely on Rahbek and Mosconi (1999) for the cointegration rank inferences with $e_t$ and $R_t$ being the supposedly non-modelled stationary regressors in our case. We then perform weak exogeneity tests on $e_t$ and $R_t$ by means of standard $\chi^2$ inference after the value of the rank and the estimates of the cointegrating vector (s) are determined. Our starting point of the cointegration analysis and the tests that follow is thus an equilibrium correction representation of a four-dimensional VAR (henceforth CVAR) of order $k$ having the form

$$\Delta x_t = \sum_{i=1}^{k-1} \theta_i \Delta x_{t-i} + \sum_{i=0}^{k} \phi_i y_{t-i} + \sum_{i=0}^{k} \phi_i z_{t-i} + \pi x_{t-1} + \mu + \psi S_t + \eta D_t + \varepsilon_t, \quad (12)$$

where $x_t = \left(tm_t, pp_t, vuc_t^d, \left(h - x^d\right)_t\right)'$ is a vector of the modelled variables, $y_t$=$e_t$ and $z_t$=$R_t$ are the stationary explanatory variables, $\mu$ is a vector of constants, $S_t$ is a vector of centered seasonals (labelled $S_{1t}$, $S_{2t}$ and $S_{3t}$), $D_t$ contains a price stop dummy (labelled $PSTOP_t$) with a value of unity in price regulation periods and minus unity in catch-up periods during the second half of the 1970s and $\varepsilon_t$~IN(0,Σ). Assuming $x_t$ to be $I(1)$, presence of cointegration implies $0<r<4$, where $r$ denotes the rank or the number of cointegrating vectors of $\pi$. The null hypothesis of $r$ cointegrating vectors may be formulated as $H_0$: $\pi = \alpha\beta'$, where $\alpha$ and $\beta$ are 4x$r$ matrices, $\beta' x_t$ comprises $r$ cointegrating $I(0)$ linear combinations and $\alpha$ contains the adjustment coefficients. We let the constants, the seasonals, the price stop dummy and the stationary regressors enter unrestrictedly in (12). However, as pointed out by Rahbek and Mosconi (1999), the asymptotic distribution of the trace test in this model depends on nuisance parameters due to the presence of stationary regressors. Hence, the approach suggested by Rahbek and Mosconi (1999) is to analyse the extended model given by

$$\Delta x_t = \sum_{i=1}^{k-1} \theta_i \Delta x_{t-i} + \sum_{i=0}^{k} \phi_i y_{t-i} + \sum_{i=0}^{k} \varphi_i z_{t-i} + \pi^* \begin{pmatrix} x_{t-1} \\ \sum_{i=1}^{t} y_i \\ \sum_{i=1}^{t} z_i \\ t \end{pmatrix} + \mu + \psi S_t + \eta D_t + \varepsilon_t, \quad (13)$$

where $\pi^* = \alpha\beta^{*\prime}$. After the rank is determined using critical values tabulated in Harbo et al. (1998), we test the linear restrictions that there are no accumulated level of the exchange rate series and the interest rate series and no linear trend in the cointegrating relations by considering the hypothesis $\beta^* = (\beta, 0)'$ with standard $\chi^2$ inference. The likelihood ratio tests for this hypothesis may, in line with Rahbek and Mosconi (1999), be regarded as misspecification tests of the model in (12). Strictly speaking, the cointegration rank needs not be determined from (12) once it has been determined from (13). Nevertheless, we compare the cointegration rank inference from (13) with the cointegration rank inference from (12) as a robustness check. As guidance for choosing the optimal lag order ($k$) we rely on both Akaike's information criterion (AIC) and various diagnostic tests. According to AIC we should include five lags ($k$=5) in both (12) and (13), albeit AIC is a borderline case with respect to $k$=4 and $k$=5 in (12). However, although the equation for $\Delta tm_t$ has well-behaved residuals in both models, this was not the case for the other three equations. To secure valid statistical inference, we include the same set of impulse dummies in (12) and (13) to control for outliers and instabilities in the equations for $\Delta pp_t$, $\Delta vuc_t^d$ and $\Delta(h-x^d)_t$, see the Appendix for details. Diagnostic tests for the preferred VARs (with $k$=5) including the set of impulse dummies reveal no serious problems with misspecification and recursively estimated one step residuals and sequences of break-point Chow tests indicate that both systems are reasonably stable over the sample. Table 2 reports trace test statistics using (12) and (13).

Based on (13) and the trace test with a small sample adjustment ($\lambda^{a}_{trace}$), the null hypothesis of no cointegration is rejected at the 5 % significance level, whereas the hypothesis of at most one cointegrating vector among the variables involved is not. Testing the hypothesis $\beta^* = (\beta, 0)'$, assuming $r$=1, gives $\chi^2(1)$=4.324 ($p$-value= 0.038), $\chi^2(1)$=0.081 ($p$-value=0.776) and $\chi^2(1)$=4.858 ($p$-value=0.028) for the accumulated level of the exchange rate series, the accumulated level of the interest rate series and the linear trend, respectively. Accordingly, the hypothesis of no accumulated level of the exchange rate series and no linear trend in the cointegrating relationship, although not accepted at the 5 % significance level, is not strongly rejected. In this sense, we argue that (12) to a large degree passes the misspecification tests. Based on (12), we also notice that $\lambda^{a}_{trace}$ supports the hypothesis of only one cointegrating vector between $tm_t$, $pp_t$, $vuc_t^d$ and $(h-x^d)_t$, the same conclusion about the value of the rank using (13). We therefore apply (12) in the successive cointegration analysis assuming $r$ =1. The estimate of the *unrestricted* cointegrating vector (normalised on $tm_t$) is given by (standard errors in parenthesis)

$$\underset{}{tm_t} = \widehat{\alpha}_0 + \underset{(0.175)}{0.384}pp_t + \underset{(0.176)}{0.529}vuc_t^d + \underset{(0.057)}{0.055}(h-x^d)_t, \tag{14}$$

which is interpretable as an equation for trade margins as the estimated coefficients for purchasing prices, marginal costs and the ratio between self-employed hours worked and production are economically reasonable with expected signs. Besides, weak exogeneity tests give $\chi^2(1)$=5.504 ($p$-value=0.019), $\chi^2(1)$=2.659 ($p$-value= 0.103), $\chi^2(1)$=4.949 ($p$-value=0.026) and $\chi^2(1)$=0.093 ($p$-value=0.761) for $tm_t$, $pp_t$,

**Table 2** Tests for cointegration rank

| Hypothesis | Model (13) | | Model (12) | |
|---|---|---|---|---|
| | $\lambda_{trace}$ | $\lambda^{a}_{trace}$ | $\lambda_{trace}$ | $\lambda^{a}_{trace}$ |
| $r$ =0 | 105.39 (80.9) | 86.40 (80.9) | 60.05 [0.002] | 49.23 [0.035] |
| $r \le 1$ | 61.81 (56.3) | 50.68 (56.3) | 32.34 [0.024] | 26.51 [0.117] |
| $r \le 2$ | 32.76 (35.5) | 26.85 (35.5) | 10.21 [0.270] | 8.37 [0.434] |
| $r \le 3$ | 11.16 (17.9) | 9.15 (17.9) | 0.38 [0.537] | 0.31 [0.576] |

Sample period: 1971$Q$2 – 1998$Q$4. The underlying VARs are of order 5. $r$ denotes the cointegration rank. The $\lambda_{trace}$ and $\lambda^{a}_{trace}$ are the trace test statistics without and with degrees-of-freedom-adjustments, respectively. The critical values in parenthesis, which correspond to the 5 % significance level, are from Table 2 in Harbo et al. (1998). The $p$-values in brackets, which are reported in OxMetrics, are based on the approximations to the asymptotic distributions derived by Doornik (1998). The critical values produced by OxMetrics are only indicative as the inclusion of the two conditioning variables in (12) affects the asymptotic distribution of the trace test, see Rahbek and Mosconi (1999).

$vuc_t^d$ and $(h - x^d)_t$, respectively, which imply that the cointegrating vector enters the $\Delta tm_t$-equation (albeit also the $\Delta vuc_t^d -$ equation , but not so significantly). The sum of the estimated coefficients of $\gamma$ and $(1-\gamma)$ in (14) is not far from unity. To complete the cointegration analysis, we thus tested for, and could not reject, homogeneity between $tm_t$, $pp_t$ and $vuc_t^d$ . Imposing the homogeneity restriction and weak exogeneity of $(h - x^d)$ gives $\chi^2(2)=3.557$ ($p$-value=0.169) and the following *restricted* estimate of the cointegrating vector (standard errors in parenthesis):

$$\begin{array}{c} tm_t = \widehat{\alpha}_0 + 0.365pp_t + 0.635vuc_t^d + 0.123(h - x^d)_t \\ \qquad\qquad\qquad\qquad (0.186) \qquad (0.019) \end{array} \tag{15}$$

Recursively estimated parameters of $vuc_t^d$ and $(h - x^d)_t$ are reasonably constant and a sequence of $\chi^2(2)$ test statistics confirms the validity of the homogeneity restriction and the weak exogeneity status of the ratio between self-employed hours worked and production for any sample ending between 1985 and 1998. Also, the restricted cointegrating vector is virtually unchanged from the unrestricted one. Before exploiting results from the cointegration analysis within a dynamic equilibrium correction model of trade margins we shall test formally that the exchange rate series and the interest rate series are weakly exogenous and not just an assumption imposed at the outset. We carry out likelihood ratio tests based on a CVAR system using deviations from (15) as an equilibrium correction mechanism (*EqCM*)

$$\Delta x_t = \sum_{i=1}^{4} \theta_i \Delta x_{t-i} + \delta EqCM_{t-1} + \mu + \psi S_t + \eta D_t + \vartheta DUM_t + \varepsilon_t, \tag{16}$$

where $x_t$ now, as opposed to (12), also contains $e_t$ and $R_t$ and $DUM_t$ includes the set of impulse dummies described above. The system in (16) is estimated by FIML. Testing the hypothesis of weak exogeneity of $e_t$ and $R_t$ involves zero restrictions on the associated adjustment coefficients inherent in $\delta$. The likelihood ratio test gives $\chi^2(2)=1.751$ ($p$-value=0.417), which means that the hypothesis of weak exogeneity of $e_t$ and $R_t$ with respect to the long run parameters in (12) is not rejected by data.

## 3.4 Short run Dynamics

We now focus on (*i*) the dynamic adjustment of trade margins to changes in purchasing prices, variable unit costs and the ratio between self-employed hours worked and production and (*ii*) the role of the exchange rate and the interest rates as separate explanatory variables in the short and medium term. For this purpose, we derive a dynamic equilibrium correction model for trade margins based on a general-to-specific modelling strategy. We recall that the weak exogeneity tests imply that both $tm_t$ and $vuc_t^d$ are error correcting (albeit the latter less significantly), whereas $pp_t$ and $(h - x^d)_t$ are not. Consistent with the cointegration analysis, we therefore start with a general system

$$
\begin{aligned}
\Delta tm_t = {} & \kappa_{tm} + \sum_{i=1}^{4} \varpi_{1i}\Delta tm_{t-i} + \sum_{i=0}^{4} \varpi_{2i}\Delta pp_{t-i} + \sum_{i=0}^{4} \varpi_{3i}\Delta vuc^d_{t-i} + \sum_{i=0}^{4} \varpi_{4i}\Delta(h-x^d)_{t-i} \\
& + \sum_{i=0}^{4} \varpi_{5i}\Delta e_{t-i} + \sum_{i=0}^{4} \varpi_{6i}\Delta R_{t-i} + \lambda_{tm} EqCM_{t-1} + \psi_{tm} S_t + \eta_{tm} D_t + \vartheta_{tm} DUM_t + \varepsilon_{tm,t} \\
\Delta vuc^d_t = {} & \kappa_{vuc^d} + \sum_{i=0}^{4} \omega_{1i}\Delta tm_{t-i} + \sum_{i=0}^{4} \omega_{2i}\Delta pp_{t-i} + \sum_{i=1}^{4} \omega_{3i}\Delta vuc^d_{t-i} + \sum_{i=0}^{4} \omega_{4i}\Delta(h-x^d)_{t-i} \\
& + \sum_{i=0}^{4} \omega_{5i}\Delta e_{t-i} + \sum_{i=0}^{4} \omega_{6i}\Delta R_{t-i} + \lambda_{vuc^d} EqCM_{t-1} + \psi_{vuc^d} S_t + \eta_{vuc^d} D_t + \vartheta_{vuc^d} DUM_t + \varepsilon_{vuc^d,t}
\end{aligned}
\tag{17}
$$

The system in (17) is estimated by FIML. To exactly identify the two equations in the system, the impulse dummy $D74Q1$ and the price stop dummy $PSTOP_t$ are excluded from the equation for trade margins and variable unit costs, respectively. We find a parsimonious model by stepwise elimination of insignificant variables in the system. It turned out that this general-to-specific system analysis produces a dynamic model for trade margins which is close to the economic content and statistical significance of a dynamic model derived from a general-to-specific single equation analysis of (17). For this reason, we focus in the following on the dynamic model for trade margins derived from the single equation analysis. We thereby follow the argument by Boswijk and Urbain (1997), that one may apply single equation analysis with the long run relationship(s) estimated and deduced from a VAR in cases where the conditioning variables are error correcting, but weakly exogenous for the short run parameters. Also, an equation for variable unit costs, through the definition in (5) in Section 2 and CVARs for wages and labour demand, is already embedded in the macroeconometric model.

The specific dynamic model derived from the single equation analysis is presented in (18). The conditioning on $\Delta pp_t$ and $\Delta(h-x^d)_t$ may pose caveats because they need not be weakly exogenous for the short run parameters. Adding the predicted counterparts to $\Delta pp_t$ and $\Delta(h-x^d)_t$ from the VAR, both individually and jointly, yield $p$-values of 0.307 and 0.545 and $\chi^2(2)$=1.386 ($p$-value=0.500), and may be taken as evidence that $\Delta pp_t$ and $\Delta(h-x^d)_t$ indeed are weakly exogenous for the short run parameters in (18). Consequently, the parameters in (18) are consistently estimated by OLS.[8,9]

$$
\begin{aligned}
\Delta tm_t = const. & - \underset{(0.066)}{0.188}\Delta_2 tm_{t-1} + \underset{(0.113)}{0.607}\Delta pp_t + \underset{(0.071)}{0.258}\Delta_2 pp_{t-2} + \underset{(0.041)}{0.092}\Delta vuc^d_{t-3} \\
& + \underset{(0.026)}{0.046}\Delta(h-x^d)_t - \underset{(0.107)}{0.252}\Delta e_t + \underset{(0.010)}{0.022}\Delta R_{t-2} + \underset{(0.011)}{0.033}\Delta R_{t-4} \\
& - \underset{(0.054)}{0.245}[tm - 0.365pp - 0.635vuc^d - 0.123(h-x^d)]_{t-1} \\
& - \underset{(0.006)}{0.019}S_{1t} - \underset{(0.004)}{0.022}PSTOP_t
\end{aligned}
\tag{18}
$$

$OLS$, $T$=111(1971$Q$2–1998$Q$4), $R^2$=0.601, $\sigma$ =1.73 %

| | |
|---|---|
| $AR_{1-5}$ | $F(5, 94)$=0.935 [0.462] |
| $ARCH_{1-4}$ | $F(4, 103)$=1.789 [0.137] |
| $NORM$ | $\chi^2(2)$=1.036 [0.596] |
| $HET$ | $F(21, 89)$=0.679 [0.843] |
| $RESET$ | $F(2, 97)$=1.676 [0.193] |

[8] Square brackets [..] and parenthesis (..) contain $p$-values and standard errors, respectively.

[9] Based on statistical inference we have simplified the dynamics in (18) such that $\Delta_2 tm_{t-1} = \Delta tm_{t-1} + \Delta tm_{t-2} = tm_{t-1} - tm_{t-3}$ and $\Delta_2 pp_{t-2} = \Delta pp_{t-2} + \Delta pp_{t-3} = pp_{t-2} - pp_{t-4}$ .

Below (18) we report several test statistics.[10] None of the diagnostics are significant at the 1 % significance level. The economic variables entering (18) are all significant and the *EqCM* appears in the model with a *t*-value of −4.54, adding force to the results obtained from the cointegration analysis. We notice that (18) implies rejection of *dynamic* homogeneity and that the mark-up rates decrease with higher inflation, a finding which is in line with previous studies based on European and American data, see e.g. Bénabou (1992), Bowitz and Cappelen (2001) and Banerjee and Russell (2004). Empirical evidence of constancy of (18) is supported by one-step residuals, one-step Chow tests, break-point Chow tests, forecast Chow tests and recursively estimated coefficients, which do not reject constancy between 1979 and 1998. Besides, the impulse dummies used to account for outliers in the $\Delta pp_t$, $\Delta vuc_t^d$ and $\Delta(h - x^d)_t$ equations in the VAR are all *in*significant when added to (18), both individually and jointly. These findings are evidence against relevance of the Lucas critique in our context, see e.g. Favero and Hendry (1992), and we claim that the degree of exchange rate pass-through to trade margins has remained fairly constant throughout the estimation period.

