OXFORD BULLETIN OF ECONOMICS AND STATISTICS, 77, 5 (2015) 0305–9049
doi: 10.1111/obes.12078

Inflation perceptions and expectations in Sweden – Are media reports the missing link?*

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Abstract

We analyze the interrelation between consumers' quantitative inflation perceptions and expectations as well as actual inflation rates in Sweden. The role of media reports about inflation is emphasized. Structural vector error correction models show stable cointegration between actual, perceived and expected inflation. Impulse responses and forecast error variance decompositions suggest strong interaction between perceived and expected inflation, with a lesser role for actual inflation. Media effects are generally small, but imply an asymmetric reaction of inflation expectations and perceptions to news on increasing vs. decreasing inflation. Thus, to anchor inflation expectations, central banks should explore better communication channels to inform consumers about actual inflation.

I. Introduction

Ever since the rational expectations revolution in macroeconomics, policy makers have emphasized the importance of inflation expectations for monetary policy making. As a consequence, a large literature on the formation of inflation expectations has emerged and many economies have introduced implicit or explicit inflation targets in an attempt to anchor expectations around the target. While most theoretical models assume that agents form expectations rationally based on observed actual inflation rates, the empirical observation of potentially large deviations between actual and perceived inflation raises the question of the relationship between expected and perceived inflation.[1] The role they play in their respective formation process then becomes an important issue for policy makers and theoretical economists alike.

In this paper, we analyze the interrelation between actual inflation and the perceived and expected inflation rates of Swedish households. In order to account for the

*The author would like to thank two anonymous referees, Christopher Carroll, Ulrich Fritsche, Michael Funke, Bernd Lucke, Jan-Oliver Menz, Jan-Egbert Sturm, Artur Tarassow as well as participants at the 2011 Annual Meeting of the German Economic Association and seminar participants at the KOF Swiss Economic Institute, ETH Zurich, and the University of Hamburg for very helpful comments and suggestions and thankfully acknowledges financial support from the German Research Foundation (grant number FR 2677/2-2). All remaining errors are mine.
JEL Classification numbers: C32, E31, E37.

[1] Especially after the Euro introduction, a large gap between actual inflation and households' inflation perceptions was observed in Euro Area countries, see for instance Ehrmann (2006).

cointegration between the inflation variables, we estimate structural vector error correction models (SVECMs) and evaluate long-run cointegration relations as well as responses to structural shocks. Moreover, throughout the analysis special emphasis is given to the role of social amplification via media reports on inflation, which are measured in a unique data set. The study is conducted using monthly quantitative survey data of households' inflation expectations and perceptions from the Swedish Consumer Tendency Survey and media data on reports about inflation in the Svenska Dagbladet from the media research institute MediaTenor.

Overall, our results show considerable interaction between actual, perceived and expected inflation. In particular, inflation expectations of households are more affected by structural shocks to perceived, rather than actual inflation. Media effects are generally found to be small. When concentrating on the effects of 'bad' news on rising inflation vs. 'good' news on falling inflation, we find a stronger media impact. Additionally, there is some evidence of an asymmetric media effect, since potentially 'good' news on decreasing inflation also lead to an increase in perceived and expected inflation and, hence, seem to be perceived as 'bad' news. In that sense, the media do seem to provide a link for the relation between actual, perceived and expected inflation, but may also cause a bias. We thus argue that monetary policy aiming to anchor inflation expectations should also explore other communication channels in order to have a more direct impact on households' inflation perceptions and expectations.

The analysis of the relationship between expectations and perceptions of inflation has received surprisingly little attention in the literature so far. An early contribution, Jonung (1981) reports a significantly positive correlation coefficient between perceived and expected inflation of about 0.5, analyzing the micro data from an older version of the Swedish Consumer Tendency Survey. A similar result is also reported by van der Klaauw *et al.* (2008) for a micro data survey of consumers in the US. By contrast, this paper uses aggregate time series data. To our knowledge, the conceptual framework in Ranyard *et al.* (2008) comes closest to our model of the interrelation between actual, perceived and expected inflation and provides theoretical intuition for the feedback effects. In their model, the authors propose interrelations between actual, perceived and expected inflation, with an additional amplification from the media or social interaction: They hypothesize that inflation perceptions are not only influenced by the direct experience of price changes but also by social amplification via the media or word of mouth. Via agents' spending behavior, inflation perceptions then feed back into actual and expected inflation rates. Inflation expectations, on the other hand, are based on inflation perceptions and economic forecasts and may also be influenced by social amplification. Finally, expectations feed back into actual inflation through saving, spending and investment decisions.

The remainder of the paper is structured as follows: The relevant literature is discussed in section II. The data used are described in section III. Section IV presents results of rationality tests for inflation expectations and perceptions. The nature of interrelations between actual, perceived and expected inflation as well as the role of media reports on inflation is analyzed in SVECM estimations in section V, where we present long-run cointegration relations and short-run media effects, impulse responses to structural shocks, forecast error variance decompositions (FEVDs) and Granger causality tests. Finally, section VI concludes.

II. Literature survey

This paper is related to several approaches in the literature which aim at evaluating the relation between expected and perceived inflation both empirically and theoretically.

Analyzing the formation process of inflation perceptions and expectations separately, rationality of households' perceived and expected inflation is generally rejected, as both fail either the condition of unbiasedness or of efficiency, or both. Studies that reject the rationality of inflation expectations include, *inter alia*, Forsells and Kenny (2002) for the Euro area as well as Thomas (1999), Mankiw, Reis and Wolfers (2004) and Souleles (2004) for the US, where the latter uses micro survey data instead of aggregate data. Jonung and Laidler (1988) as well as Lein and Maag (2011) reject the rationality of inflation perceptions for a panel of European countries.

Regarding the interrelation of perceived and expected inflation, a number of empirical studies suggest that households often form inflation expectations based on their perception of past inflation. Analyzing qualitative responses of the February 2008 issue of the Bank of England/GfK NOP Inflation Attitudes Survey, Benford and Driver (2008) report that almost 50% of respondents stated that inflation perceptions both over the past six months and over the past year and longer were 'very important' when forming expectations. More recently, Maag (2010) finds in a Gaussian mixture model, using micro data from the Swedish Consumer Tendency Survey, that about 51% of households form static inflation expectations on the basis of perceived inflation, while only 19% form forward-looking expectations based on actual inflation. Further evidence of inflation perceptions feeding into expectations is presented by Blanchflower and Kelly (2008) who report that groups with biased perceptions also form biased expectations.

