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International Transmissions of Inflation Expectations in a Markov Switching Structural VAR Model

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International Transmissions of Inflation Expectations in a Markov Switching Structural VAR Model

February 27, 2015

Abstract

This paper extends the discussion of international comovements of actual inflation rates to inflation expectations. Financial market expectations about inflation rates in the United States (US) and Euro Area (EA) are modeled in a structural vector autoregression (SVAR). We demonstrate how the heteroscedasticity of the expectations data enables a flexible and data-driven statistical identification of the model. A multi-step procedure is proposed to explore the economic nature and geographical source of structural shocks. We emphasize the SVAR's ability to derive shocks that disentangle US specific, EA specific and global components. Our main empirical finding indicates that so-called 'global inflation' translates to *short horizon* inflation expectations. In contrast, *long* expectations horizons are mostly driven by domestic shocks, thus, appear rather local. Results support the view of credible monetary policy strategies that anchor inflation expectations.

Keywords: Spillover, monetary policy, financial crisis, break-even inflation, identification through heteroskedasticity.

JEL classification: E31, F42, E52.

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1 Introduction

International comovements of the actual rate of inflation are a long-known and extensively studied fact, see Darby and Lothian (1983) for an early contribution. International trade, technology shocks, monetary policy and migration are considered most important in explaining cross country correlations, see Melitz and Ottaviano (2008), Henriksen et al. (2013) and Bentolila et al. (2008), respectively. For developing macroeconomic models and designing monetary policy rules, the *origin* of transmissions have recently aroused great interest. This has stimulated a new angle of empirical literature put forward by Ciccarelli and Mojon (2010), that aims at decomposing inflation rates into domestic and global components.

In this paper we extend the discussion about domestic and global components in inflation rates to inflation expectations. As demonstrated by Clarida et al. (2000) and Leduc et al. (2007), inflation expectations have attracted great attention by policy makers and academia. Especially in recent times of stable inflation and exceptionally expansive policy regimes, the key role played by inflation expectations in the conduct of monetary policy has become more explicit.¹ The New Keynesian Phillips Curve suggests a close connection between actual inflation rates and inflation expectations. The natural question arises if the strong international component in inflation, found by e.g. Neely and Rapach (2011) and Mumtaz and Surico (2012), is also present in inflation expectations.² Henriksen et al. (2013) suppose a positive answer to that question. Their international business cycle model draws a direct link between cross country correlations in expectations and cross country correlations in output and inflation. Based on empirical assessments, the authors confirm that similarities in expectations across industrialized countries increase with the expectations horizon. In contrast, studies on the anchoring of inflation expectations by Gürkaynak et al. (2010b) or Strohsal and Winkelmann (2015) consider inflation expectations to be independent across countries. They argue that inflation expectations at long horizons are country specific and mostly driven by the credibility of a central bank's inflation target. Regressions by Castelnuovo (2010) and Bayoumi and Swiston (2010) appear to support that expectations at different expectation horizons have only a weak international component.

The most frequently employed approach to derive domestic and global components from inflation rates is based on a reduced form dynamic latent factor model. Inspired by the concluding remarks of Ciccarelli and Mojon (2010), we follow the idea

¹The Federal Reserve (FED) defines medium term inflation expectations smaller 2.5% and anchored longer-term inflation expectations as two criteria to keep the target range for the federal funds rate at the zero lower bound; see FED (2012).

²The term 'global inflation' was introduced by Ciccarelli and Mojon (2010) and expresses that the variance of inflation rates of OECD countries are explained by up to 70% by a global factor.

that the question about international components is at its heart a structural one. We propose a parsimonious structural vector autoregressive (SVAR) analysis of inflation expectations data. Our modeling goal is to derive shocks that separate the structural drivers behind different expectations horizons and at the same time isolate their country specific origin. Similar to actual inflation in standard macroeconomic models, we think about demand, supply and monetary policy shocks as one group of potential structural drivers of inflation expectations. Following Leduc et al. (2007) and Beechey et al. (2011), such shocks affect especially short horizon expectations. At long horizons their impact should have already decayed. Besides shocks that are important for actual inflation, Bomfim and Rudebusch (2000) propose that credibility shocks with respect to a central bank's inflation target form a second group of structural drivers. Since it is difficult to find convincing just-identifying restrictions to set up a conventional structural VAR with inflation expectations, we exploit the heteroscedasticity in the data. The Markov switching SVAR model proposed by Lanne et al. (2010) appears suitable in the present context since it is data-driven and requires only mild assumptions. Furthermore, the statistical identification can be evaluated by formal tests, see Lütkepohl and Netšunajev (2014). However, the procedure of attaching labels to shocks is generally much more involved compared to classical identifying techniques since unique orthogonal shocks may lack economic interpretation. We propose a multi-step procedure which includes regressions of structural shocks on the surprise component of macroeconomic data releases and joint parameter tests to learn about their economic nature and geographical origin. We study a five variable SVAR comprising of short and long horizon financial market inflation expectations from the world's two largest economies, the United States (US) and Euro Area (EA), and a (global) commodity price index. Weekly data is considered in a period from 2004 to 2012. We show that the heteroscedasticity inherent in the data adds identifying information that avoids putting exclusion-, sign- or long run restrictions, compare Lütkepohl (2011). Volatility regimes governed by the Markov states provide time-varying transmissions that can be attributed to pre-crisis (around 2004 to 2007) and crisis periods. Shocks of the SVAR model do not only differ in their relative variances but also respond to different surprise variables of US and EA macroeconomic data releases. In combination with impulse responses and forecast error variance decompositions we attain a solid ground to confirm the economic meaning of shocks. We find that the variance of short horizon inflation expectations is mainly explained by structural shocks that pool information about demand, supply and monetary policy shocks. The responsiveness of structural shocks to subsets of US and EA specific data releases further allow conclusions about their US, EA or global origin. In contrast, structural shocks most relevant for long horizon expectations are insensitive to the surprise components of macroeconomic