The estimated impact response of purchasing prices (0.6) is somewhat larger than its long run counterpart. Apparently, the trade margins overshoot with respect to changes in domestic as well as import prices on traded commodities in the short run. The pass-through from variable unit costs to the trade margins is, however, delayed and incomplete in the short run. Turning to the exchange rate itself, the estimated impact elasticity of −0.25 shows that trade margins are significantly affected by exchange rate fluctuations in the short run. These *direct* effects of the exchange rate work in the opposite direction compared to the *indirect* effects of purchasing prices and variable unit costs. If the exchange rate appreciates by 10 %, say, then trade margins increase immediately by 2.5 %, but at the same time decrease with changes in purchasing prices (and variable unit costs with some delay) that are caused by decreased import prices of tradable goods and material inputs. With periods of large fluctuations in the exchange rate, firms may find it difficult to perceive whether the changes are transitory or permanent. Hence, it is likely that firms are reluctant to change their prices in response to exchange rate fluctuations for reasons such as menu costs and stock of products with different purchasing price than the current price. Under such circumstances, firms may increase their sales considerably by leaving the trade margins unchanged in periods of exchange rate appreciation. Some firms (importers) may also secure themselves against exchange rate fluctuations, either through financial instruments or price agreements, thereby contributing to modest exchange rate pass-through to trade margins and further to consumer prices. The short run dynamics of the trade margins are also in a quantitatively important way influenced by past changes in interest rates. The adjustment towards steady state is

[10] The reported statistics are as follows: $T$, $R^2$ and $\sigma$ are the number of observations, the squared multiple correlation coefficient (adjusted) and the residual standard errors, respectively. $AR_{1-5}$ is Harvey's (1981) test for up to 5th order residual autocorrelation, $ARCH_{1-4}$ is the Engle (1982) test for up to 4th order autoregressive conditional heteroskedasticity in the residuals, *NORM* is the normality test described in Doornik and Hansen (1994), *HET* is a test for residual heteroskedasticity due to White (1980) and *RESET* tests for functional form misspecification [cf. Ramsey (1969)]. $F(\cdot)$ and $\chi^2(\cdot)$ represent the null distributions of $F$ and $\chi^2$, with degrees of freedom shown in parenthesis. The null hypothesis underlying the various diagnostic tests is that the residuals are white noise.

rather slow as reflected by the small estimated magnitude of the loading parameter (−0.24). Summing up, the estimated dynamics of (18) imply that trade margins act as cushions to exchange rate fluctuations during the first year due to the short run effects of the exchange rate. In later years there is a gradual and smooth increase in trade margins due to the fact that variable unit costs also increase gradually, in particular wages according to (9) in Section 2.

### 3.5 Out-of-Sample Forecasting

We now study the *out-of-sample* forecasting performance of (18) to shed light on its robustness with respect to the financial crises and the monetary policy regime shift. If the distributors pricing behaviour has changed significantly following the economic events in the last decade, we should expect instabilities in (18) as, for example, indicated by poor *out-of-sample* forecasting ability. To assess the forecasting performance of (18), we employ forty seven quarters (1999$Q$1−2010$Q$3) of *out-of-sample* observations, including the periods of both the financial crises and the formal change in monetary policy regime. Figure 3 depicts actual values of $\Delta tm_t$ together with one-step ahead forecasts, adding bands of 95 % confidence intervals to each forecast in the forecasting period.[11] As many as forty two actual values of $\Delta tm_t$ stay clearly within their corresponding confidence intervals over the forecasting period. Interestingly, none of the five forecasting failures coincide with the point in time of the formal change in monetary policy. Thus, the *out-of-sample* forecasting ability of (18) is reasonably good with respect to the major shift in monetary policy from exchange rate targeting to inflation targeting in late March 2001. The regime robustness is further evidence that the Lucas critique lacks force in our case.

That said, we observe that the actual values of $\Delta tm_t$ are outside their corresponding confidence intervals in 2008$Q$4 and 2009$Q$1, two quarters in which the exchange rate was particularly volatile during the financial crisis, cf. Figure 2. However, the forecasting ability of (18) is thereafter reasonably good in a period with repercussions of the financial crisis still present. Hence, we interpret the two mentioned forecasting failures to be “outliers” rather than evidence of structural breaks in the distributers pricing behaviour caused by the financial crises. Noticeably, (18) is virtually unchanged with respect to both economic content and diagnostics when estimated with a sample period ending in 2010$Q$3 rather than in 1998$Q$4. We conclude that the dynamic model of trade margins in the distribution sector shows a high degree of historical constancy and contains some interesting properties with respect to exchange rate pass-through. In the next section, we show the degree of exchange rate pass-through to the overall consumer price index when the model of trade margins is analysed within the macroeconometric model of the Norwegian economy.

[11] The forecast for period $s$ is $\widehat{y}_s = x'_s\widehat{\beta}_t$ , where $x_s$ is the observed value of $x$ for period $s$, $\widehat{\beta}_t$ is estimated from the first $t$ observations of data and $s>t$. In our case $s$ spans the period 1999$Q$1−2010$Q$3, while $t$ covers the period 1971$Q$2−1998$Q$4.

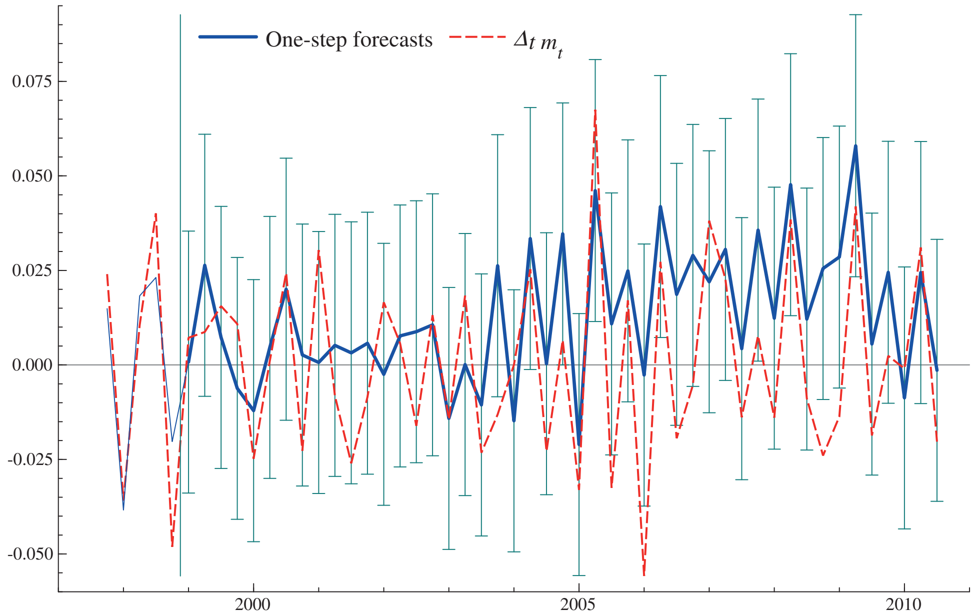


**Fig. 3** Actual values of $\Delta tm_t$ and one-step ahead forecasts with 95 % bands

## 4 Simulation Results

The purpose of this paper is to study the pass-through of exchange rate changes to consumer prices. We therefore assume an exogenous exchange rate in our simulation study although the exchange rate is clearly not an exogenous variable in an economy where monetary policy is based on inflation targeting. In the simulations, the import-weighted exchange rate is permanently increased (depreciated) by 10 % compared to a reference path, which roughly follows the historic development of the Norwegian economy until 2010.[12] The real after tax interest rate is set equal to the value in the reference path by adjusting the nominal interest rate according to changes in CPI-inflation. The degree of exchange rate pass-through to the Norwegian economy is studied in a full model simulation, thereby taking account of all modelled channels through which the exchange rate affects domestic prices.

We assume fiscal policy to be unchanged in real terms compared to the reference path despite the fact that Norway introduced a fiscal policy rule in 2001. The fiscal policy rule states that over time the government can include 4 % (equal to the expected real rate of return) of the value of the Petroleum Fund in the net balance for higher expenditures or lower taxes.[13] In our simulations the exchange rate shock

[12] The simulations begin in the first quarter of 2000, but the simulated effects reported below are for practical purposes independent of the actual time period. In line with actual "policy rules" in Norway, all excise taxes in equation (8) adjust in accordance with the consumer price index to avoid any nominal price inertia (due to these taxes) to affect the simulation results.

[13] When the rule was formally introduced actual fiscal policies were already in line with the rule. Hence, the formal introduction of the rule did not bring about a change in policy in that year. The same can be said for monetary policy, which formally changed at the same time from exchange rate targeting to inflation targeting, see the discussion in Section 3. Inflation was close to the target chosen in 2001 and had been so for nearly a decade. The Petroleum Fund was set up in 1991, but revenues transferred to the fund from 1996. The magnitude of the fund was 1.2 times GDP in 2011.

changes the value of the fund by the same rate in domestic currency and one may ask whether fiscal policy should become more expansionary as a consequence. We think not. First, government expenditures also increase in nominal terms since we keep real expenditures fixed. Second, a depreciation of the currency increases GDP in the short run. Fiscal policy should accordingly be slightly less expansionary according to the rule which states clearly that discretionary policies should be pursued in addition to relying on automatic stabilisers. Third, the guidelines for the fiscal policy rule state that policies should be smoothed when the fund is affected by nominal shocks. In the longer term prices and costs will be higher. Therefore, the assumption of unchanged fiscal policy in real terms is not unrealistic.

## 4.1 Effects on Prices

Figure 4 shows simulation results for aggregate price indices. The exchange rate pass-through to import prices is quite rapid in the short run, although still incomplete in the medium to longer run (10 years). This finding is in accordance with the pricing-to-market hypothesis. The exchange rate pass-through to export prices is also rather fast, but nevertheless slower than the pass-through to import prices in the short to medium run. One reason is that domestic costs affect export prices and domestic costs are slow to react to exchange rate impulses. The moderate response of export prices affects wage bargaining and slows down the pass-through of prices to wages and further from wages in the tradable goods sector to wages in the service industries This explains partly why the impact of exchange rate changes on consumer prices is fairly modest both in the short and medium term, an empirical finding in line with what is usually found in the literature, see e.g. Berben (2004) and references therein.

Figure 4 shows export prices of traditional goods, which do not include the major export items crude oil, natural gas and shipping services. Prices of total exports

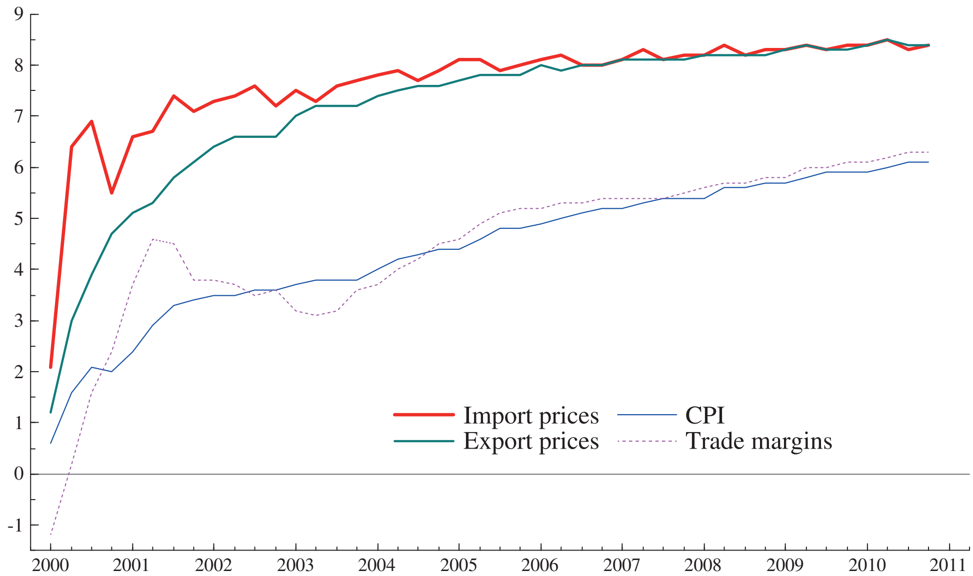


**Fig. 4** Pass-through to prices of a 10 % exchange rate depreciation. Deviations from reference path in per cent

increase much more rapidly than export prices of traditional goods due to the open economy assumption of some important export prices (see Table 4 below). The crude oil price in USD is, as mentioned in Section 2, assumed exogenous for Norwegian producers. When the exchange rate depreciates the price of oil in domestic currency follows the exchange rate.

The exchange rate pass-through to consumer prices is noticeably delayed by the effects on trade margins in the distribution sector, which act as cushions to exchange rate fluctuations in the short run. With these cushions effects included in the CVAR model for trade margins, CPI inflation increases by 1.6 percentage points compared to the reference path during the first year (see Table 3 below). If the trade margins instead had increased in parallel with other consumer prices following the 10 % depreciation of the exchange rate, then the first year effect on CPI would have been 0.3 percentage points higher. Similarly, if trade margins had been constant as a rate on a weighted price average of goods passing through the distribution sector, then the first year effect on CPI would also have been 0.3 percentage points higher.[14] In other words, the value added that fully endogenous trade margins bring to the analysis of overall exchange rate pass-through in the Norwegian economy is estimated to nearly 20 % of the CPI-effect during the first year. We thus argue that the simulation results prove the importance of the distribution sector and the CVAR modelling for trade margins when the purpose is to study the overall exchange rate pass-through process in the Norwegian economy. Otherwise, the study of pass-through will ignore some important channels through which the exchange rate is likely to operate according to our estimated CVAR model.

That said, the pricing-to-market effects from import prices, leading to a moderate increase in the import prices (in domestic currency) of only 5.3 % during the first year (see Table 4 below), are even more important than the buffer effects from the trade margins in slowing down the exchange rate pass-through to consumer prices. If the 10 % depreciation of the exchange rate instead had been immediately and completely passed through to import prices, then CPI inflation would have increased roughly by 3.0 percentage points the first year and not by 1.6 percentage points as with pricing-to-market effects included in the macroeconometric model. According to Fig. 4, trade margins overshoot somewhat during the second year, while no significant contribution from trade margins to the CPI is present after the third year. The exchange rate pass-through to consumer prices is far from complete even within a ten-year horizon despite the fact that the wage-price block inherent in the macroeconometric model (cf. Section 2) is homogenous of degree one, which implies complete pass-through of an exchange rate depreciation in the very long run. Table 3 presents simulation results for the various price components of private consumption, trade margins and some price aggregates.

Generally speaking, prices on consumer goods react faster to changes in the exchange rate than prices on consumer services. This applies in particular to consumer prices for durable goods due to relatively high import shares or high degree of import penetration. The exchange rate pass-through to consumer prices for *Other services*, which constitute around 20 % of total private consumption, is slow both in

[14] This would correspond to a model of trade margins being $TM$ = constant $PP$, cf. equation (11) in Section 3.

**Table 3** Consumer price effects of a 10 % exchange rate depreciation. Deviations from reference path in per cent unless otherwise noted

| | 1. year | 2. year | 3. year | 4. year | 10.year |
|---|---|---|---|---|---|
| Food | 1.7 | 3.5 | 3.9 | 4.1 | 6.4 |
| Beverages | 1.8 | 3.5 | 4.1 | 4.5 | 6.6 |
| Tobacco | 1.6 | 3.5 | 3.9 | 4.0 | 6.2 |
| Petrol, etc. | 2.5 | 4.6 | 5.3 | 5.7 | 7.6 |
| Electricity, heating oil etc. | 1.8 | 3.7 | 4.1 | 4.6 | 6.3 |
| Clothing and footwear | 3.3 | 5.3 | 5.3 | 5.4 | 7.1 |
| Cars | 3.8 | 4.9 | 5.3 | 5.6 | 7.2 |
| Furniture, el. appliances | 2.7 | 4.7 | 5.0 | 5.2 | 7.1 |
| Housing | 0.0 | 1.0 | 1.9 | 1.8 | 3.9 |
| Public transport | 0.1 | 0.7 | 1.3 | 1.7 | 3.7 |
| Health care | 0.8 | 1.9 | 2.7 | 3.2 | 5.6 |
| Other goods | 2.6 | 4.8 | 5.0 | 5.1 | 7.1 |
| Other services | 0.8 | 1.9 | 2.6 | 3.1 | 5.5 |
| Trade margins[1] | 0.8 | 4.2 | 3.7 | 3.3 | 6.0 |
| Consumer price index (CPI) | 1.6 | 3.0 | 3.6 | 3.8 | 5.8 |
| CPI-inflation[2] | 1.6 | 1.4 | 0.6 | 0.2 | 0.2 |
| Core-inflation[2, 3] | 1.6 | 1.4 | 0.6 | 0.2 | 0.2 |

[1] Margins on services in the distribution sector

[2] Percentage points

[3] CPI without energy goods and adjusted for real changes in indirect taxation

the short and medium run. The first year effects on these components of CPI are close to half of the average CPI effect. After 10 years the exchange rate pass-through to prices on *Other services* is still lower than the average pass-through to consumer prices. The reason for the slow exchange rate pass-through to *Other services* is mainly modest pass-through to wages, as production is labour intensive in most services. Also, the direct import share for this consumption category is small.