However, there exists also empirical evidence of causality running from inflation expectations to perceptions: Traut-Mattausch *et al.* (2004) conduct experiments with restaurant menus denoted in Euro and in D-Mark and find that in all studies price trend perceptions are significantly biased towards price increases. This bias is not due to memory biases or inaccurate recall, but persists even when the original prices in the past are provided. The authors, therefore, attribute their finding to selective outcome correction resulting in the so-called 'expectancy confirmation hypothesis': Agents that expect prices to rise, will also perceive the price increases since calculation errors are more thoroughly corrected when they disconfirm the initial expectations than otherwise. Evidence in line with expectancy confirmation in the context of inflation is also provided for instance by Hofmann, Kirchler and Kamleitner (2007).

The conceptual framework in Ranyard *et al.* (2008) nicely shows the complexity of these interactions and the possibility for causality in both directions. Moreover, the aspect of social amplification via the media or social interaction is put forward. Therefore, we base our empirical analysis on the model in Ranyard *et al.* (2008) and aim at investigating empirically both the interactions and the direction of causality between actual, perceived and expected inflation as well as the role of the media.

The role of media news on the formation of inflation expectations is also put forward in several theoretical approaches in the context of information frictions. In models of rational inattention as in Sims (2003) and in models of sticky information as in Mankiw and Reis (2002), agents face information frictions either due to a limited capacity of processing information or due to an exogenous constraint on the arrival of new information. They are thus forced to form expectations with limited information. In his epidemiology model, Carroll (2001) then proposes the media to be the most important source of information about inflation developments, linking the intensity of news reporting to the share of agents using the most recent information set in a sticky-information setting. Similarly, Sims (2003) highlights the importance of the media in providing information coding services. In this line of argument, more news reporting should thus lead to better informed inflation expectations.

However, while the theories discussed above regard the media as an exogenous and unbiased factor, the agenda-setting theory developed in media studies by McCombs (2004) discusses how the intensity of media reporting and the interpretation of topics by the media may be explained. Thus, the media are not only regarded as a sender of information, but may add their own interpretation of topics or push new topics on the agenda, i.e. the public discussion. Moreover, as media outlets pursue economic interests, they may have an interest in emphasizing 'bad' over 'good' news as these sell better, thus adding an interpretation and a possible bias to the information content of the news. Indeed, Soroka (2006) shows that the media report negative news more extensively than positive news, resulting in asymmetric news coverage.

With regard to the empirical effect of the media on inflation expectations and perceptions, Lamla and Lein (2008, 2010) highlight the importance of media reports both as a transmission mechanism of information and as a possible cause for a bias. Especially with regard to inflation expectations, the authors find using German survey data that the 'tone' of an article may bias expectations, as they react more strongly to negative news.

III. Data description

Data for monthly inflation perceptions and expectations are obtained from the Swedish Consumer Tendency Survey. During the first two weeks of each month, a random sample of about 1,500 individuals is interviewed via telephone, where the target population is the Swedish public aged 16–84.[2]

The Swedish Consumer Tendency survey coincides with the Joint Harmonized EU Program of Business and Consumer Surveys conducted by the European Commission. However, in addition to questions Q5 and Q6 asking for a qualitative measure of inflation perceptions and expectations, the Swedish survey also asks respondents for a quantitative evaluation of perceived and expected inflation. The questions asking for quantitative inflation perceptions and expectations read as follows:

5a–b. 'Compared with 12 months ago, how much higher in percent do you think that prices are now? (Average)'

6a–b. 'Compared with today, how much in percent do you think that prices will go up (i.e. the rate of inflation 12 months from now)?'

The quantitative questions are located in the questionnaire directly after the qualitative questions asking about 'prices in general', so that the framing of the questions with regard

[2] For further information on the Swedish Consumer Tendency Survey and the full questionnaire, see http://www.konj.se. The survey is also discussed in detail in Palmqvist and Strömberg (2004).

to consumer price inflation seems well identified. We use mean responses to questions 5a–b. and 6a–b. as our measure of inflation perceptions (π^{p}) and inflation expectations (π^{e}) respectively.[3]

The data on media report about inflation is taken from a unique data set for Sweden assembled by the media research institute MediaTenor.[4] For the time span of January 1998 to December 2008, all articles related to inflation that were published in the Swedish newspaper 'Svenska Dagbladet' were coded according to a codebook in line with the standards of media content analysis.[5] Further details on the coding of our media data set are provided in Data S1. In addition to the variable *vol_articles* including all articles on inflation in a given month, we can distinguish between the number of articles on increasing (*vol_increase*) and decreasing (*vol_decrease*) inflation. This allows us to differentiate between a general news effect and, possibly asymmetric, news effects from 'bad' news on rising inflation and 'good' news on declining inflation rates.

In addition to the data on inflation expectations, perceptions and media reports, a number of macroeconomic control variables are incorporated into the analysis. These include actual HICP inflation rates (π),[6] an indicator of industrial production (*prod_ind*), the harmonized monthly unemployment rate (u), long-term interest rates defined as the 10-year yield on government bonds (i^{long}), short-term interest rates defined as 3-months treasury bills (i^{short}) and the growth rate of money supply M2 ($m2$). All macroeconomic variables are obtained from the OECD Main Economic Indicators Database.

Our media data set covers the time period 1998*m*1–2008*m*12. However, in 2008 rising food and energy prices caused a spike in actual, perceived and expected inflation, which was accompanied by a strong increase in media reports about inflation.[7] Due to the end-of-sample problem, in the following we present empirical estimates for the stable inflation period 1998*m*1–2007*m*12 and leave an analysis of later periods for future research.

IV. Rationality tests

Before proceeding to the econometric analysis, we test all variables for unit roots and cointegration. The results are available in Data S1. We find some evidence of non-stationarity of actual inflation and stronger evidence of a unit root in expected and perceived inflation from ADF, GLS-detrended DF, Phillips–Perron and KPSS tests, while the media variables appear

[3] The Swedish Consumer Tendency survey does not provide the median of answers across respondents. Rather, mean responses corrected for extreme values of inflation perceptions and expectations are available. We checked for robustness of all our results with respect to the corrected measures. Since we found no significant changes in the results, we use the encompassing mean of all survey answers in this study.

[4] See http://www.mediatenor.com.

[5] Due to resource restrictions, unfortunately it was not possible to extend the media data set to include the most recent periods. We leave this for future research.

[6] Note that we take HICP inflation to represent the true inflation rate, which should be captured by inflation perceptions and expectations. However, HICP inflation is likely to be biased upwards due to quality improvements and product innovations, which may not yet be accounted for in the underlying basket of consumer goods. Nevertheless, as HICP inflation is the measure of inflation most frequently discussed by policy officials and the media, we take it as a natural benchmark against which perceptions and expectations can be evaluated.

[7] The data over the whole sample period are shown graphically and discussed in Data S1.

to be stationary.[8] Thus, we proceed to test for cointegration between the inflation variables with a Johansen test. All tests suggest the existence of two cointegration relations when testing in a trivariate vector error correction model (VECM) between actual, perceived and expected inflation. This result is in line with the findings in Lein and Maag (2011), who report evidence of panel cointegration between actual and perceived inflation for a European sample.