data releases. We assign the labels FED and ECB credibility shocks, respectively. Overall, we emphasize the identification of US specific, EA specific and global shocks. In line with evidence on actual inflation by e.g. Neely and Rapach (2011), global shocks account for up to 71% (US) and 48% (EA) of the variance of short horizon inflation expectations. Spillovers of US (EA) specific shocks explain at most 19% of EA (6% of US) short run inflation expectations. During the global financial crisis, the great recession and the European sovereign debt crisis the role played by global shocks declines and country specific shocks, especially ECB credibility shocks, gain in importance. Compared to short horizon expectations, long horizon inflation expectations are far less determined by foreign shocks. For example, US and global shocks explain around 3% of the variance of EA long horizon inflation expectations only. Thus, in contrast to short horizon expectations, inflation expectations at long horizons appear rather local.

Our findings document that the debate about increasing international components in inflation rates and a related weakening of inflation control by national central banks translates to short horizon inflation expectations. In contrast, long horizon inflation expectations are less exposed to foreign shocks, hence, appear more directly controllable by national central banks. In the context of anchoring criteria, our results support the view of credible monetary policy strategies that anchor inflation expectations.

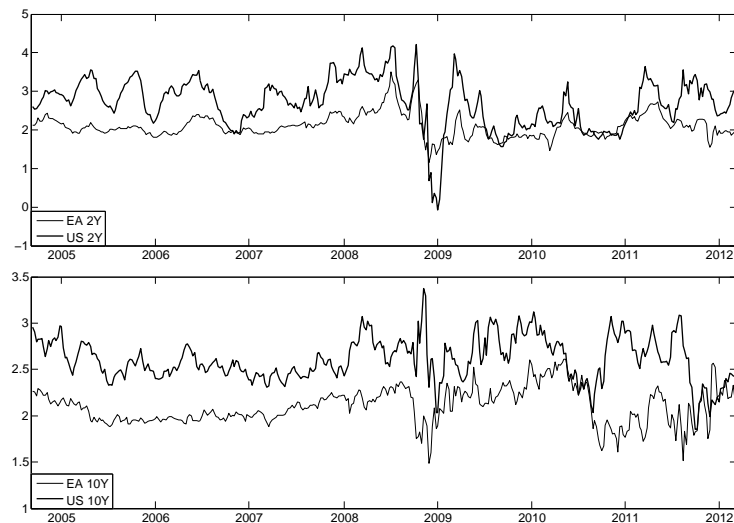
The rest of the paper is organized in 4 upcoming sections. Section 2 presents the data. Section 3 introduces the Markov switching SVAR model. The main part is Section 4. We first describe the estimation and identification and then present the results in terms of impulse responses and a forecast error variance decomposition. Section 6 concludes.

2 Data

Opposed to actual inflation rates, inflation expectations are not directly observable. A number of measures exist among which inflation surveys and financial market instruments are the most prominent sources. In this paper, we refer to financial market measures as they provide timely information about inflation expectations over a variety of expectations horizons.

The spread between yields of nominal and real (inflation indexed) government bonds, known as break-even inflation, is the basis of our expectations data. Because of differences in risk premia between nominal and real bonds, break-even inflation rates are not a pure measure of inflation expectations. Adjustment procedures of Gürkaynak et al. (2010a), Christensen et al. (2010) or Hördahl and Tristani (2012)

Figure 1: US and EA short and long horizon inflation expectations.