The effect of the exchange rate deprecation on housing rents is also an important factor behind the overall delayed pass-through to consumer prices. We observe that the third year effect on housing rents is roughly half of the effect on CPI, while the effects after 10 years are two thirds of the CPI effect. Housing rents are determined by user costs of housing capital in the long run, but are mainly indexed to the CPI in the short run as rents typically are based on contracts with an explicit clause linking changes to the CPI. Overall, CPI inflation increases by close to 1.5 percentage points in the first and secondyear compared to the reference path. There is not much difference between the effects on total CPI and core-inflation. We see that prices on the two energy categories (Petrol, etc. and Electricity, heating oil etc.) increase a bit more than CPI, but their total weight in CPI is only 8 %. In addition, excise taxes increase in line with CPI by assumption, such that the effects on core inflation and CPI inflation are similar. Nevertheless, the simulated effects on inflation are clearly important for an inflation targeting central bank.

**Table 4** Macroeconomic effects of a 10 % exchange rate depreciation. Deviations from reference path in per cent unless otherwise noted

| | 1. year | 2. year | 3. year | 4. year | 10.year |
|---|---|---|---|---|---|
| Household consumption | −0.8 | −0.9 | −1.3 | −1.3 | −1.2 |
| Real investments, mainland economy | 0.1 | 0.5 | −0.2 | −0.3 | −0.2 |
| Housing | −0.7 | −1.5 | −1.7 | −1.4 | −0.9 |
| Business sector | 0.6 | 1.8 | 0.5 | 0.1 | 0.1 |
| Exports, traditional goods | 1.4 | 1.5 | 1.5 | 1.4 | 1.1 |
| Imports | −1.0 | −1.1 | −1.4 | −1.5 | −1.3 |
| GDP, mainland economy | 0.2 | 0.3 | 0.1 | 0.1 | 0.0 |
| Manufacturing | 1.9 | 2.5 | 2.7 | 2.7 | 2.2 |
| Employed persons | 0.3 | 0.5 | 0.4 | 0.4 | 0.3 |
| Rate of unemployment[1] | −0.2 | −0.3 | −0.2 | −0.2 | 0.0 |
| Import prices, total | 5.3 | 7.0 | 7.4 | 7.5 | 8.3 |
| Export prices, total | 6.4 | 7.7 | 8.2 | 8.5 | 8.9 |
| Export prices, traditional goods | 3.3 | 5.6 | 6.6 | 7.1 | 8.3 |
| Consumer price index (CPI) | 1.6 | 3.0 | 3.6 | 3.8 | 5.8 |
| Average wage rate | 0.8 | 1.6 | 2.3 | 2.8 | 5.2 |
| Money market rate[2] | 3.0 | 1.7 | 0.4 | 0.3 | 0.3 |

[1] Percentage points

[2] Percentage points pro anno

## 4.2 Effects on the Real Economy

Table 4 presents simulation results for some main macroeconomic variables of the real economy.[15] We see that the effects on mainland GDP and unemployment are small following a 10 % depreciation of the exchange rate. These findings originate from the moderate nominal exchange rate pass-through, which translates into expansive effects with respect to the level of activity. The effects on non-oil exports and import shares are important in this respect. We notice that the positive effects on exports decline over time as the pass-through to export prices increases. Real wages also decline during the entire simulation period. The fall in real wages is, however, most noticeable in the short and medium run, implying that nominal wages adjust more slowly than prices. The institutional system of wage bargaining in Norway contributes to the slow adjustment in nominal wages. Reduction in labour productivity partly explains the fall in real wages in the longer term. The decline in real wages causes demand from households to decline as well. This counteracts the expansionary effects on exports such that mainland GDP does not change much. Consequently, neither employment nor unemployment change much either.

[15] Norway is a large producer and exporter of crude oil and natural gas. Production of petroleum is exogenous in the macroeconometric model and equal to capacity output based on detailed information from oil companies. With domestic demand endogenous, exports of petroleum are endogenous as a residual. In 2010 crude oil and natural gas alone made up roughly 45 % of total exports.

Overall, exchange rate changes do not translate into substantial effects on mainland GDP and unemployment in the short and medium term. After 10 years the effects are close to zero. The main explanation is that the negative effects on household consumption and housing investments are nearly counteracted by the positive effects on non-oil exports and lower import shares due to improved competitiveness. Comparing changes in the CPI with changes in the export prices, we notice much quicker pass-through to export prices. Thus, using export prices rather than consumer prices when calculating the real exchange rate, the shock in the nominal exchange rate produces only moderate effects on the real exchange rate. Even so, we observe that changes in the industry structure are likely to take place in the face of an exchange rate shock of the magnitude studied here as sizeable increase in manufacturing production is simulated both in the short and medium term.

## 5 Conclusions

Much of the new open economy literature examines exchange rate pass-through using relatively aggregated price models. We have followed recommendations of empirically based macroeconomics in Colander et al. (2008) and studied exchange rate pass-through by means of CVAR models for different prices – trade margins in particular – within a large scale macroeconometric model of the Norwegian economy. Our approach has thus taken into account numerous channels through which the exchange rate is likely to operate in a small open economy. We found that exchange rate pass-through to import prices in domestic currency is quite rapid, but nevertheless incomplete within a ten-year horizon. Even if complete exchange rate pass-through to consumer prices follows from the macroeconometric model in the long run, costs of adjustment and other inertia in the pricing behaviour cause pass-through of exchange rate changes to consumer prices to take considerable time. Econometric analysis and model simulations show the importance of the pricing behaviour in the distribution sector as trade margins do play an important buffer role in delaying pass-through to consumer prices in the short run.

The finding that exchange rate pass-through to consumer prices is not fully completed within a ten-year horizon may be interpreted as supporting the LCP hypothesis. Changes in the exchange rate have substantial effects on CPI inflation even in the short run, an insight important for inflation targeting central banks. Since nominal exchange rate changes and inflation seem to move together (at least partly), we may argue that stability in the one implies stability in the other. However, large first and second year effects and relatively high persistent effects in the medium term can make a strategy of stabilising inflation within a two-year horizon, say, by controlling fluctuations in the nominal exchange rate a difficult task. That exchange rate changes do not seem to produce substantial variability in other targeting variables, like GDP growth and unemployment in the short run, suggests that a floating exchange rate has little effects on the real economy. However, exchange rate fluctuations are likely to produce changes in the industry structure of an open economy like the Norwegian. For an economy like the Norwegian that tries to avoid Dutch Disease problems likely to follow from spending large oil revenues, exchange rate fluctuations that result in de-industrialisation may seem just as undesirable.

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## Appendix. Data definitions and sources

| | |
|---|---|
| $TM$ | Trade margins index in the distribution sector (2002=1). The index covers trade margins on all services delivered from the sector. The data on trade margins are taken from the Norwegian national accounts and are based on detailed surveys of trade margins by product every fifth year or so. In between these surveys, average margins by broader product categories are inflated in the national accounts using annual statistics on wholesale and retail trade. Source: Statistics Norway. |
| $PP$ | Price index of purchasing prices faced by distributors in the wholesale and retail trade sector (2002=1). The index weights together domestic and import prices on main commodities traded in the sector. Source: Statistics Norway. |
| $VUC^d$ | Variable unit costs in the wholesale and retail trade sector, defined as in (5) in the text. Source: Statistics Norway. |
| $H$ | Hours worked by self-employed in the wholesale and retail trade sector, expressed in 1000. Source: Statistics Norway. |
| $X^d$ | Gross production in the wholesale and retail trade sector (at fixed 2002 prices), expressed in millions. Source: Statistics Norway. |
| $E$ | Import-weighted nominal exchange rate index (1995=1). The index covers the main Norwegian trading partners, i.e., the 44 countries with the largest weights in Norway's imports of goods. The weight of each trading partner is the imports from this partner as a share of total imports. Source: Central Bank of Norway. |
| $R$ | Average nominal interest rates on bank loans in the private sector. Source: Statistics Norway. |
| $S_{it}$ | Centered seasonal dummy for quarter $i$, equals 0.75 in quarter $i$, −0.25 otherwise, $i$=1,2,3. |
| $D74Q1$, $D78Q2$ | Dummy variables used to account for outliers and instabilities in the equation for $\Delta pp_t$ in the VARs. $D74Q1$ and $D78Q2$ equal unity in the first quarter of 1974 and in the second quarter of 1978, respectively, zero otherwise. |
| $D78Q1$ | Dummy variable used to account for an outlier and instability in the equations for $\Delta pp_t$, $\Delta vuc_t^d$ and $\Delta(h - x^d)_t$ in the VARs. Equals unity in the first quarter of 1978, zero otherwise. |
| $D96Q3$, $D98Q2$ | Dummy variables used to account for outliers and instabilities in the equation for $\Delta(h - x^d)_t$ in the VARs. $D96Q3$ and $D98Q2$ |

| | |
|---|---|
| | equal unity in the third quarter of 1996 and in the second quarter of 1998, respectively, zero otherwise. |
| *PSTOP* | A combined regulation/catch-up dummy variable used to control for governmental price regulations during the second half of the 1970s. Equals unity in years of regulations, minus unity in years of catch-up and zero otherwise, see Bowitz and Cappelen (2001) for details. |

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# 6 Chapter 6

## Trade liberalisation and exchange rate pass-through: the case of textiles and wearing apparels

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# Trade liberalisation and exchange rate pass-through: the case of textiles and wearing apparels

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**Abstract** Studies on the relationship between exchange rates and traded goods prices typically find evidence of incomplete pass-through, usually explained by pricing-to-market behaviour. Although economic theory predicts that incomplete pass-through may also be linked to the presence of non-tariff barriers to trade, variables reflecting such a link is rarely included in empirical models. In this paper, we estimate a pricing-to-market model for Norwegian import prices on textiles and wearing apparels, controlling for non-tariff barriers to trade and shift in imports from high- to low-cost countries. We apply the cointegrated VAR approach and develop measures of foreign prices based on superlative price indices (including the Törnqvist and Fischer price indices) and a data calibration method necessary to approximate relative price levels across countries. Our measures of foreign prices thereby account for inflationary differences *and* varying import shares and price level differences (known as the China effect) among trading partners. We show that these measures of foreign prices, unlike standard measures used in the pricing-to-market literature, are likely to produce unbiased estimates of pass-through. Once the China effect is controlled for, we find little evidence that pass-through has changed alongside trade liberalisation.

**Keywords** Trade liberalisation · The China effect · Import prices · pricing-to-market · Exchange rate pass-through · Vector autoregressive models

**JEL Classification** C22 · C32 · C43 · E31

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## 1 Introduction

A key topic in monetary economics of interest for policy makers in general and inflation targeting central banks in particular is the responsiveness of prices of internationally traded goods to changes in nominal exchange rates. Empirical research on the degree of exchange rate pass-through (henceforth pass-through) is abundant. Typically, existing studies find evidence of incomplete pass-through, which is often explained by pricing-to-market behaviour under conditions of imperfect competition and segmented markets, see e.g. Menon (1995a), Goldberg and Knetter (1997), Gil-Pareja (2003), Herzberg et al. (2003), Campa and Goldberg (2005), Atkeson and Burstein (2008), Bugamelli and Tedeschi (2008), Thomas and Marquez (2009) and Gust et al. (2010). Also, empirical studies of small open economies show that import prices do not fully respond to changes in exchange rates and that domestic market conditions influence the price setting behaviour of foreign firms, see e.g. Menon (1995b), Menon (1996), Naug and Nymoen (1996), Alexius (1997), Kenny and McGettigan (1998) and Doyle (2004).

However, previous studies usually ignore the Bhagwati hypothesis that the presence of non-tariff barriers to trade may affect pass-through, see Bhagwati (1991). The hypothesis says that in the presence of quantity restraints on imports a small depreciation of the exchange rate is likely to be absorbed into the quota rents extracted by the exporter rather than being reflected in import prices. If the depreciation, on the other hand, is large enough to push import prices above the point where the quantity restraints are no longer binding, then pass-through will be positive, but incomplete.

In this paper, we estimate a model for Norwegian import prices on textiles and wearing apparels (henceforth clothing) that controls for the shift in imports from high- to low-cost countries and the gradual removal of non-tariff barriers to trade experienced in the clothing industry since the mid 1990s. The model is based on the pricing-to-market theory by Krugman (1987) and is estimated on quarterly time series data over the period 1986–2008. We apply the cointegrated VAR framework to quantify the degree of pass-through and pricing-to-market, thereby paying attention to the time series properties of the variables involved.

The motivation of our study follows from the fact that low consumer price inflation observed over several years in Norway coincides well with a simultaneous fall in import prices on clothing. The development in import prices on clothing during the last two decades may partly be explained by conventional factors such as shifts in exchange rates, international prices (measured in foreign currency) and domestic market conditions. However, it should also be viewed in light of the trade liberalisation, which led to the massive increase in imports of clothing from China and other low-cost countries at the expense of imports from high-cost countries, the euro area in particular. The significant deflationary effect on traded goods prices of shifts in the country composition of imports has been dubbed the *China effect* and is likely to be important when quantifying pass-through in regression models. The gradual removal of quota restrictions on trade may in accordance with the Bhagwati hypothesis have pushed the estimate of pass-through upwards, an empirical question which we pursue in the present paper.

To answer this question, we construct three different measures of foreign prices to be used in the estimation of pass-through. Index number theory advocates the use of the so-called superlative index number formulas when the aim of the aggregation problem is to account for flexible substitution effects between commodities caused by relative price level changes, see e.g. Diewert (1976, 1978). Our first two measures of foreign prices are thus based on the Törnqvist and Fischer price indices, which both belong to the class of superlative price indices. The appealing aggregation properties are, however, somewhat counterbalanced by the fact that available data on foreign prices on clothing are price *indices* and not price *levels*, which makes the superlative price indices (like any other index number formulas) not directly ready for numerical calculations in our context. If the available set of price indices is plugged directly into the superlative price indices, only inflationary impulses implied by price changes and substitution between goods with different price changes are accounted for in the final price aggregate. We, therefore, suggest a data calibration method based on purchasing power parities to account for not only *inflationary differences* as is typical in the pricing-to-market literature, but also *varying import shares* and *differences in price levels*—that is the China effect—among trading partners when constructing the superlative price index measures of foreign prices.[1] Our third measure of foreign prices is based on the often used geometric mean price index with constant import shares as weights, a measure which fails to take account of the China effect. By comparing the estimates of pass-through that come out of modelling the import price of clothing with the alternative measures of foreign prices, we are able to shed some light on the potential problem of omitted variable bias in empirical tests of pricing-to-market.

We find that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high- to low-cost countries on average has reduced the international price impulses on imports of clothing by around 2 percentage points per year since the early 1990s. Controlling for these effects by means of the superlative price index measures of foreign prices, we estimate import price models consistent with the pricing-to-market hypothesis. Specifically, the pass-through and pricing-to-market elasticities are significantly estimated to 0.44 and 0.56, respectively, irrespective of using the Törnqvist or the Fischer price index measure of foreign prices in the regression model. In contrast, we find that the use of the geometric mean price index measure of foreign prices with constant weights biases the estimates due to international price impulses being substantially overestimated. We also establish that the estimated dynamic model is reasonably stable in sample

[1] To our knowledge, no previous studies have estimated pricing-to-market models with a superlative price index measure of foreign prices. Generally, there are few academic papers which examine the impact of increased imports from China and other low-cost countries on traded goods prices and overall inflation in developed countries. Thomas and Marquez (2009) estimate a pricing-to-market model for aggregated US import prices with a geometric Paasche price index (with varying weights) measure of foreign prices, but do not decompose the price measure into its different inflationary and price level components as we do in the present paper. Nickell (2005) computes the China effect on traded goods prices based on the geometric Paasche price index, but does not estimate a model when analysing the impact of a changing trade pattern on import prices and overall consumer price inflation in the UK. Wheeler (2008) also calculates the China effect on traded goods prices following the operational route in Nickell (2005). In addition, Wheeler (2008) estimates panel regressions of UK inflation by goods category on the level and growth of the import share from China as the main determinants.

and exhibits no serious forecasting failures around the dates of the shifts in trade policy. That no serious structural breaks are detected may reflect that likely pass-through effects of changes in trade policy are controlled for through the superlative price index measures of foreign prices. Consequently, once the effect of shifts in imports towards low-cost countries is controlled for, we find little evidence that the properties of the import price equation have changed alongside trade liberalisation.

Based on our findings, we claim that the choice of aggregation formula for foreign prices matters for the quantification of an import price model, an issue typically ignored in the pricing-to-market literature. We emphasise that the empirical example of clothing is *not* a special case as trade liberalisation has led to increased exports of several product categories (not just clothing) from China and other low-cost countries to Norway and other high-cost countries over the last two decades or so. There exist a large parallel literature on trade that explicitly discusses and demonstrates the importance of choosing the relevant price index to incorporate new products (or countries entering or leaving the market for a particular good) into an aggregate of international prices, see e.g. Feenstra (1994) and Broda and Weinstein (2006). Gaulier et al. (2008) present extensive empirical evidence that the choice of aggregation method matters for the calculation of international prices and show that the Törnqvist and Fischer price indices provide similar results in practice.