Next, we test for rationality of inflation perceptions and expectations. In line with the literature, we evaluate three different aspects, namely accuracy, unbiasedness and efficiency: First, we analyze accuracy by reporting mean absolute errors (MAE) and root mean squared errors (RMSE), which are normalized by average inflation over the sample period. Next, we test for a bias with respect to inflation perceptions and expectations respectively. Under the null hypothesis of no bias, today's inflation perceptions as well as inflation expectations one year ago should be an unbiased predictor of today's actual inflation rate. Hence, we estimate the regressions

$$\pi_t = \alpha + \beta \pi_t^{\mathrm{p}} + \epsilon_t \tag{1}$$

$$\pi_t = \alpha + \beta \pi_{t-12}^{\mathrm{e}} + \epsilon_t \tag{2}$$

and test for the joint null hypothesis $H_0 : (\alpha, \beta) = (0, 1)$. In order to account for the cointegration between π and π^{p} or π^{e}, respectively, we estimate equations (1) and (2) in a VECM and test for H_0 regarding the cointegration relations contained in the cointegrating vector. H_0 is tested with a Wald test using the Johansen maximum likelihood estimator (see Lütkepohl and Krätzig, 2004, pp. 103–105 & 121–122).[9] Furthermore, we present results of Wilcoxon signed-rank tests for the null hypotheses of $H_0 : E(\pi_t - \pi_t^{\mathrm{p}}) = 0$ and $H_0 : E(\pi_t - \pi_{t-12}^{\mathrm{e}}) = 0$ respectively. With the advantage of being unaffected by non-normal distributions, this non-parametric test procedure tests for the equality of matched pairs of observations by assuming that both distributions are symmetric.

Finally, we evaluate whether perceptions and expectations efficiently incorporate available information. First, weak-form efficiency is rejected if perception or expectation errors are persistent or significantly correlated with lagged inflation rates. Second, strong-form efficiency is tested by regressing errors on a larger set of macroeconomic and monetary aggregates. All variables are lagged by 13 months in order to exclude the possibility of overlapping errors within the 12-month horizon of forecasts/backcasts and to allow for one publication lag. To avoid spurious results, we use first differences of all variables that were found to be non-stationary and report robust Newey–West standard errors.

Table 1 presents results of the three rationality tests for inflation perceptions in Sweden. Both the normalized MAE (0.386) and RMSE (0.485) are relatively high, considering that inflation perceptions are based on actual inflation and should thus be less prone to inaccuracy than inflation expectations. Lein and Maag (2011) report similar results using Swedish data for the time span of 1993–2007. With respect to a bias in perceptions, both the Wald test from a VECM estimation and the non-parametric Wilcoxon signed-rank test reject the null hypothesis of no bias in inflation perceptions at the 1% level. Finally, we find evidence of both weak-

[8] With respect to macro and money aggregates, we found a unit root in prod_ind, u, i^{long} and i^{short} and therefore use differences of these variables in all regressions. Results for the unit root tests can be obtained from the author upon request.

[9] The Wald tests on restrictions in the cointegration relation are carried out with the software JMulTi, version 4.24 (http://www.jmulti.de), which implements this test with the Johansen maximum likelihood estimator.

TABLE 1

Testing the rationality of inflation perceptions

Accuracy:	*MAE*		*RMSE*
	0.386		0.485
Bias: $\pi_t = \alpha + \beta\pi_t^p$	*Wald/z-stat.*	*Probability*	
$H_0\colon (\alpha, \beta) = (0, 1)$	44.261	0.000	VECM, 12 lags
$H_0\colon E(\pi_t - \pi_t^p) = 0$	−5.594	0.000	Wilcoxon signed-rank test
Efficiency: π^p*_error*	*(1)*	*(2)*	*(3)*
π^p_error$_{t-13}$	−0.256*		
	(0.156)		
$d(\pi)_{t-13}$		−0.436***	−0.362***
		(0.112)	(0.129)
$d(\pi^p)_{t-13}$			0.179
			(0.163)
$d(\pi^e)_{t-13}$			−0.069
			(0.146)
d(prod_ind)$_{t-13}$			0.021
			(0.020)
$d(u)_{t-13}$			0.041
			(0.089)
$d(i^{\text{long}})_{t-13}$			−0.599**
			(0.247)
$d(i^{\text{short}})_{t-13}$			1.261***
			(0.369)
$m2_{t-13}$			−0.719
			(4.225)
vol_articles$_{t-13}$			0.008
			(0.009)
ADF test resid.	−4.170	−4.334	−4.956
1% critical value	−2.599	−2.599	−2.599

Notes: Newey–West standard errors in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively. Sample period: 1998m1–2007m12.

and strong-form inefficiency: Perception errors are significantly negatively correlated with own past errors and with past inflation changes. This latter finding remains robust when we include a larger set of explanatory variables, where we find additional significant effects of past changes in short- and long-run interest rates.

Results of the rationality tests for inflation expectations in Sweden are given in Table 2. Regarding the accuracy of inflation forecasts, both the MAE (0.537) and RMSE (0.675) are larger than those found for perceptions. This is not surprising, since inflation expectations are formed with respect to an uncertain future. Testing for a bias in inflation expectations, both the Wald test in a VECM setting and the Wilcoxon signed-rank test reject the null of no bias at the 1% level. However, in contrast to our results for inflation perceptions, we find only weak evidence of inefficiency of expectations with a marginally significant effect of past changes in short-run interest rates.

TABLE 2

Testing the rationality of inflation expectations

Accuracy:		*MAE*	*RMSE*
		0.537	0.675
Bias: $\pi_t = \alpha + \beta \pi^{e}_{t-12}$	*Wald/z-stat.*	*Probability*	
			VECM, 12 lags
$H_0\colon (\alpha, \beta) = (0, 1)$	30.921	0.000	Wilcoxon signed-
$H_0\colon E(\pi_t - \pi^{e}_{t-12}) = 0$	−4.984	0.000	rank test
Efficiency: π^{e}*_error*	(1)	(2)	(3)
π^{e}_error$_{t-13}$	0.167		
	(0.162)		
$d(\pi)_{t-13}$		−0.099	0.018
		(0.216)	(0.277)
$d(\pi^{e})_{t-13}$			−0.107
			(0.187)
$d(\pi^{p})_{t-13}$			0.064
			(0.178)
d(prod_ind)$_{t-13}$			−0.012
			(0.042)
$d(u)_{t-13}$			0.017
			(0.134)
$d(i^{\text{long}})_{t-13}$			−0.655
			(0.531)
$d(i^{\text{short}})_{t-13}$			1.705*
			(0.945)
$m2_{t-13}$			2.666
			(5.817)
vol_articles$_{t-13}$			0.025
			(0.018)
ADF test resid.	−2.956	−2.963	−3.320
1% critical value	−2.602	−2.599	−2.599

Notes: Newey–West standard errors in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively. Sample period: 1998m1–2007m12. Note that the sample period for the efficiency tests is reduced to 2000m2–2007m12 in model (1) because realized expectations errors in our sample period begin in 1999m1.