Notes: Two-year spot rate (upper figure) and one-year forward nine years ahead (lower figure). Weekly averages (Monday to Friday) of liquidity adjusted break-even inflation (BEI) rates (391 observations). For illustration purposes, adjusted BEI rates are centered around the sample mean of inflation in the EA (2.1%) and US (2.6%).

are advocated to obtain valid expectations. In this paper, we study weekly US and EA data in the time period from September 2004 to March 2012. A two-year spot rate and a one-year forward nine years ahead model the short and long expectations horizons. While the spot rate is meant to capture drivers important for building expectations over short horizons, the forward is meant to emphasize important drives for building long horizon expectations. We follow Gürkaynak et al. (2010a) and adjust each break-even inflation rate by regressions on country and horizon specific risk measures, compare also Söderlind (2011). The residuals of such regressions constitute the inflation expectations measure used in our structural VAR analysis.³ To account for a global driver of inflation expectations, we follow results provided by Leduc et al. (2007) and Ehrmann et al. (2014) and incorporate global commodity prices given by the S&P GSCI index as a fifth variable.⁴

Figure 1 illustrates the sample paths of the inflation expectations measures. Week-by-week expectations appear quite persistent. Conventional unit root tests suggest that the expectations measures as well as the commodity price index are stationary.

³US BEI rates are taken from the database of Gürkaynak et al. (2010a). EA BEI rates are obtained from the ECB and contain German, French and Italian bonds. More details about the liquidity adjustments are provided in Appendix A.

⁴The S&P GSCI index tracks global futures prices for agricultural goods, energy and industrial metals. The index is provided by the Macrobond database. We also tried other global price measures like oil prices. Results reported in Section 4 remain very similar.

The figure indicates that US expectations are more volatile than EA expectations. For the whole sample period, the standard deviation of the two-year US expectations is with 0.64 percentage points twice as large as corresponding EA expectations. Standard deviations of actual inflation rates in the US (1.6 percentage points) and EA (0.9 percentage points) share a very similar relative order of magnitude. Overall, a heteroscedastic pattern is clearly visible in the sample paths. In the following, we aim at exploiting the heteroscedasticity to identify the structural drivers behind inflation expectations.

3 The Markov switching SVAR model

The identification through heteroscedasticity is a powerful option to support identification of shocks in SVAR models, see Rigobon (2003) or Lanne and Lütkepohl (2008), among others. In comparison to classical identifying techniques like short run, long run or sign restrictions, the identification through heteroscedasticity is a more data oriented approach. This is also in sharp contrast to the identification strategies for latent dynamic factor models, previously applied to inflation series in e.g. Mumtaz and Surico (2012), where factor loadings are restricted to zero such that country specific factors are easily characterized by having no impact on foreign inflation. With the SVAR model and the identification through heteroscedasticity, we attempt to explore transmission channels and the economic nature of driving forces more deeply. We let the data speak about the statistical identification and check in a second step whether some economic meaning can be attached to the individual structural shocks. In general, our statistical procedure allows some country specific shock to transmit to foreign inflation expectations.

Given our data vector of two- and ten-year US and EA inflation expectations and commodity prices, $Y_t = (\pi_t^e(\text{EA 2Y}) \pi_t^e(\text{US 2Y}) \pi_t^e(\text{EA 10Y}) \pi_t^e(\text{US 10Y}) \text{Cmnty}_t)'$, we aim at identifying shocks ε_t through a structural VAR model with p lags:

$$Y_t = \nu + A_1 Y_{t-1} + \dots + A_p Y_{t-p} + B \varepsilon_t, \quad (1)$$

where ν is a constant intercept and the A_j s ($j = 1, \dots, p$) are 5×5 coefficient matrices. We follow Lanne et al. (2010) and model the heteroscedasticity of ε_t via a discrete Markov process s_t with states $1, 2, \dots, M$, transition probabilities $p_{ij} = \Pr(s_t = j | s_{t-1} = i)$, $i, j = 1, \dots, M$ and conditional distribution $\varepsilon_t | s_t \sim N(0, \Lambda_{s_t})$. The matrix $\Lambda_{s_t} = \text{diag}(\lambda_1, \dots, \lambda_5)$ is normalized such that the $\varepsilon_{t,k}$, $k = 1, 2, \dots, 5$, have unit conditional variance in the first state. Standard matrix algebra determines the

matrix B of impact effects:

$$\Sigma_1 = BB', \quad \Sigma_{s_t} = B\Lambda_{s_t}B', \quad s_t = 2, 3, \dots, M, \quad (2)$$

where the reduced form error covariance matrix Σ_{s_t} is conditioned on the same process s_t as its structural counterpart Λ_{s_t} . The standard linear combination $\varepsilon_t = B^{-1}U_t$ gives the relation between structural and reduced form errors. The decomposition (2) imposes testable restrictions on the covariance matrices. In case of $M > 2$, it is possible to check whether the data is compatible with the decomposition and, thus, a time-invariant B can be used to transform reduced form errors into structural shocks. Lanne et al. (2010) shows that the model (1) and decomposition (2) give a unique (apart from ordering and sign) B (and thus ε_t) if structural shocks' variances are distinct across variables and states, i.e. for any two subscripts $k, l \in \{1, \dots, 5\}$, $k \neq l$, there is a $j \in \{2, \dots, M\}$ such that $\lambda_{jk} \neq \lambda_{jl}$.