The rest of the paper is organised as follows: Sect. 2 outlines the pricing-to-market theory for a small open economy and discusses the effects of non-tariff barriers to trade on pass-through. Section 3 presents the construction of the alternative measures of foreign prices and the data used in the empirical analysis. Section 4 describes and reports results from the cointegrated VAR modelling, while Sect. 5 presents the estimated dynamic model for import prices on clothing. Section 6 concludes.

## 2 The theoretical framework

The underlying theoretical model for the behaviour of import prices on clothing is based on the pricing-to-market theory by Krugman (1987). Markets for clothing are typically characterised by imperfect competition between firms producing differentiated products. Furthermore, these markets are segmented due to trade barriers, transportation costs and imperfect information. Profit maximisation under these circumstances normally implies that foreign exporters can charge different markups over their marginal costs, and hence can charge different prices, depending on the conditions in each particular market. The following exposition of the pricing-to-market model and the relationship between pass-through and the presence of (and removal of) non-tariff barriers to trade build on Naug and Nymoen (1996) and Menon (1996).

### 2.1 Pricing-to-market

Consider a representative foreign firm producing a differentiated product of clothing exported to $n$ segmented markets or countries ($i = 1, \ldots, n$). The product is assumed to be weakly separable from all other competing goods in the consumer's utility function. The demand faced by the firm in each export market may then be expressed as

$X_i = X_i(PX_i \cdot ER_i, PQ_i, DP_i)$, where $PX_i$ is the firm's export price measured in the exporter's currency, $ER_i$ is the bilateral exchange rate with respect to country $i$, $PQ_i$ is an index of prices on competing products and $DP_i$ represents other factors affecting demand (henceforth referred to as demand pressure). The profit of the firm is given by

$$\prod(PX_1, \ldots, PX_n) = \sum_{i=1}^{n} PX_i \cdot X_i\left(PX_i \cdot ER_i, PQ_i, DP_i\right) - C\left[\sum_{i=1}^{n} X_i\left(PX_i \cdot ER_i, PQ_i, DP_i\right), W\right], \tag{1}$$

where $C[\cdot]$ is the cost function depending on production and input prices ($W$). Time arguments are provisionally suppressed for simplicity. Profit maximisation generates the following first order conditions

$$PX_i = \lambda_i MC, \quad i = 1, \ldots, n. \tag{2}$$

Hence, the foreign firm sets each export price as a markup ($\lambda_i$) on the common marginal costs ($MC$) measured in the currency of the exporter. Generally speaking, $\lambda_i = \eta_i/(\eta_i - 1)$, where $\eta_i = \eta_i((PX_i \cdot ER_i), PQ_i, DP_i)$ is the elasticity of demand in market $i$. As every export price reflects conditions in each particular market, profit maximisation typically leads to price discrimination, and thus market-specific markups. The import price ($PI_i$) measured in the currency of the importing country $i$ is obtained by multiplying through (2) with the bilateral exchange rate $ER_i$.

$$PI_i = ER_i PX_i = ER_i \lambda_i MC, \quad i = 1, \ldots, n. \tag{3}$$

Following Naug and Nymoen (1996), we abstract from competition between foreign firms in market $i$ to simplify matters and specify the destination specific markup as $\lambda_i = K_i(PD/PI)_i^{\gamma_{1i}} DP_i^{\gamma_{2i}}$, where $K_i$ is a constant, $PD_i/PI_i$ is the price on competing goods *produced* in market $i$ relative to the import price and $DP_i$ is the demand pressure in the importing country. Economic theory predicts that $\gamma_{1i} \geq 0$ because higher prices on competing goods imply a potential for increasing markups. The sign of $\gamma_{2i}$ is, however, undetermined from theory. An increase in the demand pressure may rise the scope for an increase in the markup, but may very well also increase economies of scale in production and distribution, and hence pave the way for a decrease in the markup. Substituting the expression for $\lambda_i$ into (3) and using lower case letters to indicate natural logarithms, we obtain[2]

$$pi_i = \kappa_i + (1 - \psi_i)(mc + er_i) + \psi_i pd_i + \delta_i dp_i, \quad i = 1, \ldots, n, \tag{4}$$

where $\kappa_i = \ln K_i/(1 + \gamma_{1i})$, $\psi_i = \gamma_{1i}/(1 + \gamma_{1i})$ and $\delta_i = \gamma_{2i}/(1 + \gamma_{1i})$. When $\psi_i > 0$ domestic prices ($pd_i$) matter for the determination of import prices, and changes in

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

marginal costs and the exchange rate are not entirely passed through to import prices. This phenomenon is what Krugman (1987) labelled pricing-to-market. The degree of pass-through from $mc$ and $er_i$ to $pi_i$ is given by the coefficient $(1 - \psi_i)$. In the special case when $\psi_i = 0$, the pass-through from $mc$ and $er_i$ is complete, and $pd_i$ has no role in the determination of import prices. Conversely, $\psi_i = 1$ implies zero pass-through.

The law of one price (henceforth LOP) is the standard assumption of import pricing in theoretical models of small open economies, and follows as a special case of (4). As pointed out by Naug and Nymoen (1996), the absolute version of LOP requires full pass-through ($\psi_i = 0$), no effects from domestic demand pressure ($\delta_i = 0$) and the same markup ($\kappa_i = \kappa > 0$) in all countries, which implies that $PX_i = PX$ in all markets. The relative version of LOP, on the other hand, only requires ($\psi_i = \delta_i = 0$) in all countries. Hence, the relative version of LOP allows price discrimination through a varying constant ($\kappa_i$)—which under both versions of LOP equals the markup ($\lambda_i$)—across markets.

The pricing-to-market model outlined here is based on foreign firms' price setting behaviour and two channels through which domestic factors in the importing country may affect import prices on clothing, namely through competitive pressure ($pd_i$) and demand pressure ($dp_i$) in the importing country. Another motivation for including $pd_i$ and $dp_i$ in the model would be when importers of clothing act as agents and find domestic factors important in price negotiations with foreign producers. The model implicitly assumes, on the other hand, that markets for clothing are segmented due to inter alia presence of non-tariff barriers to trade. As previously mentioned, such trade barriers may limit the degree of pass-through according to the Bhagwati hypothesis, an issue which we now turn to.

### 2.2 Non-tariff barriers to trade

Before the Uruguay Round in 1986, the clothing industry was among the most strictly regulated manufacturing sectors, both in terms of tariffs and quantity restrictions on trade. During the 1970s and 1980s, the Norwegian market for clothing was mainly regulated through the Multi-Fibre Agreement, an agreement that allowed importers to negotiate bilateral export restraint quotas with low-cost countries. The Uruguay Round, however, led to major changes in the trade policy and it was decided that quota regulations should be eliminated between 1995 and 2005. Norway was relatively quick in liberalising the quota system and the last quantity restrictions on trade with clothing were abolished in 1998. The removal of quotas has no doubt contributed significantly to further increase in imports of clothing from low-cost countries during the last 10–15 years. Substantial reduction over time in tariff rates on imports of clothing has likewise pulled in the same direction.[3]

Here, we shall focus on the link between non-tariff barriers to trade and pass-through as the empirical analysis is based on import prices of clothing exclusive tariffs. The

[3] For instance, the average ordinary tariff rate was reduced from about 20 % in 1994 to 12 % in 2004. See Melchior (1993) and Høegh-Omdal and Wilhelmsen (2002) for summaries of clothing trade policies in Norway.

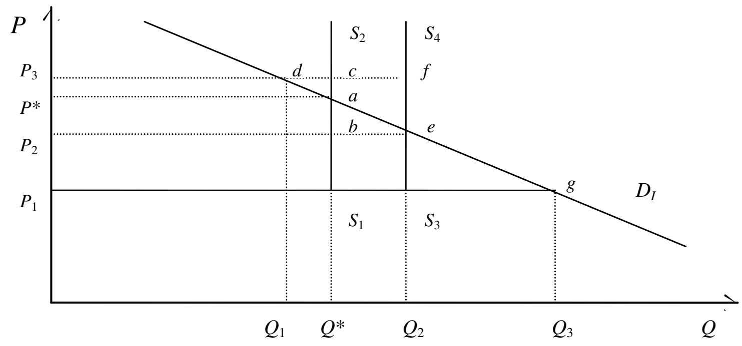


**Fig. 1** Pass-through with presence and removal of quota restrictions

effects of quantity restrictions on pass-through do not depend on particular market structures. We, therefore, extend Menon (1996) analysis and highlight the relationship between pass-through and gradual removal of non-tariff barriers to trade by means of a small country being a price taker with respect to its imports. Figure 1 illustrates the implications of the Bhagwati hypothesis for pass-through in the presence and removal of quantity restrictions on trade.

The demand curve for imports is represented by $D_I$, whereas the supply curve consists of the horizontal line $P_1S_1$ and the vertical line $S_1S_2$. The supply curve is perfectly elastic at $P_1$ (reflecting the small country assumption) and becomes perfectly inelastic when the quantity restrictions on trade are met at $Q^*$. The initial equilibrium is at point $a$ with quantity $Q^*$ and price $P^*$. At point $a$, the seller is able to pull out $P_1S_1aP^*$ in quota rents due to the presence of quantity restrictions.

A small depreciation of the importing country's currency will shift the horizontal part of the supply curve upwards, while the vertical part is unchanged. For example, a depreciation of the currency to $P_2$ will neither affect equilibrium quantity nor market price, but will reduce the quota rents to $P_2baP^*$. It follows that the depreciation is entirely absorbed into the quota rents and that pass-through is zero. However, if the depreciation is large enough to push the market price above $P^*$ to say $P_3$, the horizontal supply curve ($P_3c$) will shift to a level where the quantity restrictions are no longer binding. At the new equilibrium point, $d$ quantity falls below the quota limit to $Q_1$ and the market price increases from $P^*$ to $P_3$. Hence, some part of the currency depreciation is now passed through to the import price. Specifically, the degree of pass-through in this situation equals the change in the market price relative to the magnitude of the currency depreciation, that is $(P_3 - P^*)/(P_3 - P_1) < 1$ as $P^* > P_1$.

Suppose instead that trade liberalisation takes place so that quantity restrictions on trade are effective at $Q_2$ rather than at $Q^*$. Consequently, the horizontal supply curve is represented by the line $P_1S_3$, whereas the vertical supply curve (which shifts to the right alongside the reduction in the quota restrictions) is represented by the line $S_3S_4$. The new initial equilibrium is at point $e$ with quantity $Q_2$, price $P_2$ and quota rents $P_1S_3eP_2$. We notice that $P_1S_1aP^* > P_1S_3eP_2$. The possibilities to absorb currency depreciations into the quota rents are reduced in situation $e$ compared to situation $a$ as $P^* > P_2$. If a currency depreciation again pushes the market price to $P_3$, so that the horizontal supply curve shifts to the line $P_3f$, the equilibrium point $d$ is still reached.

However, both the quantity and the market price will change relatively more for a given currency shock when the initial equilibrium is at point $e$ rather than at point $a$, where no reduction in the quota restrictions has yet taken place. In other words, a reduction in the quota restrictions from $Q^*$ to $Q_2$ implies that pass-through to import prices will be higher, other things equal. To see this, we notice that the degree of pass-through in situation $e$ equals $(P_3 - P_2)/(P_3 - P_1)$, which is greater than $(P_3 - P^*)/(P_3 - P_1)$ because $P_2 < P^*$. Pass-through is still incomplete in situation $e$ as $P_2 > P_1$. Only when the quantity restrictions on trade are entirely removed, as in situation $g$ in Fig. 1, will pass-through be complete.

To summarise, a currency depreciation in the presence of non-tariff barriers to trade will generally reduce the quota rents first, hence absorbing much of its impact, before it is reflected in the market price. It is only when the depreciation is large enough to push the market price above the point where the quota restrictions are no longer binding that pass-through will be positive, but incomplete according to the Bhagwati hypothesis. Finally, if incomplete pass-through is inter alia linked to the presence of non-tariff barriers to trade, gradual removal of such barriers will push pass-through upwards, other things equal.

## 3 From theory to empirics

Because the focus is on aggregated time series for one destination country, namely Norway, we first translate (4) into a testable empirical representation by replacing the index $i$ with the subscript $t$ to denote time. We further replace marginal costs, which are not directly observable, with three measures of foreign prices based on (i) the Törnqvist price index $(pf_t^T)$ with varying import shares as weights, (ii) the Fischer price index $(pf_t^F)$ with varying import shares as weights and (iii) the geometric mean price index $(pf_t^G)$ with constant import shares as weights. Besides, we approximate domestic prices and demand pressure with variable unit costs $(vc_t)$ and the unemployment rate $(UR_t)$, respectively, and add a disturbance term $(u_t)$ to (4). The following empirical representation of (4) emerges:

$$pi_t = const. + (1 - \psi)(pf^i + er)_t + \psi vc_t + \delta UR_t + u_t, \quad i = T, F, G, \tag{5}$$

where $\psi = \gamma_1/(1 + \gamma_1)$ and $\delta = \gamma_2/(1 + \gamma_1)$. Because the unemployment rate enters (5) without a logarithmic transformation the markup is specified as $\lambda_t = K(VC/PI)_t^{\gamma_1} \exp^{(\gamma_2 UR_t)}$. A testable implication of LOP from (5) is that $(pi - pf^i - er)_t$ is stationary or forms a long run cointegration relationship when the variables involved all are non-stationary. If imports and domestic products of clothing are close substitutes, we expect LOP to be a reasonable approximation and pass-through to be nearly complete. The long run version of PPP implies similarly that $(vc - pf^i - er)_t$ is stationary. As noted by Naug and Nymoen (1996), the long run versions of LOP and PPP may be consistent with (5) rewritten as $(pi - pf^i - er)_t = const + \psi(vc - pf^i - er)_t + \delta UR_t + u_t$. We see that this equation is balanced when $\psi > 0$, $\delta \neq 0$ and $(pi - pf^i - er)_t$, $(vc - pf^i - er)_t$, $UR_t$ and $u_t$ all are stationary variables. If both LOP and PPP hold in the long run, then pricing-to-market is only a short run phenomenon and (5) predicts the existence of two cointegrating vectors relating the variables. On the other hand, if $(pi - pf^i - er)_t$ and $(vc - pf^i - er)_t$ are nonstationary, neither LOP nor PPP holds in the long run, and $(pi - pf^i - er)_t - \psi(vc - pf^i - er)_t$ is stationary. In this case, pricing-to-market is a long run phenomenon.

We also notice that (5) imposes the same coefficient on $pf_t^i$ and $er_t$ as well as unit homogeneity between $pi_t$, $(pf^i + er)_t$ and $vc_t$. In practice, however, these restrictions need not hold. Exchange rates are typically more volatile than costs, and foreign exporters may be more willing to absorb into their markups changes in exchange rates (which are likely to be permanent) than changes in costs. We test the parameter restrictions in the empirical analysis rather than imposing them from the outset.

Finally, Naug and Nymoen (1996) emphasise that the use of a geometric mean of export prices proxying marginal costs induces measurement errors as the disturbance term contains the foreign producers' markups. The disturbance term $u_t$ is thus correlated with the export price measure and (5) only forms a cointegration relationship when the measurement errors are stationary. We show in Sect. 4 that a well-specified VAR model and a significant cointegrating vector exist using our data set. Thus, judged by statistical criteria, measurement errors seem to be a minor problem. The disturbance term in (5) also contains domestic producers' markups when we replace domestic prices by variable unit costs. As pointed out by Naug and Nymoen (1996), these measurement errors may be correlated with the unemployment rate representing demand pressure. If markups of foreign firms are affected by domestic demand pressure, we expect that markups of domestic firms also are influenced. We therefore acknowledge that effects of demand pressure would be overestimated in (5) to the extent that $u_t$ is correlated with $UR_t$. Again, our estimated import price model is well specified, indicating that $u_t$ and $UR_t$ are not much correlated.

### 3.1 The measure of foreign prices

As numerous index number formulas with different aggregation properties exist, see e.g. Balk (2008) for a survey, we are faced with the problem of which to choose to account properly for the China effect in the measure of foreign prices. Two commonly used index number formulas in the literature on trade are the Laspeyres and Paasche price indices, which both measure the evolution of prices between a base and a comparison period for a given basket of goods. These price indices, however, produce biased measures of price evolutions for two main reasons. Firstly, as the assigned weights are from a single period only, they fail to capture any substitution effect among different goods from one period to another. Whereas the Laspeyres price index tends to overestimate price growth because it uses the weights from the base period, the Paasche price index tends to underestimate price growth by assigning larger weights to products with increased quantity following a relative price decrease. Secondly, both the Laspeyres and Paasche price indices fail to capture disappearance of old or appearance of new products between the base period and the comparison period. Ignoring new products generally leads to overestimation of price growth, see e.g. Feenstra (1994) and Broda and Weinstein (2006).

Both sources of bias discussed above may be dealt with by chaining the Laspeyres and Paasche price indices. Such chained indices account for changes in the basket of goods, including new products, through variations in the weights assigned to each product entering the basket. However, the chained indices still suffer from a measurement bias as they only use information from one of the two periods in each elementary index. Feenstra (1997) show empirically that the Laspeyres and Paasche price indices represent the upper and lower bounds of the real price development.