V. Structural vector error correction models

After having rejected rationality, we turn to evaluating the interrelation between actual inflation, perceptions and expectations, while allowing for media effects. This is done in the framework of SVECMs, where we model the cointegration between actual, perceived and expected inflation. Additionally, we account for the possible endogeneity of the media with respect to inflation by including media variables as endogenous regressors in the model. Since the media variables are found to be stationary, we construct artificial non-stationary media series, $\Delta^{-1} media_t$, by cumulating the original data. These are then included in the SVECMs, where we impose zero restrictions on $\Delta^{-1} media_t$ in the cointegration matrix β.

Throughout the following analysis, we compare a model with the general media variable $\Delta^{-1}vol\_articles$ to a model including the cumulated numbers of articles on increasing and decreasing inflation, $\Delta^{-1}vol\_increase$ and $\Delta^{-1}vol\_decrease$. In line with the information criteria, all models are estimated with two lags.[10]

The VECM may thus be stated as follows (Kirchgässner and Wolters, 2008, p. 229):

$$\Delta y_t = \alpha\beta' y_{t-1} + \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} + C_0 D_t + u_t \tag{3}$$

where $y_t = (\pi_t \quad \pi_t^{\mathrm{p}} \quad \pi_t^{\mathrm{e}} \quad \Delta^{-1}media_t)'$ is the vector of endogenous variables, D_t includes the constant and seasonal dummies and u_t is the vector of reduced-form residuals. The cointegration relations between the variables are estimated in the matrix β. α contains the loading coefficients and Γ_1, Γ_2 and C_0 are coefficient matrices.

The reduced-form VECM given in equation (3) has the Beveridge–Nelson MA representation (omitting the vector D_t of deterministic variables for ease of exposition):

$$y_t = \Xi \sum_{i=1}^{t} u_i + \sum_{j=0}^{\infty} \Xi_j^* u_{t-j} + y_0^* \tag{4}$$

where the matrices Ξ_j^* are absolutely summable, so that they converge towards zero as $j \to \infty$, and y_0^* is a vector of initial values. As shown in Lütkepohl (2005, p. 369), long-run effects of the shocks on y_t are captured by the common trends in $\Xi \sum_{i=1}^{t} u_i$, where $\Xi = \beta_\perp \left[\alpha'_\perp \left(I_K - \sum_{i=1}^{p-1} \Gamma_i\right) \beta_\perp\right]^{-1} \alpha'_\perp$.[11]

The SVECM may then be identified by imposing restrictions on the contemporaneous relations in the model. The structural shocks, denoted by $\tilde{u}_t$, are typically identified with a B-Model (Lütkepohl, 2005, p. 369):

$$u_t = B\tilde{u}_t \quad \text{with} \quad \tilde{u}_t \sim (0, I_K) \tag{5}$$

where the matrix B captures the contemporaneous effects of the shocks and includes the restrictions and $\tilde{u}_t$ is a white noise error vector. The long-run effects of the structural shocks are then given by ΞB. The estimate of B can be used to identify impulse-responses and FEVDs in response to structural shocks.

We estimate all models with two cointegration relations in β in line with the Johansen cointegration tests, where the first column excludes actual inflation π_t and the second column excludes perceived inflation π_t^{p}.[12] The first coefficient of each cointegration relation is normalized to 1 and both columns include zero restrictions for $\Delta^{-1}media_t$. All reduced-form VECMs are estimated in a two-stage procedure, where the cointegration matrix β is estimated with the S2S estimator in the first stage and the remaining coefficients of the VECM are estimated with OLS in the second stage (Lütkepohl and Krätzig, 2004, pp. 103–105).

[10] Generally, all VECMs satisfy stability conditions and we find no evidence of autocorrelation or heteroscedasticity in the residuals. Test results are available from the author upon request.

[11] In our VECMs, we have the lag length $p = 2$ and the number of endogenous variables $K = 4$, i.e. $K = 5$ in the second model. Note that $\beta_\perp$ and $\alpha'_\perp$ denote the orthogonal complements of β and α'.

[12] All SVECM estimations in this and the following sections are carried out with the open-source software package JMulTi, version 4.24 (http://www.jmulti.de, see also Lütkepohl and Krätzig, 2004).

The B-matrix identifying the SVECM is estimated with maximum likelihood using the concentrated likelihood function. We impose restrictions on B by assuming it to be lower triangular. This identification is chosen based on theoretical and empirical observations: We argue that actual inflation is unlikely to be affected by shocks to perceived and expected inflation rates (or the media) in the same month due to the generally observed stickiness of prices incorporated into most modern macroeconomic models. Also, since several authors find empirical evidence that inflation expectations are largely formed on the basis of perceived inflation (see, e.g. Benford and Driver, 2008; Maag, 2010), we allow for a contemporaneous effect of shocks to perceptions on expectations, but not *vice versa*. Finally, in order to account for a possible publication lag, we only allow lagged effects of media shocks on actual, perceived and expected inflation. In the model with $\Delta^{-1}vol\_increase$ and $\Delta^{-1}vol\_decrease$, we additionally exclude any contemporaneous effect between the two media variables, so that the second SVECM is overidentified.

In the following subsections, we evaluate the models with respect to the cointegration relations and the short-run media effects, analyze impulse response functions and FEVDs and test for long-run and short-run Granger causality.

Cointegration relations and short-run media effects

Tables 3 and 4 present estimates of the cointegration relations in the matrix β as well as the reduced-form loading coefficients in α, including the aggregate media variable $\Delta^{-1}vol\_articles$ in the first model, and $\Delta^{-1}vol\_increase$ and $\Delta^{-1}vol\_decrease$ in the second model. As expected, we find cointegration between inflation perceptions and expectations as well as between actual and expected inflation in all models, where the coefficients in the vector β_1 suggest a one-to-one cointegration relation between perceived and expected inflation.

Analyzing the loading coefficients contained in α, we find that inflation perceptions adjust to the first long-run equilibrium between π^p and π^e. Similarly, actual inflation adjusts to the second long-run equilibrium between π and π^e. This result is robust in both models.