Besides approving that the structural shocks are unique, orthogonal and heteroscedastic, the statistical procedure does not necessarily provide economically interpretable shocks. The motivation behind (1) is that the approach extracts not only statistically unique shocks but also decomposes the distinctive natures of the data Y_t . Since the distinguishing features of the data are the expectations horizons (economic content) and geographical source (US, EA and global), the different $\varepsilon_{t,k}$, $k = 1, 2, \dots, 5$ are meant to isolate some of these different characteristics. In accordance with implications of macroeconomic models presented by Bomfim and Rudebusch (2000) and Beechey et al. (2011) and related economic intuition, we evaluate the economic meaning of shocks in Section 4.⁵

Stemming on the conditional normality of the reduced form residuals, we estimate the MS-SVAR model via maximum likelihood. The full algorithm can be found in the Appendix B. Tests for statistical identification and confidence bands for impulse response function are computed as suggested in Lütkepohl and Netšunajev (2014).

4 US and EA inflation expectations spillovers

In this section, we document the model selection procedure and how we achieve the identification of US and EA specific and global shocks. Given the identified structural shocks, we study their impact on US and EA inflation expectations via impulse responses and variance decompositions.

⁵Due to the numerical complexity of the estimations and the weekly data frequency, our empirical strategy does not consider variables like GDP, inflation, unemployment or interest rates. We stress similarities of structural shocks derived from (1) with shocks of reduced rank SVARs and structural FAVARs. The basic idea is that shocks in (1) combine economic shocks. For interpretations see also Appendix C.

Table 1: Markov switching VAR model selection.

Model	$\log L_T$	AIC	SC
VAR(3) without MS	2012	-3834	-3458
MS(2)-VAR(3)	2343	-4466	-4030
MS(3)-VAR(3)	2451	-4652	-4157*
MS(4)-VAR(3)	2472	-4664*	-4109

Notes: L_T is the value of the likelihood function, $AIC = -2 \log L_T + 2 \times \#$ free parameters, $SC = -2 \log L_T + \log T \times \#$ free parameters. Sample: 2004 - 2012 ($T = 391$ obs.).

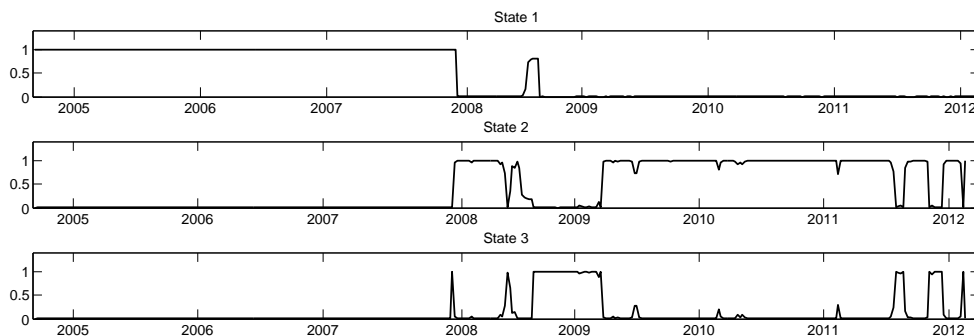
4.1 Model specification and identification

To specify an appropriate model for the identification of structural shocks from the inflation expectations and commodity price index, we first choose the lag length of a reduced form VAR with constant parameters for the whole sample period from 2004 to 2012. We follow the suggestion of the Schwarz criterion (SC) and continue with a VAR with three lags. We then implement the switching variance for different numbers of states M . Determination of the number of states by means of information criteria has been analyzed by Psaradakis and Spagnolo (2003, 2006). The information criteria are reported to perform well when the parameter changes are not too small. Building on these findings, we base the selection of M on the information criteria. Table 1 shows the log-likelihood and values of the Akaike (AIC) and SC for different models. Clearly, the likelihood is increasing in the flexibility of the model. We choose the model with three variance states since the MS(3)-VAR(3) is preferred by the SC.⁶

The estimated smoothed state probabilities of the MS(3)-VAR(3) model are shown in Figure 2. State 1 is the lowest volatility regime and State 3 the highest volatility regime. It can be seen that the first part of the sample until late 2007 is associated with state 1, while state 2 and 3 dominate the second part of the sample. The period since 2008 is well known to coincide with the global financial crisis, the global recession and the European sovereign debt crisis. We label state 1 as a non-crisis state and state 2 and 3 as crisis states. State 3 captures the timing of key events like the failure of the investment banks Bear Stearns (March 2008) and Lehman Brothers (September 2008), the home loan mortgage corporation Fannie Mae and Freddie Mac (July 2008) as well as the intensification of the European sovereign debt crisis, affecting Italy and Spain and coupled with increased banking sector strains,

⁶Note that with four regimes the statistical identification is not obvious and the states are more difficult to label. A two regime model provides the same qualitative results presented in the next subsection.