The superlative Törnqvist and Fischer price indices, on the other hand, use information from both the base and the comparison period by composing the Laspeyres and Paasche price indices, see e.g. Diewert (1976, 1978). Whereas the Törnqvist price index is defined as the geometric mean of the geometric Laspeyres and Paasche price indices, the Fischer price index is defined as the geometric mean of the arithmetic Laspeyres and Paasche price indices, see e.g. Balk (2008). The fact that all information at hand is utilised motivates us to apply the Törnqvist and Fischer price indices as the underlying index number formulas for the measure of foreign prices. The Törnqvist price index ($PF^T$) in our context equals

$$PF_t^T \equiv \left(\prod_{j=1}^{n} PF_{j,t}^{s_{j,t-1}} \prod_{j=1}^{n} PF_{j,t}^{s_{j,t}}\right)^{1/2} = \prod_{j=1}^{n} PF_{j,t}^{\bar{s}_{j,t}}, \tag{6}$$

where the expressions $\prod_{j=1}^{n} PF_{j,t}^{s_{j,t-1}}$ and $\prod_{j=1}^{n} PF_{j,t}^{s_{j,t}}$ are the geometric Laspeyres and Paasche price indices, respectively, $\bar{s}_{j,t} \equiv \frac{s_{j,t}+s_{j,t-1}}{2}$ for $j = 1, \ldots, n$, $s_{j,t-1}$ and $s_{j,t}$ are the value shares of imports from trading partner $j$ in the base period $t-1$ and the comparison period $t$, respectively, $0 \leq s_{j,h} < 1$ and $\sum_{j=1}^{n} s_{j,h} = 1$ for $h = t-1, t$. We observe that $PF_t^T$ is a weighted geometric average of the foreign price indices ($PF_{j,t}$), the weights being the arithmetic means of the value import shares of the base and comparison period. Aggregating the foreign price indices by means of (6) directly will only capture inflationary impulses because a price index by construction measures the percentage change in a price relative to a base period. We, therefore, suggest a calibration method based on purchasing power parities to approximate relative price levels to accommodate both inflationary impulses and price level differences across countries in the measure of foreign prices.

The first step of our data calibration method involves constructing calibration coefficients for each trading partner, labelled $\lambda_j$, by the formula

$$\lambda_j = \frac{\frac{GDP_j^{NOM}}{GDP_j^{PPP}}}{\frac{GDP_{numeraire}^{NOM}}{GDP_{numeraire}^{PPP}}}, \tag{7}$$

where $GDP_j^{NOM}$ and $GDP_j^{PPP}$ are nominal GDP and purchasing power parity adjusted volume of GDP for trading partner $j$, respectively, and $GDP_{numeraire}^{NOM}$ and $GDP_{numeraire}^{PPP}$ are corresponding GDP figures for the numeraire country. We point out that the calibration coefficients are unitless and easy to interpret for our purposes. For instance, a $\lambda_j$ equal to 0.5 would imply that the overall price level in country $j$ is 50 % of that in the numeraire country.

The second step of our data calibration method involves multiplying the calibration coefficients from (7) with the corresponding price indices from (6). Formally, we rewrite (6) as

$$PF_t^T = \prod_{j=1}^{n} PF_{j,t}^{*\bar{s}_{j,t}}, \tag{8}$$

where $PF_{j,t}^* = \lambda_j PF_{j,t}$. The calibrated price indices $PF_{j,t}^*$ in (8)—which are to be interpreted as relative price levels—equal the relative price levels calculated from (7) in the base period, in which the price indices $PF_{j,t}$ are set equal to unity. In all other periods, the calibrated price indices develop according to the development of the levels of respective original price indices. Formula (8) is an aggregate of foreign prices that accounts for the *total* price effects of the shift in imports towards low-cost-countries.

Analogous to (8), the Fischer price index ($PF^F$) in our context equals

$$PF_t^F = \left( \frac{\sum_{j=1}^{n} s_{j,t-1} PF_{j,t}^*}{\sum_{j=1}^{n} s_{j,t} \frac{1}{PF_{j,t}^*}} \right)^{1/2}, \tag{9}$$

where the expressions $\sum_{j=1}^{n} s_{j,t-1} PF_{j,t}^*$ and $\frac{1}{\sum_{j=1}^{n} s_{j,t} \frac{1}{PF_{j,t}^*}}$ are the Laspeyres and Paasche price indices (calibrated with $\lambda_j$ for our purposes), respectively. Aggregating the foreign price indices by means of (9) will, just like (8), produce a final index number that measures the *total* price effects of the shift in imports towards low-cost-countries. We show below that (8) and (9) generate similar foreign price aggregates in our case. Hence, we shall in the following decomposition of the *total* price effects into inflation and price level effects concentrate on the Törnqvist price index as the underlying index number formula. To simplify matters in the exposition without loss of generality, we only consider two trading partners ($j = 1, 2$). First, taking natural logarithms of (8) and differencing once, we obtain

$$\begin{aligned} \Delta pf_t^T &= pf_t^T - pf_{t-1}^T \\ &= \bar{s}_{1,t} pf_{1,t}^* + \bar{s}_{2,t} pf_{2,t}^* \\ &\quad - \bar{s}_{1,t-1} pf_{1,t-1}^* - \bar{s}_{2,t-1} pf_{2,t-1}^*, \end{aligned} \tag{10}$$

where $\Delta$ indicates the first difference operator. Then, adding and subtracting $\bar{s}_{1,t} pf_{1,t-1}^*$ and $\bar{s}_{2,t} pf_{2,t-1}^*$ to the right hand side of (10), making use of the adding up condition of the value shares of imports and collecting terms, we get an expression for the percentage change in the aggregate foreign price in period $t$ that reads as

$$\begin{aligned} \Delta pf_t^T &= \bar{s}_{1,t} \Delta pf_{1,t}^* + \bar{s}_{2,t} \Delta pf_{2,t}^* \\ &\quad + \Delta \bar{s}_{1,t} \left( pf_{1,t-1}^* - pf_{2,t-1}^* \right). \end{aligned} \tag{11}$$

By calculating $\Delta pf_t^T$ in this way, we allow for inflationary and price level differences as well as varying import shares among the main Norwegian trading partners. The first two terms on the right hand side of (11) show that increasing inflation on clothing from each of the trading partners contribute to increasing inflationary impulses faced by Norwegian importers. The larger the price increase and the larger the import share, the larger is the inflationary impulse (measured in foreign currency) in $\Delta pf_t^T$. The last term on the right hand side of (11) constitutes the total effect of the price level differences, that is the China effect. If the import share is changing in favour of a low-cost country, the last term becomes negative. The larger the change in the import share and the larger the difference in price levels, the larger is the deflationary impulse in $\Delta pf_t^T$. We notice that the China effect is zero only in the special cases when the import shares are constant ($\Delta \bar{s}_{1,t} = 0$) and when the composition of trade changes between countries with identical price levels ($pf_{1,t-1}^* - pf_{2,t-1}^* = 0$). Although the bilateral distribution of the China effect can be sensitive to the choice of numeraire country, the size of the *aggregated* China effect calculated from (11) is not. The level of aggregate foreign prices may now be calculated by setting $PF_t^T$ equal to unity in the base period and letting the level of the price index from then on be determined consecutively by the measured growth rates from (11).

The standard practise in related studies is to weight together some proxy for foreign prices by means of a geometric mean price index with constant import shares as weights, see e.g. Naug and Nymoen (1996), Kenny and McGettigan (1998), Herzberg et al. (2003) and Campa and Goldberg (2005). The geometric mean price index ($PF^G$) in our context, analogous to (8), equals

$$PF_t^G = \prod_{j=1}^{n} PF_{j,t}^{*\bar{s}_j}, \tag{12}$$

where the exponent $\bar{s}_j$ now is the constant value share of imports from trading partner $j$, $0 \leq \bar{s}_j < 1$ and $\sum_{j=1}^{n} \bar{s}_j = 1$. Following Naug and Nymoen (1996), we set $\bar{s}_j$ equal to the average of each import share over the sample period. Again, differencing once the natural logarithms of (12), adding and subtracting $\bar{s}_1 pf_{1,t-1}^*$ and $\bar{s}_2 pf_{2,t-1}^*$, making use of the adding up condition of the value shares of imports and collecting terms, we get

$$\Delta pf_t^G = \bar{s}_1 \Delta pf_{1,t}^* + (1 - \bar{s}_1)\, \Delta pf_{2,t}^*. \tag{13}$$

We see that (13) only accounts for inflationary differences among the trading partners. Accordingly, international price impulses are overestimated to the extent that China effects are present. On this background, we expect that the estimate of the degree of pass-through will reflect an omitted variable bias when $PF_t^G$ is used instead of the superlative price indices in a regression model for import prices of clothing, other things equal. One way to remedy this potential econometric problem may be to add a linear trend to approximate the price level term in (11). However, we thereby implicitly assume that the China effect has been constant over the sample period, a strict assumption to impose on the regression model from the outset. A linear trend is just a 'measure of ignorance' and at best it represents an omitted variable in the regression model. We, therefore, argue in this paper that a more flexible and reliable approach is to allow the China effect, and thereby also the consistency of the degree of pass-through, to be entirely controlled for through the superlative price indices.

### 3.2 Data[4]

We use quarterly, seasonally unadjusted time series covering the period 1986Q1–2008Q1. The import price ($pi_t$) is an implicit deflator for imports of clothing with Norwegian substitutes measured in Norwegian currency. The products comprising the deflator are priced *cif* at the Norwegian border. Hence, prices include costs of insurance and freight, but exclude tariffs. The deflator is a chained geometric mean price index calculated by weighting together each unit price, which is based on the value and volume of each single imports, with the corresponding import share (measured in value) of each trading partner. Because the import shares are continuously updated in accordance with the development in the country composition of clothing imports, the deflator reflects the shifts in imports from high- to low-cost countries over time.

To construct the superlative price indices ($pf_t^T$ and $pf_t^F$), we need data on import shares, export prices and price levels for each one of the main trading partners. The foreign trade statistics provide time series of import shares by country. The main exporters of clothing to Norway are China ($ch$), the euro area ($eu$), the United Kingdom ($uk$), Denmark ($dk$), Sweden ($sw$), Hong Kong ($hk$) and Turkey ($tr$). Together, these countries covered nearly 80 % of Norwegian imports of clothing as an average over the sample period.[5] Because the euro area is treated as one country, we abstract from any import substitution from high- to low-cost countries within this area.

It proved difficult to find long and consistent proxies for export prices for China and Turkey. We, therefore, approximate Chinese export prices by connecting producer prices on clothing available from 1997$Q$1 together with consumer prices on all products available from 1986$Q$1. The fact that these two time series are highly correlated during the period 1997Q1–2008Q1 may make consumer prices a fairly good proxy for producer prices of clothing during the first half of the sample period. Similarly, we connect Turkish export prices on clothing available from 2004$Q$1 together with export prices on manufactures available from 1995$Q$1 and import prices on all products available from 1986$Q$1.

Price level differences among the trading partners should ideally be based on comparable price levels on clothing that reflect the level of production costs corrected for the level of productivity in each country. Because such data are not available, we use purchasing power parity-adjusted GDP figures provided by IMF. Table 1 shows the

[4] See the Appendix for details about data definitions and sources.

[5] The rest of exports of clothing to Norway came from countries with relatively small import shares during the 1980s and 1990s. Indeed, Bangladesh was represented by an import share of about 8 % in 2008, but is left out of the analysis due to lack of relevant price data.



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**Table 1** Average relative price levels ($\lambda_j$). 1991–2008

| *dk* | *sw* | *uk* | *eu* | *hk* | *tr* | *ch* |
|---|---|---|---|---|---|---|
| 1.30 | 1.25 | 1.05 | 1.00 | 0.91 | 0.56 | 0.41 |

*Sources*: IMF and Statistics Norway

average calculated international relative price levels ($\lambda_j$) over the period 1991–2008 based on (7).[6]

As the euro area is chosen as numeraire country, $\lambda_{eu}$ equals unity. Our calculations show that the overall price level in China is 41 % of that in the euro area. The corresponding figure for Turkey is 56 %. Hence, both China and Turkey stand out as low-cost countries in our study. We recognise that the relative price levels in Table 1 are good proxies only to the extent that relative price levels on clothing are similar to relative GDP deflators across countries, an assumption that needs not hold in practise. For instance, it may be the case that exporters of clothing from low-cost countries set their prices somewhat below the competitors' prices to gain market shares. Consequently, the price level of imports from low-cost countries may be higher than that calculated from the purchasing power parity-adjusted GDP deflators. If this is the case, the calculated superlative price index measures of foreign prices, based on the figures in Table 1 will overestimate the true negative price level impulses to the Norwegian economy. The superlative price indices may, on the other hand, overestimate the true international price impulses as consumer prices, which also include mark-ups on domestic costs of distribution not faced by Norwegian importers, approximate Chinese export prices of clothing in the first half of the sample period. We shed some light on the sensitivity of the superlative price indices, and thereby the sensitivity of the estimates of pass-through and pricing-to-market, when the relative price levels for China and Turkey in Table 1 are increased and decreased by 50 %, other things equal. Nevertheless, the relative price levels in Table 1 are used as benchmarks to calibrate the respective export price indices in (8) and (9).

Figure 2 displays the country-specific export prices ($pf_{j,t}$), measured in foreign currency and normalised to unity in 1986$Q$1. We observe that the export prices of clothing from high-cost countries increased quite substantially during the first half of the sample period, possibly reflecting high economic growth and steady demand in their export markets. In the wake of the Asian financial crises, which started in Thailand in July 1997, high-cost countries generally faced reduced export possibilities and stronger price competition from the Asian countries with depreciated currencies. The price competition among trading partners was further amplified by the increased presence of low-cost countries on international markets following the trade liberalisation after the Uruguay Round. Additionally, imports from China increased when the country joined the WTO in 2001 and the international economic downturn in 2002 gave rise to reduced export possibilities for most high-cost countries in the successive

[6] Data for purchasing power parity-adjusted GDP are not available on a quarterly basis, and only from 1991 onwards for the euro area. Because the calculated price level series appear relatively stable we assume constant price levels equal to the average over the period 1991–2008.

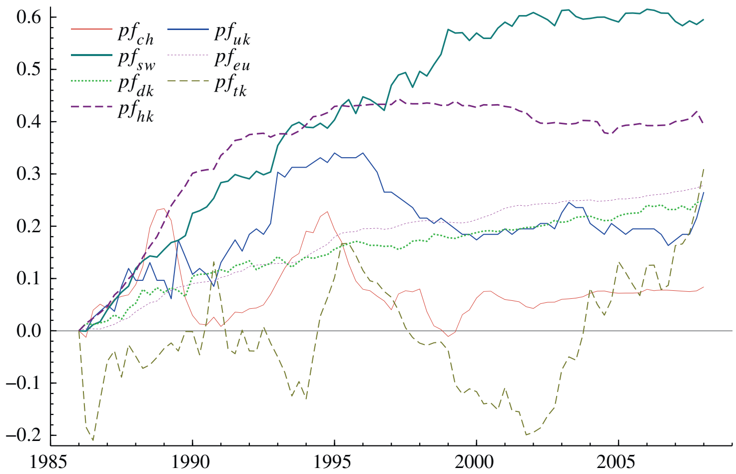


**Fig. 2** Time series for foreign prices ($pf_{j,t}$). Source: See the Appendix

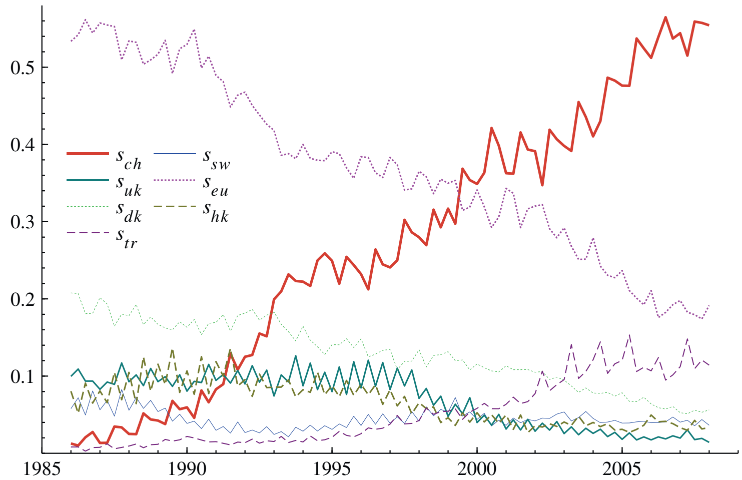


**Fig. 3** Time series for import shares ($s_{j,t}$). Source: See the Appendix

years. Together, these economic features generally may have led exporters of clothing in high-cost countries to lower their markups over costs during the second half of the sample period.