TABLE 3

Cointegration relation and loadings in the structural vector error correction model with news on inflation

	α_1	α_2	β_1	β_2
π	0.102	−0.177***	–	1.000
	(0.117)	(0.060)		–
π^p	−0.216**	0.073	1.000	–
	(0.089)	(0.046)	–	
π^e	−0.002	0.003	−0.957***	−0.637**
	(0.098)	(0.050)	(0.144)	(0.279)
$\Delta^{-1}vol\_articles$	−0.950	−0.120	–	–
	(1.456)	(0.748)		

Notes: Standard errors in parentheses. *,** and *** significance at the 10%, 5% and 1% level respectively. The model is estimated with zero restrictions on the media variables in the cointegrating vector β. Sample period: 1998*m*1–2007*m*12.

TABLE 4

Cointegration relation and loadings in the structural vector error correction model with news on increasing or decreasing inflation

	α_1	α_2	β_1	β_2
π	0.131	−0.205***	–	1.000
	(0.115)	(0.060)		–
π^{p}	−0.191**	0.068	1.000	–
	(0.090)	(0.047)	–	
π^{e}	0.041	−0.013	−1.135***	−1.121***
	(0.096)	(0.051)	(0.152)	(0.252)
Δ^{-1}*vol_increase*	−1.031**	0.316	–	–
	(0.500)	(0.262)		
Δ^{-1}*vol_decrease*	−0.009	0.140**	–	–
	(0.139)	(0.073)		

Notes: Standard errors in parentheses. *,** and *** significance at the 10%, 5% and 1% level respectively. The model is estimated with zero restrictions on the media variables in the cointegrating vector β. Sample period: 1998m1–2007m12.

TABLE 5

Short-run media effects

Model	*Media variables*	$\Delta\pi$	$\Delta\pi^{\mathrm{p}}$	$\Delta\pi^{\mathrm{e}}$
(1)	$vol\_articles_{t-1}$	−0.005	−0.003	−0.001
		(0.007)	(0.006)	(0.006)
	$vol\_articles_{t-2}$	0.004	0.005	0.003
		(0.007)	(0.006)	(0.006)
(2)	$vol\_increase_{t-1}$	−0.007	−0.004	0.024
		(0.022)	(0.017)	(0.018)
	$vol\_increase_{t-2}$	0.016	0.027	0.010
		(0.022)	(0.017)	(0.018)
	$vol\_decrease_{t-1}$	0.151**	0.007	0.062
		(0.071)	(0.055)	(0.059)
	$vol\_decrease_{t-2}$	0.184***	0.012	0.104*
		(0.070)	(0.055)	(0.059)

Notes: Standard errors are in parentheses, where *, ** and *** denote significance at the 10%, 5% and 1% level respectively. Sample period: 1998m1–2007m12.

Regarding the effect of the media, we find significant loadings only in the second VECM with Δ^{-1}*vol_increase* and Δ^{-1}*vol_decrease*.

Next, we evaluate the short-run effects of the media in more detail. These are presented in Table 5. Since short-run effects in a VECM are from the differenced endogenous variables, in the case of our cumulated media variables these measure in fact the short-run impact of the level media variables *vol_articles* as well as *vol_increase* and *vol_decrease*.

While short-run effects from the total number of articles on inflation are insignificant in the first model, we find some media effects on actual and expected inflation when distinguishing between articles on increasing or decreasing inflation in the second VECM. These hint at a certain asymmetry in the effect of news regarding rising or falling inflation: While we find no effects from articles on increasing inflation, the results suggest a positive effect of past reports on decreasing inflation on actual and expected inflation. This suggests that past reports on reduced inflation rates may be perceived as 'negative' rather than 'positive' news, thus introducing a media bias. A media bias from news with a negative tone is also found by Lamla and Lein (2008) in their analysis of German households' inflation expectations.

Impulse responses

Next, we analyze the impulse responses to structural shocks identified with the matrix B. The impulse responses of the two SVECMs are shown in Figures 1–2, where the columns present responses to shocks in actual inflation π_t, perceptions π_t^{p}, expectations π_t^{e} and the cumulated media variable(s) $\Delta^{-1} media_t$, respectively.[13] Generally, impulse responses suggest considerable interaction between actual, perceived and expected inflation. As expected, a positive shock to the actual inflation leads to an increase of both perceived and expected inflation, where the effect on expectations is less pronounced. Conversely, actual inflation is positively affected by a shock to perceived, but not to expected, inflation. Interestingly, both SVECMs suggest a strong and persistent positive effect of a shock to perceptions on expected inflation, while the reverse effect only becomes significant after 12–14 months.

Overall, this implies that consumers' inflation expectations are affected more by changes in perceived, rather than actual inflation. This result is in line with the findings in the survey by Benford and Driver (2008) and the results from the Gaussian Mixture model in Maag (2010), using the Swedish Consumer Tendency survey data also analyzed here. However, we also find some weak support for the reverse hypothesis of expectancy confirmation in Traut-Mattausch *et al.* (2004), since a shock to inflation expectations has a minor positive effect on perceived inflation in both SVECMs.

Regarding the effect of media shocks on actual, perceived and expected inflation, the impulse response functions confirm our result of the previous section: We see no effect of shocks to $\Delta^{-1} vol\_articles$, but find some interaction between the media and the inflation variables in the second SVECM with the cumulated numbers of articles on increasing or decreasing inflation. Effects of actual, perceived and expected inflation on $\Delta^{-1} vol\_increase$ and $\Delta^{-1} vol\_decrease$ are generally as expected. By contrast, the impulse responses of π_t, π_t^{p} and π_t^{e} to media shocks again suggest some asymmetry in the effect of 'positive' vs. 'negative' inflation news: A sudden increase in news on rising inflation only affects inflation expectations, where we see a marginally significant positive effect. Conversely, a sudden increase in the news coverage on decreasing inflation has a significantly positive effect on actual inflation, perceptions and expectations, while a negative relation would be expected. The media effects on households' perceptions and expectations suggest that media news might be perceived with a negative bias.

Forecast error variance decomposition

FEVDs are presented in Tables 6–7 for the two SVECMs. In line with the results from the impulse responses in the previous section, FEVDs from both SVECMs suggest a strong

[13] Note that impulse responses are calculated as the effect of structural shocks on the level variables in the SVECMs. Since these are non-stationary, a temporary shock translates into a permanent effect on the level variable.