Figure 2: State probabilities of MS(3)-VAR(3) model.



Notes: Three volatility regimes: state 1 the lowest, state 3 the highest volatility.

Table 2: Tests for equality of structural variances across states.

H_0	LR statistic	p -value
$\lambda_{21} = \lambda_{22}, \lambda_{31} = \lambda_{32}$	15.6	0.00
$\lambda_{21} = \lambda_{23}, \lambda_{31} = \lambda_{33}$	36.8	0.00
$\lambda_{21} = \lambda_{24}, \lambda_{31} = \lambda_{34}$	12.9	0.00
$\lambda_{21} = \lambda_{25}, \lambda_{31} = \lambda_{35}$	25.7	0.00
$\lambda_{22} = \lambda_{23}, \lambda_{32} = \lambda_{33}$	41.0	0.00
$\lambda_{22} = \lambda_{24}, \lambda_{32} = \lambda_{34}$	2.95	0.23
$\lambda_{22} = \lambda_{25}, \lambda_{32} = \lambda_{35}$	48.8	0.00
$\lambda_{23} = \lambda_{24}, \lambda_{33} = \lambda_{34}$	26.4	0.00
$\lambda_{23} = \lambda_{25}, \lambda_{33} = \lambda_{35}$	48.8	0.00
$\lambda_{24} = \lambda_{25}, \lambda_{34} = \lambda_{35}$	43.2	0.00

Notes: Likelihood Ratio (LR) tests for equality of normalized structural variances ($\lambda_{s_t,k}$) across states $s_t = 2, 3$ and variables $k = 1, \dots, 5$ of MS(3)-VAR(3).

from mid-2011 on.

Statistical identification

As reviewed in Section 3, we aim at using the heteroscedasticity governed by the Markov states for identification purposes. Thus, the variances of shocks, $E(\varepsilon_{t(s_t),k}^2) = \lambda_{s_t,k}$, with shock number $k = 1, \dots, 5$, and volatility state $s_t = 2, 3$, have to be sufficiently distinct, see Section 3. We follow Lanne et al. (2010) and Lütkepohl and Netšunajev (2014) and verify that condition by pairwise LR-tests. Results presented in Table 2 indicate sufficient heterogeneity in the variances. It appears that with a p -value of 0.23 only shock $\varepsilon_{t,3}$ and $\varepsilon_{t,4}$ are difficult to distinguish. The point estimates reported in Table 4 show that $\lambda_{s_t,2}$ and $\lambda_{s_t,4}$ are relatively similar compared to other pairs. However, in state 2 point estimates differ by a factor of 2 and we will

demonstrate in the following that the two shocks have very different characteristics, thus, pose no problem for identification. Despite distinct variances, we check the validity of decomposition (2) by a LR-test. A p -value of 0.37 supports that the matrix of impact effects B of structural shocks can be considered state-invariant in the three state model. Hence, we conclude that the shocks are statistically identified. Note that for the model with four Markov states, preferred by the AIC (Table 1), these requirements for the statistical identification are not met.

Labeling of identified shocks

To verify whether we can attach some economic meaning and geographical origin to the statistically identified shocks, we first set up a regression study. In separate regressions, each structural shock is modeled as a dependent variable. The choice of explanatory variables is mainly motivated by Beechey et al. (2011) and the idea that a dominating force behind short horizon inflation expectations are demand, supply and monetary policy shocks, whereas long horizon expectations should be mostly insensitive to these shocks. Thus, distinct relations between structural shocks and certain US and EA proxies for demand, supply and monetary policy shocks (macro shocks in short) may support the economic interpretation. We utilize the difference between officially released economic outcomes and a respective expected value as a measure of macro shocks, compare e.g. Gürkaynak et al. (2005). The set of so-called macroeconomic surprise variables is uniform in all regressions and contains the surprise component of consumption expenditure, income, unemployment, GDP, industrial production, trade balance, inflation, productivity and monetary policy announcements for the US and EA, respectively. Market expectations are measured by using the Consensus mean forecast published by Bloomberg the Friday before each macroeconomic data release.⁷ In total we have five regressions with 18 surprise variables on a weekly basis. Our investigations focus on two joint parameter tests.

- H_0^{US} : $\beta_1 = 0$
US surprise variables have no explanatory power for $\varepsilon_{t,k}$,
- H_0^{EA} : $\beta_2 = 0$
EA surprise variables have no explanatory power for $\varepsilon_{t,k}$, $k = 1, \dots, 5$.