Figure 3 displays the country-specific import shares ($s_{j,t}$), which sum to unity in each period. We see that the import share from China increased from a few per cent in 1986 to around 55 % in 2008. The import share from the euro area fell likewise from around 55 % in 1986 to around 20 % in 2008. After a substantial increase in the import share from the mid 1990s, Turkey supplied more than 10 % of Norwegian imports of clothing in 2008. Whereas, the import share from Sweden was relatively stable around 5 % throughout the sample period, the import shares from United Kingdom and Denmark dropped by nearly 10 percentage points each during the period 1995–2008. Hong Kong also experienced a lower import share by 5 percentage points

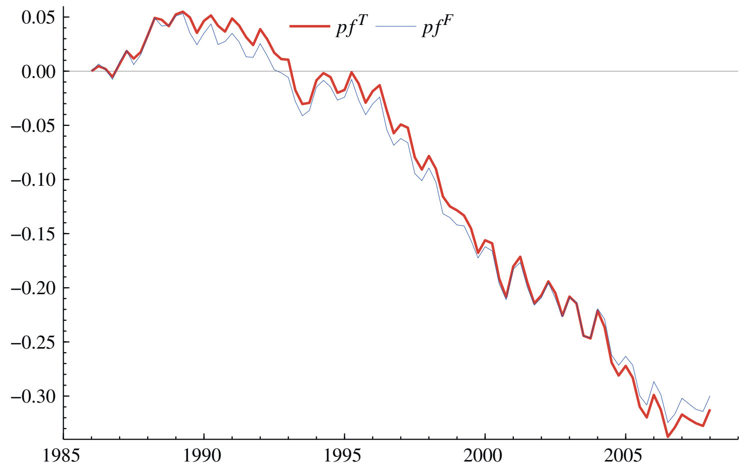


**Fig. 4** Time series for foreign prices ($pf_t^T$ and $pf_t^F$)

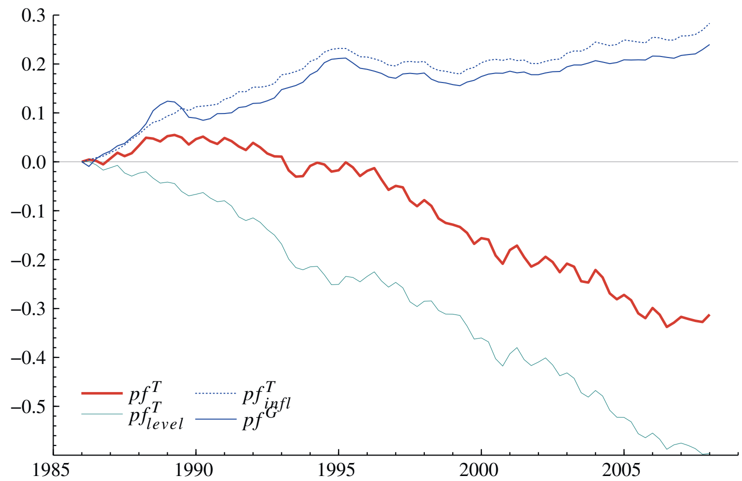


**Fig. 5** Time series for foreign prices ($pf_t^T$, $pf_{t,infl}^T$, $pf_{t,level}^T$ and $pf_t^G$)

during the same period. Overall, the shift in imports towards low-cost countries at the expense of high-cost countries was evident since the mid 1980s, but was intensified from the early 1990s and even more from around 1995 alongside the removal of the quota restrictions on trade.

Figure 4 displays the computed Törnqvist ($pf_t^T$) and Fischer ($pf_t^F$) price aggregates (measured in foreign currency) based on (8) and (9), respectively. Practically speaking, we see that the two superlative price indices generate identical aggregates with a substantial fall in international export prices on clothing during the sample period. Our calculations indicate that the price aggregates are roughly 30 % lower in 2008 compared to 1986, which implies on average a yearly decrease of around 1.3 percentage points.

Figure 5 displays the Törnqvist price index measure of foreign prices ($pf_t^T$) and its two components, the inflation effects ($pf_{t,\inf l}^T$) and the price level effects ($pf_{t,level}^T$) based on (11), together with the geometric mean price index measure of foreign prices with constant weights ($pf_t^G$) based on (12). According to our calculations, the shift in imports from high- to low-cost countries—the China effect – has on average pushed down international price impulses by around 2 percentage points each year since the early 1990s. During the second half of the 1980s, the price level effects were moderate, reflecting little substitution of imports towards low-cost countries due to strict trade regulations. The international price impulses were, however, pulled higher and dominated by inflationary effects up until 1995, before these effects became moderate and even negative in the late 1990s. Paralleling the period of trade liberalisation, the price level effects played a dominating role in the development of $pf_t^T$ from 1995 onwards. Even though the last quota restriction was lifted in 1998, the price level effects continued to pull down $pf_t^T$ during the last decade, which indicates that trade liberalisation may have had long lasting effects on international export prices on clothing. We also observe that the development in $pf_t^G$ parallels the development in $pf_{t,\inf l}^T$. More importantly though, is the substantial differences in $pf_t^T$ and $pf_t^G$. Because the latter fails to take account of the differences in price levels across trading partners, it exhibits an overall international price increase of somewhat less than 30 % through the sample period rather than a price fall of the same magnitude. We believe that $pf_t^T$ provides a better measure of the true international price development faced by Norwegian importers of clothing given the significant change in the import pattern over time.

Figure 6 displays the Törnqvist price aggregate ($pf_t^T$) based on (8) together with the Törnqvist price aggregate based on a 50 % increase ($pf_{t,high}^T$) and decrease ($pf_{t,low}^T$) in the relative price levels for China and Turkey in Table 1. We notice that the development in $pf_t^T$ is rather sensitive to different assumptions made about the relative price levels for China and Turkey. Whereas a 50 % increase in the relative price levels for China and Turkey makes an international price fall of only 5 % ($pf_{t,high}^T$) from

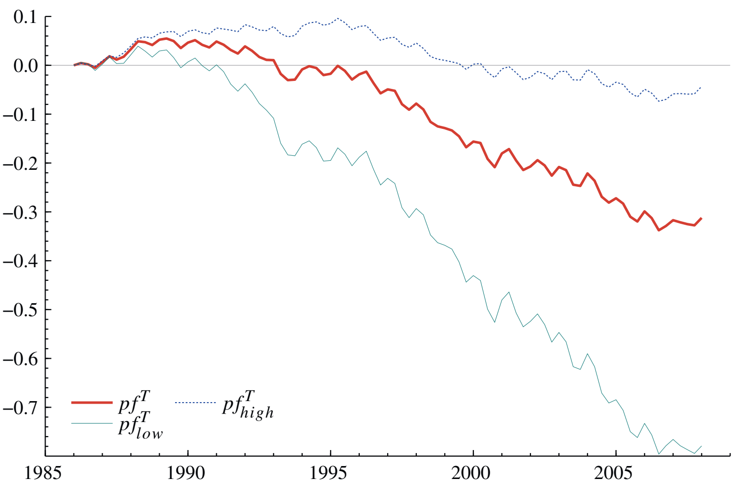


**Fig. 6** Time series for foreign prices ($pf_t^T$, $pf_{t,high}^T$ and $pf_{t,low}^T$)

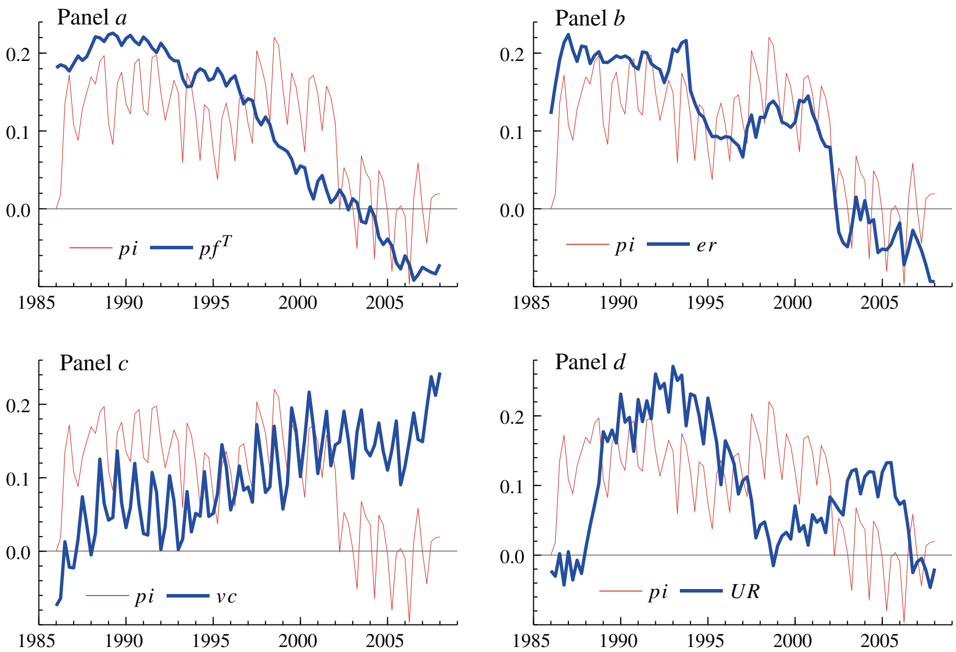


**Fig. 7** Time series for $pi_t$, $pf_t^T$, $er_t$, $vc_t$ and $UR_t$. Sources: See the Appendix

1986 to 2008, a 50 % decrease in the same relative price levels produces a price fall of as much as 55 % ($pf^T_{t,low}$) in the same period. We return to this issue below and analyse whether the sensitivity in $pf_t^T$ produces a serious sensitivity in the estimates of pass-through and pricing-to-market.

Figure 7 displays the time series for the import price of clothing ($pi_t$) together with the Törnqvist price index measure of foreign prices ($pf_t^T$) in panel *a*, the exchange rate ($er_t$) in panel *b*, the domestic variable unit costs ($vc_t$) in panel *c*, and the unemployment rate ($UR_t$) in panel *d*. The exchange rate series is a chained geometric mean index the construction of which parallels that of $pf_t^T$ in the sense that the bilateral exchange rates between Norway and the seven trading partners are weighted together with their respective (variable) import shares as weights. Domestic variable unit costs are defined as the sum of costs of variable factor inputs relative to total production of clothing and the unemployment rate is measured as the number of unemployed as a fraction of the total labour force (according to the Labour Force Survey). The scale of $pf_t^T$, $er_t$, $vc_t$ and $UR_t$ are adjusted in Fig. 7 to match that of $pi_t$, which is normalised to unity in 1986$Q$1.

It is evident that $pi_t$, $pf_t^T$ and $er_t$ all exhibit a clear downward trend throughout the sample period, whereas $vc_t$ shows some upward trend. At the same time, import prices of clothing relative to foreign prices measured in Norwegian currency $(pi - pf^T - er)_t$ increased from 1986 to 2008, which may be explained by the fact that variable unit costs relative to import prices $(vc - pi)_t$ also increased in the same period. Although consistent with the pricing-to-market hypothesis, this cannot be the full explanation for the development in $pi_t$ as $(pi - pf^T - er)_t$ increased somewhat more than $(vc - pi)_t$. As indicated by panel *d*, the development in $pi_t$ may also partly be explained by the development in the domestic demand pressure ($UR_t$). Specifically, the apparent fall in $pi_t$ during the first half of the 1990s and during the years between 1999 and 2006 coincides well with increased $UR_t$ in the same periods. Likewise, the increase in $pi_t$ during the second half of the 1990s matches rather closely with decreased $UR_t$. That the two price series, the exchange rate series and the series for variable unit costs, exhibit some trending behaviour with no apparent mean reversion points to non-stationary time series properties. The unemployment rate, on the other hand, is stationary by construction. However, we follow Bjørnstad and Nymoen (1999) in the subsequent analysis and treat $UR_t$ as if it is non-stationary within the sample period.

## 4 The econometric procedure

Because the pricing-to-market theory predicts the possibility of multiple cointegrating vectors among the variables involved, we employ the Johansen (1995, p. 167) trace test for cointegration rank determination. We start with an unrestricted $p$-dimensional VAR of order $k$ having the form

$$X_t = \sum_{i=1}^{k} \Pi_i X_{t-i} + \mu + \varpi t + \varepsilon_t, \quad t = k+1, \ldots, T, \tag{14}$$

where $X_t$ is a $(p \times 1)$ vector of modelled variables at time $t$, $\mu$ represents constants and seasonals, $\varpi$ is a $(p \times 1)$ coefficient vector of a linear deterministic trend $t$, $\Pi_1, \ldots, \Pi_k$ are $(p \times p)$ coefficient matrices of lagged level variables and $\varepsilon_{k+1}, \ldots, \varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) $(p \times p)$ covariance matrix $\Omega$. The initial observations of $X_1, \ldots, X_k$ are kept fixed. The question now is how (14) can be re-parameterised to a cointegrated VAR (henceforth CVAR) in which the pricing-to-market hypothesis can be formulated as a reduced rank restriction on the impact matrix $\Pi = -(I - \Pi_1 - \cdots - \Pi_k)$.

The way the CVAR is formulated in our context depends on the exogeneity status or otherwise of the unemployment rate series. Firstly, we consider the case when the unemployment rate series is endogenous in the system, hence (14) is a five-dimensional VAR in $X_t = (pi_t, pf_t^i, er_t, vc_t, UR_t)'$, $i = T, F, G$. Once $X_t \sim I(1)$, the first difference $\Delta X_t \sim I(0)$ implying either $\Pi = 0$ or $\Pi$ has reduced rank such that $\Pi = \alpha\beta'$, where $\alpha$ and $\beta$ are $5 \times r$ matrices and $0 < r < 5$. Herein $r$ denotes the rank order of $\Pi$. Assuming for notational simplicity that $k = 2$, the CVAR in this situation becomes

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \alpha\beta' X_{t-1} + \mu + \delta t + \varepsilon_t, \tag{15}$$

where $\beta' X_{t-1}$ is an $r \times 1$ vector of stationary cointegration relations among import prices, foreign prices, exchange rates, variable unit costs and the unemployment rate, and $\Gamma_1 = -\Pi_2$ is a $(5 \times 5)$ coefficient matrix of the lagged differentiated variables. Next, we consider the case when the unemployment rate series is weakly exogenous for the long run parameters such that valid inference on $\beta$ can be obtained from the four-dimensional system describing $pi_t$, $pf_t^i$, $er_t$ and $vc_t$ conditional on $UR_t$ without loss of information, see Johansen (1992). Following Harbo et al. (1998), we may formulate the *partial* CVAR equivalent to (15) as (again assuming $k = 2$)

$$\Delta X_{1,t} = A_1 \Delta X_{2,t} + \Gamma_{1,1} \Delta X_{t-1} + \alpha_1 \beta' X_{t-1} + \mu_1 + \delta_1 t + \varepsilon_{1,t}, \tag{16}$$

with the corresponding marginal model given by $\Delta X_{2,t} = \Gamma_{1,2} \Delta X_{t-1} + \mu_2 + \delta_2 t + \varepsilon_{2,t}$ when $X_t = (X'_{1,t}, X_{2,t})'$, $X_{1,t} = (pi_t, pf_t^i, er_t, vc_t)'$ and $X_{2,t} = UR_t$. It follows that the unemployment rate is included in the long-run part of (16) as a non-modelled variable. Because the number of relevant variables to be included in (14), and hence also the number of parameters to be estimated, is large relative to the number of observations in the available data set, it would be useful to impose weak exogeneity on the unemployment rate. However, to know whether $\beta$ can be estimated from (16), we first estimate the full system in (15) and test formally rather than assume the weak exogeneity status of the unemployment rate in that system. We follow common practice and let inference about the rank of $\Pi$ from the full system be based on unrestricted intercepts and a restricted linear trend. Likewise, dummies capturing seasonality in the data ($S1_t$, $S2_t$ and $S3_t$) enter the system unrestrictedly.

Strictly speaking, the cointegration rank does not need to be determined from the partial system once it has been determined from the full system. Nevertheless, we re-determine the cointegration rank from (16) for the sake of comparison with the rank determination from (15). However, as noted by Harbo et al. (1998), the asymptotic distribution of the trace test statistic is influenced by conditioning on weakly exogenous variables and standard critical values are thus not valid. We, therefore, use the critical values in Table 2 in Harbo et al. (1998). Also, following the suggestions in Harbo et al. (1998) for partial systems, we restrict the linear trend to lie in the cointegration space for inference purposes only. Then, after having determined the cointegration rank, we test whether the linear trend can be dropped from the cointegration relation(s) by a conventional $\chi^2$ -test. As in the full system, both the constants and the seasonals enter the partial system unrestrictedly.

We now turn to the empirical findings from the cointegration analysis based on the econometric procedure outlined above. Not surprisingly, we find that the two calculated superlative price indices produce similar cointegration analyses and near identical corresponding dynamic import price models. Accordingly, we shall below concentrate on the empirical findings based on the Törnqvist price index measure of foreign prices.[7]

## 4.1 Cointegration analysis based on $PF_t^T$

Irrespective of specifying a full five-dimensional VAR in $X_t = (pi_t, pf_t^T, er_t, vc_t, UR_t)'$ or a partial four-dimensional VAR in $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ conditional on $X_{2,t} = UR_t$ being exogenous to the system, we find that $k = 3$ produces a model with

[7] Results from the cointegration analysis and the modelling of the dynamic import price equation based on the Fischer price index measure of foreign prices are available from the authors upon request.