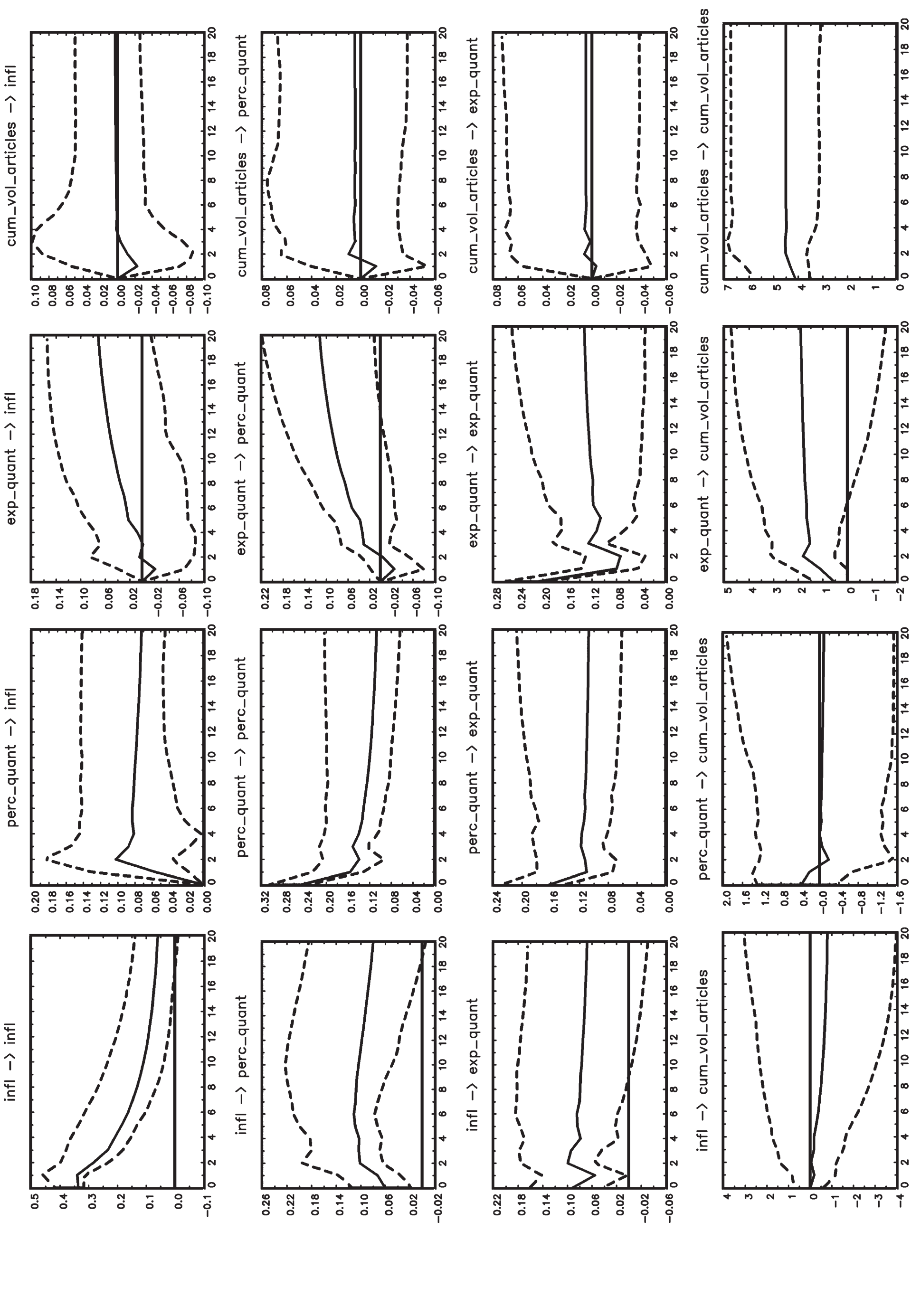

Figure 1. Structural vector error correction model impulse responses in the model with Δ^{-1} *vol_articles*

Notes: Impulse reponses with 95% confidence intervals from the Hall percentile interval with 100 bootstrap replications.

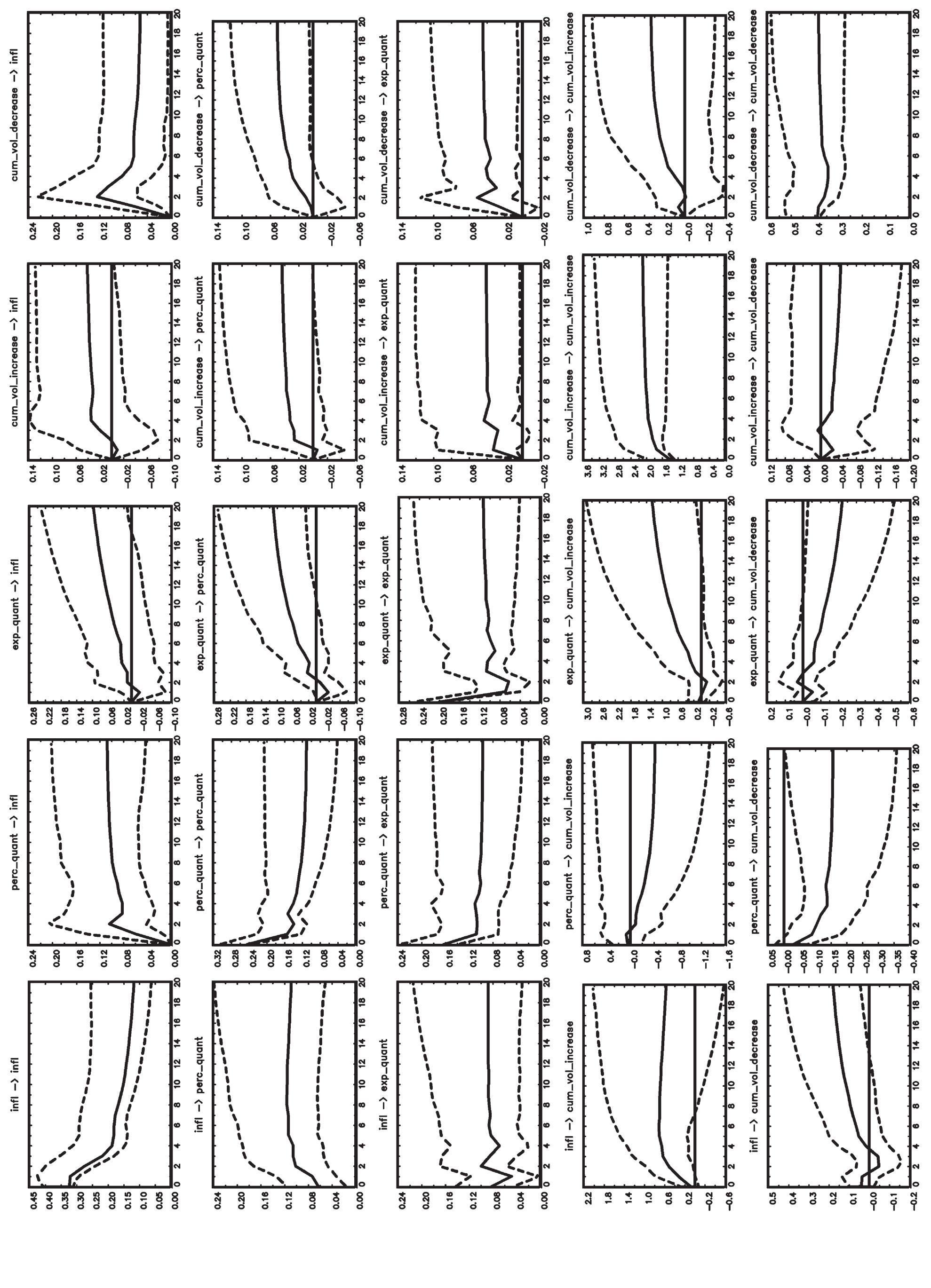

Figure 2. Structural vector error correction model impulse responses in the model with Δ^{-1}*vol_increase* and Δ^{-1}*vol_decrease*

Notes: Impulse responses with 95% confidence intervals from the Hall percentile interval with 100 bootstrap replications.