Table 3 shows the regression equation and test results. Further details and interpretations are provided in Appendix C. The regressions contribute to the economic interpretation of the structural shocks. The two tests reflect that shock 1 responds significantly to surprises about EA macroeconomic releases but not to US releases. Hence, we provisionally label structural shock 1 as a “EA macro shock”. We find that important drivers are releases about EA inflation and ECB monetary policy

⁷See further information about surprise variables and the regressions in Appendix C.

Table 3: Wald test from regressions with structural shocks on macro surprises (X).

$$\varepsilon_{t,k} = \beta_1' X_t^{(\text{US})} + \beta_2' X_t^{(\text{EA})} + u_t$$

Structural shock	F-test		R^2
	$H_0^{US} : \beta_1 = 0$	$H_0^{EA} : \beta_2 = 0$	
$\varepsilon_{t,1}$	0.78 (0.62)	13.7 (0.00)	0.10
$\varepsilon_{t,2}$	4.00 (0.00)	1.18 (0.31)	0.11
$\varepsilon_{t,3}$	0.66 (0.73)	0.71 (0.69)	0.04
$\varepsilon_{t,4}$	1.11 (0.36)	0.49 (0.86)	0.05
$\varepsilon_{t,5}$	3.28 (0.00)	8.58 (0.00)	0.06

Notes: Bold F-test statistics indicate rejections of H_0^j , $j=\text{EA}$, US. p -values are given in parentheses. Estimated standard deviations are based on HAC standard errors. Sample: 2004 - 2012 ($T = 388$ obs.).

announcements. In contrast to shock 1, for shock 2 the Null that US macroeconomic surprises have no impact is rejected, but we can not reject the insensitivity to EA surprises. Thus, for shock 2 we attach the provisional label “US macro shock”. Major drivers are news about US inflation, US industrial production and FED policy announcements. Shock 3 and 4 do not respond to either US and EA macroeconomic surprise components, so they appear as candidates for originating at longer expectations horizons. Following arguments in Bomfim and Rudebusch (2000) we label them “credibility shocks”. However, from the regression results we are not able to verify their origin.⁸ Finally, shock 5 responds to a mixture of US and EA releases, including components of FED and ECB policy announcements, US trade balance and US and EA industrial production. We propose to label the fifth structural shock a “global macro shock”, capturing the global component of news releases.

As indicated by the R^2 s the overall explanatory power of the regressions are rather low. This finding is not surprising in light of related regressions with break-even inflation rates by Gürkaynak et al. (2010b) and may reflect omitted surprise variables.⁹ It should be acknowledged that a different set of surprise variables may produce less clear-cut results. However, we pick most widely used and available data releases and find test results to be robust against moderate variations in the sample length.

⁸Note that the statistical identification is apart from ordering and sign. Thus, the ordering of structural shocks is not necessarily reflecting the ordering of variables in the data vector Y_t , see Section 3.

⁹Not that in the news regression context the omission is not likely to cause a bias of OLS- β estimates since, given the nature of forecast errors, surprise variables are usually mutually uncorrelated.

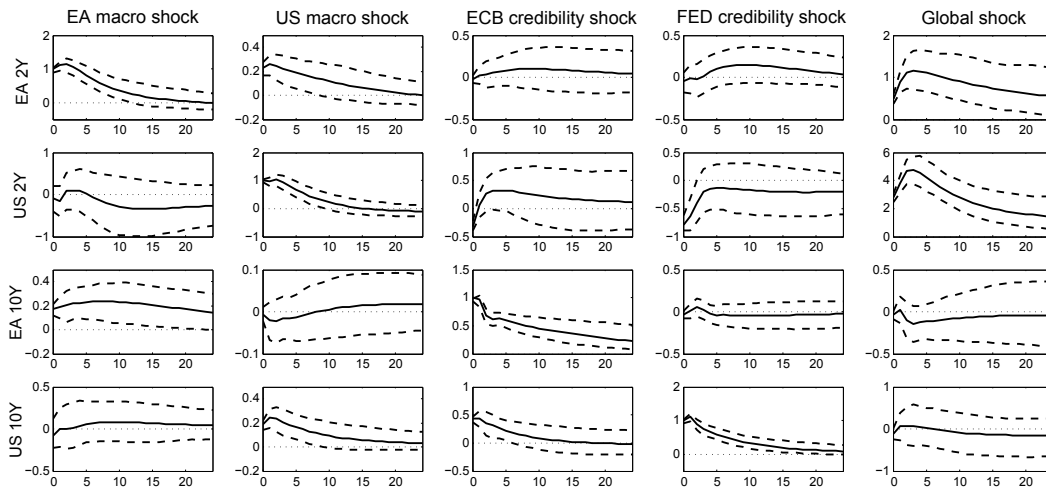
Table 4: Standard deviation and relative variance (λ) of structural shocks (ε).