**Table 2** Tests for cointegration rank based on $PF_t^T$

| Full CVAR system | | | | Partial CVAR system | | | |
|---|---|---|---|---|---|---|---|
| $r$ | $\lambda_i$ | $\lambda_{trace}$ | $p$-$\text{val}_{Ox}$ | $r$ | $\lambda_i$ | $\lambda_{trace}$ | $5\,\%_{Harbo}$ |
| $r = 0$ | 0.41 | 107.66 | 0.001 | $r = 0$ | 0.40 | 94.57 | 71.7 |
| $r \leq 1$ | 0.28 | 62.90 | 0.058 | $r \leq 1$ | 0.24 | 50.72 | 49.6 |
| $r \leq 2$ | 0.18 | 34.87 | 0.254 | $r \leq 2$ | 0.18 | 26.99 | 30.5 |
| $r \leq 3$ | 0.12 | 18.00 | 0.351 | $r \leq 3$ | 0.11 | 10.36 | 15.2 |
| $r \leq 4$ | 0.08 | 6.82 | 0.374 | | | | |

*Notes*: Sample period: 1986$Q$4–2008$Q$1. The underlying VARs are of order 3. The full CVAR consists of $X_t = (pi_t, pf_t^T, er_t, vc_t, UR_t)'$, whereas the partial CVAR consists of $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ being endogenous and $X_{2,t} = UR_t$ being exogenous. Both systems include unrestricted constants and seasonals and a restricted linear trend. $r$ denotes the cointegration rank, $\lambda_i$ are the eigenvalues from the reduced rank regressions, $\lambda_{trace}$ are the trace test statistics, $p$-$\text{val}_{Ox}$ are the significance probabilities from OxMetrics and $5\,\%_{Harbo}$ are the critical values (5 % significance level) from Table 2 in Harbo et al. (1998)

no serious misspecification as indicated by standard diagnostic tests. Certainly, the estimated residuals of the $UR_t$-equation in the full system and thus also the estimated vector residuals are borderline cases (at conventional significance levels) with respect to suffering from autocorrelation. Such a potential problem may in itself be an argument for moving to a partial system to obtain even more satisfying residual properties in our case, see Juselius (2006, p. 198). Noticeably, no impulse dummies are required to mop up any outliers to obtain Gaussian residuals.[8] Table 2 reports trace test statistics for the sample period 1986Q4–2008Q1, both in the case of the full system and the partial system with the Törnqvist price index measure of foreign prices assuming $k = 3$.

We notice that the null hypothesis of no cointegration can be rejected at the 5 % significance level, whereas the hypothesis of at most one cointegrating relationship between import prices, foreign prices, exchange rates, domestic variable unit costs and demand pressure (proxied by the unemployment rate) cannot be rejected within the full CVAR. As shown below, choosing $r = 1$ gives a cointegrating vector with interpretable properties in line with the pricing-to-market hypothesis. Testing a zero restriction on the equilibrium correction coefficient of the unemployment rate under the assumption of $r = 1$, gives $\chi^2_{(1)} = 0.91$ with a $p$-value of 0.34. Hence, $UR_t$ may be considered as weakly exogenous for the cointegrating parameters the estimates of which can then be efficiently estimated from the partial rather than the full system without loss of information. In so doing, we also obtain a more parsimonious, feasible VAR and save degrees of freedom. The formal tests in Table 2 support the hypothesis that $r = 1$ also in the case of the partial CVAR, albeit a borderline case at the 5 % significance level. Likelihood ratio tests (not shown) clearly reject the hypothesis that the modelled variables in $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ as well as $X_{2,t} = UR_t$ are excluded

[8] A VAR of order 2 produces severe autocorrelation in the vector residuals and in the residuals of the $pf_t^T$-equation and the $UR_t$-equation of the full system. Results from the diagnostic tests of the VARs and other test results not reported, here and below, are available from the authors upon request. As noted by Franses and Lucas (1998), standard cointegration tests are sensitive to atypical events such as outliers and structural breaks.

**Table 3** Tests of the pricing-to-market hypothesis based on $PF_t^T$

| Hypothesis | LR tests | $p$-value |
|---|---|---|
| $H_1$: $\alpha_{1(pi_t)} = 0$ | $\chi^2_{(2)} = 20.34$ | 0.000 |
| $H_2$: $\alpha_{1(pf_t^T)} = 0$ | $\chi^2_{(2)} = 18.14$ | 0.001 |
| $H_3$: $\alpha_{1(er_t)} = 0$ | $\chi^2_{(2)} = 2.08$ | 0.354 |
| $H_4$: $\alpha_{1(vc_t)} = 0$ | $\chi^2_{(2)} = 0.53$ | 0.766 |
| $H_5$: $(pi_t - \psi_1 pf_t^T - \psi_1 er_t - \psi_2 vc_t) \sim I(0)$ | $\chi^2_{(2)} = 2.01$ | 0.367 |
| $H_6$: $[pi_t - (1-\psi)(pf_t^T + er_t) - \psi vc_t] \sim I(0)$ | $\chi^2_{(3)} = 2.07$ | 0.558 |
| $H_7$: $(pi_t - pf_t^T - er_t) \sim I(0), \beta_{(vc_t)} = 0$ | $\chi^2_{(4)} = 29.81$ | 0.000 |
| $H_8$: $(vc_t - pf_t^T - er_t) \sim I(0), \beta_{(pi_t)} = 0$ | $\chi^2_{(4)} = 33.71$ | 0.000 |

*Notes*: Sample period: 1986$Q$4–2008$Q$1. All likelihood ratio (LR) tests are based on the partial CVAR with $r = 1$ and $\beta_{(trend)} = 0$ and with degrees of freedom in parenthesis

from $\beta$. The linear trend, however, is clearly *in*significant with $\chi^2_{(1)} = 0.523$ and a $p$-value of 0.47. It is therefore excluded from the model in the following likelihood ratio tests about the pricing-to-market hypothesis, that is, tests about $\alpha_1$ and $\beta$ in (16) assuming $r = 1$. Table 3 summarises results from these tests.

Firstly, we observe that weak exogeneity of both import prices and foreign prices for the long run parameters is strongly rejected. By way of contrast, we may assume that exchange rates and domestic costs are weakly exogenous. The hypotheses of identical parameters of foreign prices and exchange rates ($H_5$) and of long run homogeneity as an additional restriction ($H_6$) are both accepted by the data. On the other hand, the hypotheses of long run versions of LOP ($H_7$) and PPP ($H_8$), as defined in Sect. 3, are clearly rejected by the data. Finally, imposing equal parameters of $pf_t^T$ and $er_t$, long run homogeneity and weak exogeneity of $er_t$ and $vc_t$ yields $\chi^2_{(5)} = 6.21$ with a $p$-value of 0.29. Hence, we obtain the following restricted cointegrating vector (normalised on import prices)

$$pi_t = const. + 0.444 pf_t^T + \underset{(0.016)}{0.444} er_t + 0.556 vc_t - \underset{(0.003)}{0.020} UR_t, \tag{17}$$

with standard errors in parentheses. The associated vector of equilibrium correction coefficients is estimated to $\hat{\alpha}_1 = (-0.44, -0.18, 0, 0)'$. Because any deviations from (17), due to say a shock in the exchange rate, are mainly and significantly corrected through the adjustment of import prices we regard the estimated cointegrating vector as a long run import price equation for clothing consistent with the pricing-to-market hypothesis.[9] The pass-through and pricing-to-market elasticities are significantly estimated to 0.44 and 0.56, respectively. Also, the estimated import price equation includes significant negative effects of the unemployment rate such that decreases in domestic demand pressure (proxied by increases in the unemployment rate) cause prices of imports to fall somewhat.

[9] Although significantly estimated, the adjustment coefficient for $pf_t^T$ is only 40 % of that for $pi_t$.

Interestingly, Naug and Nymoen (1996) found the pass-through elasticity to be 0.63 based on data for Norwegian imports of total manufactures over the sample period 1970Q1–1991Q4. As price setting behaviour typically varies across products and the presence of non-tariff barriers to trade is not controlled for, the estimate of pass-through in Naug and Nymoen (1996) is likely to be biased. Our estimate of pass-through also differs somewhat from those found by Menon (1996) based on disaggregated Australian data over the sample period 1981Q3–1992Q2. In that study, the estimates in most cases indicate incomplete pass-through, but with substantial variation across products. Particularly, pass-through is estimated to be less than 30 % for some of the quota-protected textiles and wearing apparels studied. Menon (1996) partly views this finding in light of the Bhagwati hypothesis as significant negative effects from a quantity restriction variable are among the most convincing results. That is, exchange rate changes have to some extent been prevented from being fully passed through to import prices by the import premium associated with quotas in the Australian context. Our estimate of pass-through may also be viewed in light of the Bhagwati hypothesis. As pointed out above, the hypothesis implies increased pass-through when non-tariff barriers to trade are gradually removed, other things equal. However, once the China effect is included in the measure of foreign prices, it is likely that pass-through has not changed dramatically since the mid 1990s. Recursive estimates of the pass-through coefficient in (17) are reasonably stable in the years after 1995. Also, recursively estimated $\chi^2_{(5)}$ indicate that the restrictions in (17) are supported by the data throughout the second half of the sample period.

We complete the cointegration analysis based on $pf_t^T$ by examining potential sensitivity in the estimate of pass-through based on different assumptions made about the relative price levels for China and Turkey. As already revealed from Fig. 6, the calculated development in $pf_t^T$ is somewhat sensitive to a 50 % increase ($pf_{t,high}^T$) and decrease ($pf_{t,low}^T$) in the relative price levels for China and Turkey in Table 1. We obtain the following estimated cointegrating vectors with $pf_{t,high}^T$ and $pf_{t,low}^T$ replacing $pf^T$, all other modelling issues equal:

$$pi_t = const. + 0.604 pf_{t,high}^T + \underset{(0.019)}{0.604} er_t + 0.396 vc_t - \underset{(0.003)}{0.019} UR_t, \tag{18}$$

$$pi_t = const. + 0.306 pf_{t,low}^T + \underset{(0.015)}{0.306} er_t + 0.694 vc_t - \underset{(0.005)}{0.023} UR_t. \tag{19}$$

Similar to (17), we have imposed equal parameters of $pf_{t,i}^T$ ($i = high, low$) and $er_t$, long run homogeneity and weak exogeneity of $er_t$ and $vc_t$ in (18) and (19), which yields $\chi^2_{(5)} = 1.55$ and $\chi^2_{(5)} = 9.01$ with $p$-values of 0.91 and 0.11, respectively. We observe that the estimates of pass-through, and hence also the estimates of pricing-to-market, do not depend critically on the assumptions made about the relative price levels for China and Turkey. The estimate of pass-through increases and decreases by 33 % when $pf_{t,high}^T$ and $pf_{t,low}^T$ replace $pf_t^T$, which we consider as a rather moderate sensitivity in the estimate given the rather substantial magnitude of the shift in the relative price levels. We shed some further light on the sensitivity in the estimates of pass-through and pricing-to-market due to the potential problem of omitted variable bias in the subsequent cointegration analysis based on $pf_t^G$ rather than $pf_t^T$.

**Table 4** Tests for cointegration rank based on $PF_t^G$

| Full CVAR system | | | | Partial CVAR system | | | |
|---|---|---|---|---|---|---|---|
| $r$ | $\lambda_i$ | $\lambda_{trace}$ | $p$-val$_{Ox}$ | $r$ | $\lambda_i$ | $\lambda_{trace}$ | $5\,\%_{Harbo}$ |
| $r = 0$ | 0.32 | 89.51 | 0.042 | $r = 0$ | 0.31 | 72.14 | 71.7 |
| $r \leq 1$ | 0.23 | 56.95 | 0.166 | $r \leq 1$ | 0.18 | 39.58 | 49.6 |
| $r \leq 2$ | 0.15 | 34.60 | 0.266 | $r \leq 2$ | 0.15 | 22.88 | 30.5 |
| $r \leq 3$ | 0.15 | 20.24 | 0.218 | $r \leq 3$ | 0.10 | 8.67 | 15.2 |
| $r \leq 4$ | 0.07 | 6.06 | 0.464 | | | | |

*Notes*: Sample period: 1986$Q$4–2008$Q$1. The underlying VARs are of order 3. The full CVAR consists of $X_t = (pi_t, pf_t^G, er_t, vc_t, UR_t)'$, whereas the partial CVAR consists of $X_{1,t} = (pi_t, pf_t^G, er_t, vc_t)'$ being endogenous and $X_{2,t} = UR_t$ being exogenous. Both systems include unrestricted constants and seasonals and a restricted linear trend. $r$ denotes the cointegration rank, $\lambda_i$ are the eigenvalues from the reduced rank regressions, $\lambda_{trace}$ are the trace test statistics, $p$-val$_{Ox}$ are the significance probabilities from OxMetrics and $5\,\%_{Harbo}$ are the critical values (5 % significance level) from Table 2 in Harbo et al. (1998)

### 4.2 Cointegration analysis based on $PF_t^G$

As with the Törnqvist price index measure of foreign prices, a lag length of three is sufficient to render residuals with no serious misspecification, neither in the full nor in the partial VAR. Again, no impulse dummies are needed to achieve Gaussian residuals in the VARs. Table 4 reports trace test statistics based on the VARs of order three when the measure of foreign prices are based on the geometric mean price index with constant weights.

Again, the rank should be set to unity in the case of the full CVAR system at the 5 % significance level. Also, the unemployment rate is weakly exogenous for the long run parameters in that system under the assumption of $r = 1$, as indicated by $\chi^2_{(1)} = 0.002$ with a $p$-value of 0.97. Accordingly, we may conduct inference about the $\alpha$ and $\beta$ matrices relying on the partial CVAR. The formal tests in Table 4 also indicate existence of a unique cointegration relationship with the partial system. Besides, the hypothesis that a specific variable does not enter the cointegrating relation is rejected for all the variables $pi_t$, $pf_t^G$, $er_t$, $vc_t$ and $UR_t$, a finding in line with the analysis above using the Törnqvist price index measure of foreign prices. However, the linear trend is now needed in the cointegration space and cannot be omitted from the long run relation according to $\chi^2_{(1)} = 8.12$ and its $p$-value of 0.004. Consequently, it is *not* excluded from the reduced rank partial VAR underlying the tests about the pricing-to-market hypothesis reported in Table 5.

Overall, the test results in Table 5 are similar to those in Table 3. Note that $H_2$ is now not rejected by the data, indicating that $pf_t^G$ just like $er_t$ and $vc_t$ is exogenous for the parameters of interest. Hence, imposing the restrictions corresponding to the hypotheses $H_2$, $H_3$, $H_4$ and $H_6$ gives $\chi^2_{(5)} = 4.28$ and a $p$-value of 0.51, and the following restricted cointegrating vector (normalised on import prices)

$$pi_t = const. + 0.601 pf_t^G + \underset{(0.101)}{0.601} er_t + 0.399 vc_t - \underset{(0.005)}{0.020} UR_t - \underset{(0.00047)}{0.00207} t, \quad (20)$$

**Table 5** Tests of the pricing-to-market hypothesis based on $PF_t^G$

| Hypothesis | LR tests | $p$-value |
|---|---|---|
| $H_1: \alpha_{1(pi_t)} = 0$ | $\chi^2_{(1)} = 11.79$ | 0.001 |
| $H_2: \alpha_{1(pf_t^G)} = 0$ | $\chi^2_{(1)} = 2.48$ | 0.115 |
| $H_3: \alpha_{1(er_t)} = 0$ | $\chi^2_{(1)} = 0.09$ | 0.766 |
| $H_4: \alpha_{1(vc_t)} = 0$ | $\chi^2_{(1)} = 0.47$ | 0.492 |
| $H_5: (pi_t - \psi_1 pf_t^G - \psi_1 er_t - \psi_2 vc_t) \sim I(0)$ | $\chi^2_{(1)} = 1.34$ | 0.246 |
| $H_6: [pi_t - (1-\psi)(pf_t^G + er_t) - \psi vc_t] \sim I(0)$ | $\chi^2_{(2)} = 2.14$ | 0.343 |
| $H_7: (pi_t - pf_t^G - er_t) \sim I(0), \beta_{(vc_t)} = 0$ | $\chi^2_{(3)} = 11.25$ | 0.011 |
| $H_8: (vc_t - pf_t^G - er_t) \sim I(0), \beta_{(pi_t)} = 0$ | $\chi^2_{(3)} = 18.15$ | 0.000 |

*Notes*: Sample period: 1986$Q$4–2008$Q$1. All likelihood ratio (LR) tests are based on the partial CVAR with $r = 1$ and with degrees of freedom in parenthesis

with standard errors in parenthesis. The adjustment coefficient of import prices is now significantly estimated to $-0.42$, which is almost identical to the corresponding estimate obtained with the Törnqvist price index measure of foreign prices. More important though, when comparing (17) and (20), are the somewhat different estimates of long run pass-through and pricing-to-market that come out of the modelling with the two alternative measures of foreign prices. Another important difference between the two estimated cointegrating vectors is the linear trend, which enters significantly in (20) and *not* in (17).