TABLE 6

Forecast error variance decomposition in the structural vector error correction model with $\Delta^{-1}vol\_articles$

	Forecast horizon	*1 month*	*25 months*	*50 months*
π	% due to π	100	74	56
	% due to π^{p}	0	18	23
	% due to π^{e}	0	8	21
	% due to $\Delta^{-1}vol\_articles$	0	0	0
π^{p}	% due to π	6	26	20
	% due to π^{p}	94	51	43
	% due to π^{p}	0	23	37
	% due to $\Delta^{-1}vol\_articles$	0	0	0
π^{e}	% due to π	10	16	14
	% due to π^{p}	34	37	35
	% due to π^{e}	56	47	51
	% due to $\Delta^{-1}vol\_articles$	0	0	0
$\Delta^{-1}vol\_articles$	% due to π	0	2	3
	% due to π^{p}	1	0	0
	% due to π^{e}	2	13	14
	% due to $\Delta^{-1}vol\_articles$	97	85	83

Notes: Sample period: 1998*m*1–2007*m*12.

degree of interaction between actual, perceived and expected inflation. This is especially true for perceived and expected inflation, which both explain about 30–37% of the variance of the other variable's forecast error, while actual inflation explains a smaller fraction of the forecast error variance (14–27%).

Again, we find no media effects from the cumulated number of all articles regarding inflation and $\Delta^{-1}vol\_articles$ seems to be largely exogenous, with only a small fraction of its forecast error variance explained by inflation expectations. In the second SVECM, FEVDs suggest a stronger media impact, where in the longer run $\Delta^{-1}vol\_increase$ and $\Delta^{-1}vol\_decrease$ explain about 5% each of the forecast error variances of actual, perceived and expected inflation. In turn, both media variables' FEVD are to some degree explained by the inflation variables, especially by inflation expectations.

Granger causality

Finally, we test for Granger causality between the inflation and the media variables.[14] As in Granger (1969), a variable is said to be Granger causal for another variable if it contains useful information for predicting that latter variable. Generally, in a cointegrated system there always exists at least one Granger-causal relationship (Kirchgässner and Wolters, 2008, p. 232). The concept of Granger causality in cointegrated systems is discussed in Kirchgässner and Wolters (2008), who define the distinction between short- and long-run Granger causality in a VECM framework. We thus test for both short- and long-run causality between the variables in the VECMs with π_t, π_t^{p}, π_t^{e} and $\Delta^{-1}media_t$. Assuming

[14] All tests for Granger causality are performed in STATA. As before, all VECMs are estimated with two cointegration relations and two lags.

TABLE 7

Forecast error variance decomposition in the structural vector error correction model with Δ^{-1}vol_increase *and* Δ^{-1}vol_decrease

	Forecast horizon	1 month	25 months	50 months
π	% due to π	100	61	45
	% due to π^{p}	0	19	23
	% due to π^{e}	0	10	21
	% due to Δ^{-1}*vol_increase*	0	2	4
	% due to Δ^{-1}*vol_decrease*	0	8	7
π^{p}	% due to π	6	30	27
	% due to π^{p}	94	43	35
	% due to π^{e}	0	18	28
	% due to Δ^{-1}*vol_increase*	0	4	5
	% due to Δ^{-1}*vol_decrease*	0	5	5
π^{e}	% due to π	9	22	22
	% due to π^{p}	36	32	30
	% due to π^{e}	55	37	38
	% due to Δ^{-1}*vol_increase*	0	4	5
	% due to Δ^{-1}*vol_decrease*	0	5	5
Δ^{-1}*vol_increase*	% due to π	0	7	5
	% due to π^{p}	0	2	2
	% due to π^{e}	0	16	24
	% due to Δ^{-1}*vol_increase*	100	74	68
	% due to Δ^{-1}*vol_decrease*	0	1	1
Δ^{-1}*vol_decrease*	% due to π	1	9	13
	% due to π^{p}	1	10	9
	% due to π^{e}	0	12	20
	% due to Δ^{-1}*vol_increase*	0	1	0
	% due to Δ^{-1}*vol_decrease*	98	68	58

Notes: Sample period: 1998*m*1–2007*m*12.

two (groups of) endogenous variables y_t^1 and y_t^2 in the vector y_t, the reduced-form VECM presented in equation (3) takes the following form:

$$\begin{pmatrix} \Delta y_t^1 \\ \Delta y_t^2 \end{pmatrix} = \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{pmatrix} \left(\mathrm{ecm}_{t-1} \right) + \sum_{i=1}^{p} \begin{pmatrix} \gamma_i^{11} & \gamma_i^{12} \\ \gamma_i^{21} & \gamma_i^{22} \end{pmatrix} \begin{pmatrix} \Delta y_{t-i}^1 \\ \Delta y_{t-i}^2 \end{pmatrix} + \begin{pmatrix} u_t^1 \\ u_t^2 \end{pmatrix} \tag{6}$$

where the matrix α contains the loading coefficients, $\mathrm{ecm}_{t-1} = \beta' y_{t-1}$ denotes the two long-run cointegration relations, u_t is the vector of reduced-form residuals and the lag-length is $p = 2$. Let $\gamma_{12} := (\gamma_1^{12},\ldots, \gamma_p^{12})'$ be the vector of short-run coefficients on lags of Δy_t^2 and $\alpha_1 := (\alpha_{11}\alpha_{12})'$ be the vector of loadings in the equation for Δy_t^1. Then, y_t^2 will be Granger causal for y_t^1 if a Wald test rejects the null hypothesis $H_0 : (\alpha_1\gamma_{12})' = 0$ vs. $H_1 : (\alpha_1\gamma_{12})' \neq 0$ and *vice versa*. To distinguish between short- and long-run Granger causality, we follow Kirchgässner and Wolters (2008, p. 232) (keeping our notation from above): 'There exists only 'short-run' causality [from y_2 to y_1] if $\alpha_1 = 0$ but $\gamma_{12} \neq 0$, and only 'long-run' causality if $\alpha_1 \neq 0$ but $\gamma_{12} = 0$.' Hence, under short-run Granger causality, y_2 only significantly affects the short-run dynamics of y_1, while under long-run causality we