State		Macro shock		Credibility shock	
		EA ($k=1$)	US ($k=2$)	ECB ($k=3$)	FED ($k=4$)
1	$\text{sd}(\varepsilon_{t(1),k})$	0.045	0.139	0.030	0.062
2	$\text{sd}(\varepsilon_{t(2),k})$	0.078	0.178	0.107	0.118
	$\lambda_{2,k}$	3.03 (0.58)	1.63 (0.32)	12.8 (2.69)	3.65 (0.70)
3	$\text{sd}(\varepsilon_{t(3),k})$	0.214	0.586	0.219	0.231
	$\lambda_{3,k}$	22.8 (7.88)	17.7 (4.29)	53.5 (15.0)	14.1 (4.09)

Notes: $\text{sd}(\varepsilon_{t(s_t),k})$ are standard deviations of structural shocks calculated via $\varepsilon_t = B^{-1}U_t$, with B^{-1} having normalized unit main diagonal elements. λ s in State 1 are normalized to one. Standard deviation of estimated λ s in parentheses.

To further support our economic labels, we report standard deviations and the relative variances (λ) of structural shocks in Table 4. Two characteristics distinguish US and EA specific macro shocks. First, US macro shocks are much larger in absolute size than EA macro shocks, e.g. the standard deviation in the non-crisis state in the US is almost three times larger than in the EA. Second, the relative variance of EA macro shocks increases stronger in crisis times (state 2 and 3). The increase relative to state 1 in the EA is 22.8%, in the US 17.7%. These two distinguishing characteristics of the US and EA macro shocks are also present in the US and EA macroeconomic surprise variables used as explanatory variables in the regressions, compare related descriptive statistics reported by Autrup and Grothe (2014). This gives further support to our economic labeling of the first two shocks. We carry over this result to the credibility shocks and label shock 3 with the smaller standard deviation and larger relative variances the “ECB credibility shock” and shock 4 the “FED credibility shock”. We define a credibility shock as a shock that can not be explained by macroeconomic surprise variables but should play an important role for long horizon inflation expectations. A favorable credibility shock moves inflation expectations in the direction of an officially announced or market perceived inflation target whereas an increasing variance of credibility shocks is interpreted as higher uncertainty about the target. In crisis periods, with central banks’ key interest rates at the zero lower bound and implemented non-standard policy measures, an increase in uncertainty about the credibility of inflation targets, as documented in Table 4, appear natural. The ECB credibility with an increase of the structural error’s variance of up to 53.5% is much stronger affected compared to the 14.1% increase in FED credibility shocks. One possible explanation for this gap is the more controversial discussions about ECB policy measures in the media, which may

Figure 3: Impulse responses of inflation expectations.



Notes: MS(3)-SVAR(3). Impulse responses to a one unit structural shock with 95% bootstrap confidence bands based on 1000 replications. Full sample period (391 obs.).

have affected financial market uncertainty to a large extent, see e.g. Springer (2012). Having in mind the labeling of US and EA specific shocks and the global shock, we continue the analysis of the MS-SVAR model.

4.2 Impulse response analysis

We study how the structural shocks affect the level of short and long horizon inflation expectations in the US and EA through normalized impulse responses. To save space, we focus on the inflation expectations measures and skip results for the commodity price index.¹⁰

The two-year inflation expectations

The first two rows of Figure 3 display how the US and EA specific and the global shocks transmit to short horizon inflation expectations. Similar to spillovers between US and EA financial markets, studied by Ehrmann et al. (2011), impulse responses indicate that US macro shocks significantly affect EA expectations (first row, second column) but not vice versa (second row, first column). The global macro shock has a significant and persistent impact on both US and EA inflation expectations (first

¹⁰To check the robustness of our results, we estimate classical SVARs identified via zero restrictions on contemporaneous effects for two subsamples (before and after 2008). Zero restrictions are chosen as indicated by the impulse responses of the MS-SVAR model. These restrictions are over-identifying and supported for the two subsamples by conventional tests. Most impulse responses of the separately estimated models are not significantly different for the two sample periods. Main economic conclusion can be supported.

and second row, fifth column). The credibility shocks appear to be less important at short expectations horizons. However, US expectations respond significantly to FED credibility shocks (second row, fourth column). The negative but fast decaying impact may reflect either distinct levels of perceived inflation targets at short and long horizons or a situation where the short inflation expectations are systematically above (below) a perceived target and long horizon expectations below (above) the target.¹¹

The ten-year inflation expectations

Responses of long horizon inflation expectations are depicted in the third and fourth row of Figure 3. In contrast to short horizon inflation expectations, long horizon inflation expectations are not significantly affected by the global macro shock. Domestic shocks appear to play an important role. With respect to the anchoring criteria defined by Gürkaynak et al. (2010b), impulse responses indicate that US and EA inflation expectations are strongly anchored with respect to foreign macro shocks (fourth row, first and fifth column; third row, second and fifth column). Domestic macro shocks have a significant impact (fourth row, second column; third row, first column). From the fast decaying impulse responses, we conclude that inflation expectations are firmly anchored in both the US and EA, compare Strohsal and Winkelmann (2015).