One possible interpretation of these findings is that the effects of shifts in imports from high- to low-cost countries on internationally traded good prices (and thereby on the degree of pass-through) are likely to be controlled for through the linear trend in (20), effects which are *not* explicitly picked up by $pf_t^G$ alone. As seen from Fig. 5, the calculated price level term of $pf_t^T$ (the China effect) drifts downwards during the entire sample period and may accordingly behave like a deterministic linear trend in a regression model. Indeed, the linear trend enters significantly in (20) with a negative sign consistent with the a priori beliefs about the China effect on internationally traded goods prices. The estimate implies that the shift in imports towards low-cost countries has depressed import prices of clothing by around 0.8 percentage points yearly ($-0.00207 \cdot 400$) since 1986, approximately equal to the yearly average of around 0.9 calculated by means of the pass-through estimate of 0.44 from (17) and the 2 percentage points yearly decrease in the price level term $pf_{t,level}^T$. However, the fact that $pf_{t,level}^T$ exhibits some apparent cycles, especially around 1995 and 2000, may make a linear trend a poor proxy for the true China effect on international price impulses faced by Norwegian importers as such. For this reason, we suspect the estimates of pass-through and pricing-to-market in (20) to be somewhat biased compared to those in (17). We also find that pass-through is more or less complete and thus that pricing-to-market behaviour is absent when the trend variable is dropped from (20), results which are unlikely given the facts about the clothing industry outlined in Sect. 2. Although not accepted by the data[10], these findings also point to the likely problem of an omitted

10 The $\chi^2_{(6)} = 20.51$ with a $p$-value of 0.002.

variable bias if shifts in imports towards low-cost countries and trade liberalisation effects are *not* explicitly controlled for in the import price equation for clothing.

We have seen that using the Törnqvist price index measure of foreign prices is a flexible approach that may overcome this potential econometric problem in our empirical case, all other modelling issues equal. Based on our findings, we also believe it is a more reliable approach than using the geometric mean price index with constant weights together with a linear trend (which at best represents the China effect) in the regression model to quantify pass-through consistently.

## 5 A dynamic import price model

The degree of pass-through may, just like trade policy, be linked to the nature and magnitude of exchange rate changes. According to Froot and Klemperer (1989), foreign firms are likely to price more aggressively in the domestic market to gain higher market shares when the currency of the importing country is expected to be permanently stronger. Conversely, when a currency appreciation is believed to be temporary, foreign firms will behave less aggressively in their price setting. We shall here test the hypothesis that pass-through has changed alongside trade liberalisation and particular exchange rate fluctuations by examining stability properties of an estimated dynamic equilibrium correction model (henceforth $EqCM$). Our point of departure is a general $EqCM$ model (with the constant, the seasonals and the same lag length used in the reduced rank partial VAR) written as

$$\begin{aligned}\Delta pi_t &= const. + \sum_{i=1}^{2}\varphi_{1,i}\Delta pi_{t-i} + \sum_{i=0}^{2}\varphi_{2,i}\Delta(pf^T + er)_{t-i} + \sum_{i=0}^{2}\varphi_{3,i}\Delta vc_{t-i} \\ &\quad + \sum_{i=0}^{2}\varphi_{4,i}\Delta UR_{t-i} + \eta EqCM_{t-1} + \eta_1 S1_t + \eta_2 S2_t + \eta_3 S3_t + e_t. \qquad (21)\end{aligned}$$

The general model contains impact effects and two lags of the first difference (denoted $\Delta$) of $vc_t$, the sum of $pf_t^T$ and $er_t$, and $UR_t$. We notice that $\Delta(pf^T + er)_t$ is denominated in Norwegian currency, a short run restriction imposed from the outset in line with the corresponding long run restriction (i.e. equal parameters of $pf_t^T$ and $er_t$) accepted by the data. Also, the first difference of $pi_t$ is included in (21) with two lags, whereas the $EqCM$ term [defined in accordance with (17)] is lagged one period. The error term $e_t$ is assumed to be white noise. Simplifications from the general to the specific model is performed using the autometrics option in OxMetrics 6, see Doornik and Hendry (2009). Autometrics picks the following *specific* model in our case together with diagnostic tests[11] and the estimated standard errors below the point estimates (sample period: 1986Q4–2008Q1):

[11] $AR_{1-5}$ is Harvey (1981) test for until 5th order residual autocorrelation; $ARCH_{1-4}$ is the Engle (1982) test for until 4th order autoregressive conditional heteroskedasticity in the residuals; *NORM* is the normality test outlined in Doornik and Hansen (2008), *HET* is a test for residual heteroskedasticity due to White (1980) and *RESET* is the Ramsey (1969) test for functional form misspecification. The numbers in square brackets are $p$-values.

$$\Delta pi_t = \underset{(0.055)}{-0.336}\Delta pi_{t-1} + \underset{(0.088)}{0.479}\Delta(pf^T + er)_t - \underset{(0.006)}{0.021}\Delta UR_t - \underset{(0.006)}{0.020}\Delta UR_{t-1}$$
$$\underset{(0.067)}{-0.385}[pi_{t-1} - 0.44(pf^T + er)_{t-1} - 0.56vc_{t-3} + 0.020UR_{t-2}]$$
$$\underset{(0.024)}{+0.188} - \underset{(0.009)}{0.065}S1_t - \underset{(0.007)}{0.109}S2_t \tag{22}$$

Diagnostic tests:
$AR_{1-5} : F(5, 73) = 1.58\ [0.18], ARCH_{1-4} : F(4, 70) = 1.82\ [0.14],$
$NORM : \chi^2_{(2)} = 4.97\ [0.08], HET : F(12, 65) = 1.59\ [0.12],$
$RESET : F(1, 77) = 1.98\ [0.16].$

The equilibrium correction term enters significantly in (22). The estimated coefficient of $-0.39$ implies rapid adjustment of import prices of clothing towards the long run equilibrium level in the event of a shock in either foreign prices, exchange rates, domestic costs or demand pressure. The *EqCM* is specified with three and two lags on domestic costs and demand pressure, respectively, a reparameterisation that turned out useful to obtain reasonable short run dynamic properties. The estimated short run pass-through elasticity is somewhat greater than its long run counterpart.[12] Accordingly, import prices respond quickly and with some overshooting to shocks in foreign prices (denominated in foreign currency) and exchange rates. However, the specific model also contains significant and negative short run autoregressive effects from $\Delta pi_{t-1}$, which make the adjustment process of import prices somewhat less smooth. That is, the first quarter adjustment of import prices following a shock in say the exchange rate is corrected somewhat in the opposite direction in the next quarter due to the autoregressive effects before the adjustment process continues steadily towards the long run equilibrium level. Altogether, pass-through is almost complete within one to three quarters according to (22). The rather fast speed of adjustment of import prices identified here may reflect the fact that the exchange rate was fairly volatile during most of the sample period (cf. Fig. 7, panel *b*). If there are costs related to changing import prices, it will be rational to respond relatively fast to large fluctuations in the exchange rate that are *not* likely to be reversed in the near future. We notice further from (22) that $\Delta UR_t$ and $\Delta UR_{t-1}$ enter the model with more or less identical effects on import prices, effects which also are almost identical to the long run counterpart. Hence, foreign firms seem to absorb quickly, but with some smoothing, into their markups changes in the unemployment rate, which are normally of a somewhat permanent nature.

Our estimate of the speed of adjustment of import prices following a shift in the exchange rate accords with Menon (1996), who finds that pass-through is complete within two quarters for most products in the Australian context. However, Naug and

[12] Both $\Delta pf_t^T$ and $\Delta er_t$ enter *in*significantly as separate explanatory variables in (22), a finding which supports the hypothesis of equal short run impact effects on import prices from changes in these variables. Moreover, the *residuals* from the equations for $\Delta pf_t^T$ and $\Delta er_t$ in the partial VAR are not significant when added to (22). Hence, $\Delta pf_t^T$ and $\Delta er_t$ may be regarded as weakly exogenous for the short run parameters in the specific model. Thus, the parameters are consistently estimated by OLS, see Urbain (1992).

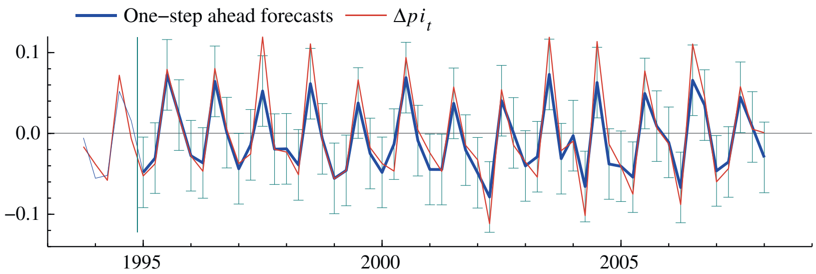


**Fig. 8** Actual values and one-step ahead forecasts of $\Delta pi_t$

Nymoen (1996) find relatively slow speed of adjustment of import prices, which may partly be viewed in light of a sample period where monetary policy was that of a fixed exchange rate regime. Small exchange rate fluctuations during that period (cf. Fig. 7, panel *b*) may thus have been viewed as transitory by foreign firms, in which case it may have been rational to respond slower, if at all.

Turning to parameter stability properties of the specific model, we first notice that the model shows no sign of misspecification as reported below (22). This model property is further confirmed by recursive break point Chow statistics and recursively estimated coefficients, which provide evidence of reasonable constancy from the early 1990s. We now ask whether the model is able to predict import prices of clothing out-of-sample to shed some more light on its robustness with respect to trade policy changes and exchange rate fluctuations during the sample period. If pass-through has changed, we should expect instabilities in the estimated model as indicated by poor out-of-sample forecasting ability. To this end, we shall use simple one-step ahead forecasts by reestimating (22) based on observations until 1994$Q$4, and leaving 53 quarters (1995Q1–2008Q1) for out-of-sample forecasts. Figure 8 depicts actual values of $\Delta pi_t$ together with its one-step ahead forecasts and 95 % confidence intervals to each forecast in the forecasting period (shown by the vertical error bars of $\pm$2SE).

We observe that the forecasts only miss significantly the observed values of $\Delta pi_t$ once, namely in the third quarter of 1997. The point in time of the forecasting failure does coincide with the time period in which a majority of the quota restrictions on trade had already been abolished. However, the particular forecasting failure may well be explained by the Asian financial crises rather than the shift in trade policy itself. As seen from Fig. 7 (panel *b*), the import prices on clothing increased following a sharp depreciation of the Norwegian currency in 1997. Nevertheless, the fact that 15 out of 16 forecasts during the period 1995–1998 are inside the confidence intervals (albeit 1998$Q$3 is a borderline case) points to pass-through being fairly constant throughout the trade liberalisation period. Also, a Chow test statistic of parameter constancy between the sample and the forecasting periods is far from being significant, as indicated by F[53,25] = 0.53 and its $p$-value of 0.98. Moreover, the reestimated model is close to the one in (22) with respect to parameter estimates and diagnostics. We, therefore, conclude that the out-of-sample forecasting ability of the estimated import price model is satisfactory despite shifts in trade policy during the forecasting period. That no serious forecasting failures are detected during the second half of the 1990s may reflect that possible effects on pass-through of changes in trade policy are controlled for through $pf_t^T$, effects which may otherwise be reflected in unstable estimates of the model.

After the introduction of inflation targeting in $2001Q1$, we should expect higher pass-through if foreign exporters believed the exchange rate changes to be more permanent in nature than before. The fact that the model exhibits no forecasting failures suggests that pass-through has remained unchanged also around the date of the shift in monetary policy. These findings may be explained by the fact that foreign firms also experienced relatively high exchange rate volatility during the 1990s, cf. Fig. 7 (panel *b*). After leaving the fixed exchange rate system in 1992 in favour of a managed floating regime, the exchange rate behaved more like free float following several episodes of speculative attacks against the Norwegian currency. It is, therefore, not surprising if foreign firms perceptions of the permanent nature of (large) exchange rate fluctuations changed, if at all, during the period of the managed floating regime. We have established, however, that the estimated import price model is stable also throughout the 1990s, which contradicts such a hypothesis.[13]

## 6 Conclusions

Economic theory predicts that the presence of non-tariff barriers to trade is potentially important when quantifying the degree of pass-through to traded goods prices. In this paper, we applied the cointegrated VAR approach and estimated a pricing-to-market model for Norwegian import prices of clothing over the period 1986–2008, controlling explicitly for potential pass-through effects of the gradual removal of non-tariff barriers to trade and the switch in imports from high- to low-cost countries. The novelty of the paper, we believe, is that the measure of foreign prices is based on superlative price indices (including the Törnqvist and Fischer price indices) and a data calibration method necessary to approximate relative price levels across countries. As such, we allowed not only for inflationary differences as is common in previous pricing-to-market studies, but also varying import shares and differences in price levels (known as the China effect) among trading partners when constructing the measure of foreign prices.

We found that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high- to low-cost countries since the early 1990s on average has reduced the international price impulses on clothing imports by around 2 percentage points per year. With the superlative price index measures of foreign prices, we established import price models for clothing consistent with the pricing-to-market hypothesis. Specifically, we found the pass-through and pricing-to-market elasticities to be 0.44 and 0.56, respectively, irrespective of using the Törnqvist or the Fischer price index measure of foreign prices in the regression model. We also found that these estimates are reasonably stable, which contradicts the implications of the Bhagwati hypothesis that gradual removal of non-tariff barriers to

[13] We also controlled for any instabilities in the estimated model by means of the outlier detection procedure available in OxMetrics 6. It turned out that no significant outliers were detected by this procedure during the periods of trade liberalisation and shift in monetary policy.

trade has pushed pass-through upwards, other things equal. That is, once the China effect is controlled for through the measure of foreign prices, we found little evidence that the long run properties of the import price model have changed significantly alongside trade liberalisation. By way of contrast, we found that the often used geometric mean price index with constant weights overestimates international price impulses and thereby produces biased estimates of pass-through and pricing-to-market. These findings thus point to the potential problem of omitted variable bias in our empirical case if the China effect is *not* explicitly controlled for in the regression model. We may approximate the China effect through a linear trend in the model together with the geometric mean price index with constant weights. However, we showed that such a model is still likely to produce some biasedness in the estimates of pass-through and pricing-to-market. Because the China effect exhibits some apparent cycles, we argue that the superlative price index measures of foreign prices are superior to the geometric price index with constant weights combined with the linear trend, which implicitly assumes that the China effect has been constant throughout the sample period. We further established that the dynamic estimated import price model is reasonably stable in-sample. Finally, a forecasting exercise on the estimated dynamic model does not lend much support to the hypothesis that pass-through has changed in the wake of the trade policy shifts during the second half of the 1990s.

An issue not addressed in this paper is the potential role for expectational dynamics arising from foreign firms being forward-looking in their price setting behaviour. If foreign firms indeed are forward-looking, the coefficients in the regression models considered herein will depend not only on the parameters in the price setting rule, but also on the parameters in the expectations mechanism. Estimating a New Keynesian import price model for clothing by means of likelihood based methods in the spirit of Boug et al. (2006, 2010) is left for future work.

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## Appendix

$PI$: Chained geometric mean price index for imports of clothing (*cif*), measured in Norwegian currency. $1986Q1 = 1$. *Source*: Statistics Norway, the Quarterly National Accounts (QNA).

$PF^T$: Törnqvist price index measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (8) in the text.

$PF^F$: Fischer price index measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (9) in the text.

$PF^G$: Geometric mean price index (with constant weights) measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (12) in the text.

$PF_{ch}$: China: Producer price index of clothing (from $1997Q1$) and consumer price index all products (from $1986Q1$), measured in Chinese currency. *Source*: Reuters EcoWin.

$PF_{eu}$: The Euro area: Producer price index of clothing, measured in EURO. *Source*: Reuters EcoWin.

$PF_{uk}$: United Kingdom: Export price index of clothing, measured in UK currency. *Source*: National statistics online, http://www.statistics.gov.uk/statbase/.

$PF_{sw}$: Sweden: Export price index of clothing, measured in Swedish currency. *Source*: National statistics online, http://www.ssd.scb.se/databaser/.

$PF_{dk}$: Denmark: Industrial output price index of clothing, measured in Danish currency. *Source*: Reuters EcoWin.

$PF_{hk}$: Hong Kong: Producer price index of clothing (from 1990$Q$1) and consumer price index all products (from 1986$Q$1), measured in Hong Kong currency. *Source*: Reuters EcoWin.

$PF_{tr}$: Turkey: Export price index of clothing (from 2004$Q$1), export price index of manufactures (from 1995$Q$1) and import price index total (from 1986$Q$1), measured in Turkish currency. *Source*: Reuters EcoWin.

$s_j$: Value import shares of clothing from country $j$ (China, the Euro area, UK, Sweden, Denmark, Hong Kong and Turkey). *Source*: Statistics Norway, the Foreign Trade Statistics.

$\lambda_j$: Calculated relative price levels for country $j$ based on nominal GDP and PPP adjusted real GDP, cf. Eq. (7) in the text. *Source*: IMF, the World Economic Outlook Database, http://www.imf.org/external/pubs/ft/weo/2009/01/weodata/.

*ER*: Chained geometric mean index for the exchange rate basket based on $s_j$ and the bilateral exchange rates between Norway and China, the Euro area, UK, Sweden, Denmark, Hong Kong and Turkey. $1986Q1 = 1$. *Source*: Statistics Norway and Norges Bank.

*VC*: Domestic variable unit costs of clothing defined as the sum of costs of variable factor inputs relative to total production of clothing. $1986Q1 = 1$. *Source*: Statistics Norway, QNA.

*UR*: Unemployment rate defined as the number of unemployed as a percentage of the labour force. *Source*: Statistics Norway, the Labour Force Survey.

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