TABLE 8

Granger causality tests

	(1) Δ^{-1}*vol_articles*	(2) Δ^{-1}*vol_increase* & Δ^{-1}*vol_decrease*
Overall causality $H_0: (\alpha_1\gamma_{12})' = 0$		
π^p, π^e, Δ^{-1}media $\rightarrow \pi$	12.43	20.58**
	(0.133)	(0.024)
π^e, π, Δ^{-1}media $\rightarrow \pi^p$	2.14	12.42
	(0.145)	(0.258)
π^p, π, Δ^{-1}media $\rightarrow \pi^E$	8.17	16.37*
	(0.417)	(0.089)
π^p, π^E, $\pi \rightarrow \Delta^{-1}$media	8.11	29.35**
	(0.423)	(0.022)
Long-run coefficients $H_0: \alpha_1 = 0$		
π^p, π^e, Δ^{-1}media $\rightarrow \pi$	7.89**	10.37***
	(0.019)	(0.006)
π^e, π, Δ^{-1}media $\rightarrow \pi^p$	6.53**	4.74*
	(0.038)	(0.094)
π^p, π, Δ^{-1}media $\rightarrow \pi^e$	0.12	0.34
	(0.941)	(0.845)
π^p, π^e, $\pi \rightarrow \Delta^{-1}$media	1.89	7.48
	(0.389)	(0.112)
Short-run coefficients $H_0: \gamma_{12} = 0$		
π^p, π^e, Δ^{-1}media $\rightarrow \pi$	3.43	12.66
	(0.753)	(0.124)
π^e, π, Δ^{-1}media $\rightarrow \pi^p$	5.33	5.57
	(0.502)	(0.695)
π^p, π, Δ^{-1}media $\rightarrow \pi^e$	7.87	13.11*
	(0.248)	(0.108)
π^p, π^e, $\pi \rightarrow \Delta^{-1}$media	4.82	23.49**
	(0.567)	(0.024)

Notes: χ^2 statistics with 8/10 degrees of freedom in the tests for overall causality, 2 degrees of freedom in the tests for long-run coefficients and 6/8 degrees of freedom in the tests for short-run coefficients in models (1) and (2), respectively, with p-values in parentheses. *, ** and *** denote rejection of H_0 at the 10%, 5% and 1% level. Sample period: 1998m1–2007m12.

only find significant loading coefficients to the long-run cointegration relations. In order to test this, we apply Wald tests on significance of the vectors α_1 and γ_{12} sequentially and interpret the results jointly.

Test results are presented in Table 8, where $\Delta^{-1}\text{media}_t$ is either Δ^{-1}*vol_articles*$_t$ or a block defined as (Δ^{-1}*vol_increase*$_t$, Δ^{-1}*vol_decrease*$_t$)'. Regarding the first VECM with Δ^{-1}*vol_article*, the overall test of no Granger causality, i.e. $H_0: (\alpha_1\gamma_{12})' = 0$, cannot be rejected for any of the blocks of variables. This is due to the fact that none of the short-run coefficients are jointly significant. However, for each cointegration relation we need at least one significant loading coefficient in α. Indeed, loading coefficients are jointly significant in the VECM equations for π and π^p. Together with the result of no significant short-run effects, this implies long-run Granger causality from (π^p, π^e, Δ^{-1}*vol_articles*) to π and from (π^e, π, Δ^{-1}*vol_articles*) to π^p.

In the second VECM including $(\Delta^{-1}vol\_increase_t, \Delta^{-1}vol\_decrease_t)'$, we can reject $H_0: (\alpha_1\gamma_{12})' = 0$ vs. $H_1: (\alpha_1\gamma_{12})' \neq 0$ in the equations for π, π^e and $\Delta^{-1}media$, where the last result tests jointly for the block of the two media variables. This implies that π, π^e and $(\Delta^{-1}vol\_increase_t, \Delta^{-1}vol\_decrease_t)'$ are Granger-caused by the other variables in the system respectively. Testing separately for long-run and short-run Granger causality, the result of long-run causality to π and π^p from the first VECM remains robust. Additionally, we find some evidence of short-run Granger causality, since we reject $H_0: \gamma_{12} = 0$ in the equations for $(\Delta^{-1}vol\_increase_t, \Delta^{-1}vol\_decrease_t)'$ and (weakly) in the equation for π^e, but cannot reject $H_0: \alpha_1 = 0$. The result of short-run Granger causality from the inflation variables to the media variables implies that the media should not be regarded as exogenous.

VI. Conclusion

In line with the conceptual framework presented in Ranyard *et al.* (2008), we find substantial interaction between actual inflation, households' perceptions and expectations. The evidence of strong effects of structural shocks to inflation perceptions on expected inflation supports the hypothesis that inflation expectations may be formed on the basis of perceived, rather than actual, inflation as suggested by Benford and Driver (2008) and Maag (2010). Nevertheless, evidence on a smaller feedback effect from expected to perceived inflation also gives some support to the hypothesis of expectancy confirmation, as put forward in Traut-Mattausch *et al.* (2004).

Furthermore, the finding that actual inflation plays a lesser role for the forecast error variance of both perceptions and expectations than the variables between themselves implies that to some extent, consumers' views on current and future inflation may be more related to their own beliefs than to the actual inflation environment. This has important policy implications for central banks aiming at stabilizing inflation: If central banks want to anchor inflation expectations and ensure that expectations are as informed as possible, they should pay attention also to consumers' inflation perceptions. In the light of discussions on macroeconomic illiteracy (Blanchflower and Kelly, 2008), our analysis thus highlights the importance of a good communication strategy from central banks to private households with the aim of informing consumers about actual inflation.

The role of the media in providing information on inflation seems, however, limited with regard to effects on inflation perceptions and expectations. While we find no significant effect of the general media variable counting all articles on inflation, small media effects emerge when we focus on news on increasing and decreasing inflation. However, these media effects might also introduce a bias, as it seems that households perceive news on decreasing inflation as negative, rather than, positive news and, hence, increase their inflation perceptions and expectations. This result is related to the findings in Lamla and Lein (2008) for media effects on inflation expectations of German households who report that a negative tone of media reports reduces consumers' forecast accuracy. Additionally, the media may react endogenously to changes in actual inflation as well as inflation attitudes. Indeed, our Granger-causality tests suggest short-run causality running from the inflation variables to the media. In that sense, while the media might provide an additional link between actual, perceived and expected inflation, our analysis suggests that central banks should explore other, more direct communication and information channels to avoid the risk of a bias from the media.

Final Manuscript Received: March 2014

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Data S1. Swedish Consumer Tendency Survey, MediaTenor dataset for Sweden and OECD Main Economic Indicators Database.