4.3 Variance decomposition

Having studied the impulse responses to structural shocks, we now turn to assess their relative importance for the variance of US and EA inflation expectations. Since variances change across the three Markov states, spillovers vary across the non-crisis (state 1) and crisis states (state 2 and 3). Spillovers are defined as the percentage of the US (EA) inflation expectations variance explained by both EA (US) shocks and the global shock. Results are summarized in Table 5.

The non-crisis period

In line with findings on actual inflation, our results show that short horizon inflation expectations are strongly affected by foreign structural shocks. In the non-crisis state 73.7% of the variance of US inflation expectations is explained by spillovers. For EA expectations the percentage is with 63.2% similarly large. Spillovers have the same order of magnitude as related global factors of actual inflation studied by Neely and

¹¹Note that the FED announced its official inflation target of 2% in 2012 – the ending of our sample period. The finding may reflect uncertainty about the level of the target. The negative impact appears quite robust across different specifications, e.g. in a two state MS-SVAR model and the model (robustness check) discussed in footnote 10.

Table 5: Variance decomposition of inflation expectations.

Volatility states	$\pi^e(\text{EA } 2\text{Y})$			$\pi^e(\text{US } 2\text{Y})$			$\pi^e(\text{EA } 10\text{Y})$			$\pi^e(\text{US } 10\text{Y})$		
	1	2	3	1	2	3	1	2	3	1	2	3
EA shocks	36.5	58.4	66.8	1.9	4.4	5.9	28.2	9.2	15.4	1.0	0.9	1.4
ECB credibility	0.3	2.2	1.4	0.5	4.6	3.4	65.1	89.6	83.2	3.2	12.1	11.0
US shocks	13.3	11.4	18.9	23.8	30.2	58.1	1.7	0.3	0.7	19.4	9.3	22.1
FED credibility	1.8	3.6	2.1	2.6	7.3	5.0	1.3	0.5	0.4	71.3	76.4	64.6
Global shock	48.1	24.4	10.8	71.3	53.5	27.6	3.6	0.4	0.2	5.0	1.4	0.9
Spillover	63.2	39.4	31.8	73.7	62.5	36.9	6.6	1.2	1.3	9.2	14.4	13.3

Note: Percentage of the inflation expectations variance explained by the structural shocks (ε) in States 1 (lowest volatility, non-crisis state) to state 3 (highest volatility, crisis state). Spillovers refer to aggregated contributions of foreign shocks. Calculated from forecast error variance decomposition at 100 weeks horizon.

Rapach (2011) and Mumtaz and Surico (2012). The SVAR model further reveals that for US short horizon expectations the main source of spillovers is the global macro shock (71.3%). EA specific shocks play with 2.4% only a marginal role for US expectations. On the contrary, EA inflation expectations are with 15.1% of its variance more exposed to US shocks. However, also for EA short horizon inflation expectations the global macro shock still plays the most significant role (48.1%).

At long expectations horizons the picture is materially different. Spillovers account for only 9.2% of the variance of US and 6.6% of EA inflation expectations. In line with ideas formalized by Bomfim and Rudebusch (2000) and Beechey et al. (2011), our results indicate that the credibility of central banks is the main driver of long horizon inflation expectations. The FED credibility shock accounts for 71.3% of US expectations while 65.1% of the variance of EA inflation expectations is explained by ECB credibility shocks. In the non-crisis state, transmissions of credibility shocks between the US and EA are negligible.

Crisis periods

The variance decomposition indicates that spillovers decrease in crisis times. We find that the main driver of the decline is the global macro shock whose impact on short horizon expectations drops down to one-third for the US and one-fourth for the EA. The decline is slightly compensated by increasing transmissions of US and EA specific shocks, e.g. the explanatory power of EA macro and ECB credibility shocks for US short horizon expectations increases up to 7 percentage points (state 1 vs. 3). Given the relative nature of the variance decomposition, decreasing spillovers imply an increasing role of domestic shocks. The variance decomposition of short horizon EA expectations provides an example. The variance explained by EA macro shocks almost doubles in crisis times and accounts for 66.8% in state 3. This pattern suggests that during the global financial crisis, the great recession and the European sovereign debt crisis the relative exposure of inflation expectations to foreign shocks has declined resulting in a stronger focus on the domestic economy.

The US long horizon expectations provide an exemption to the decreasing spillovers. US expectations are with a slight increase from 9.2% up to 14.4% stronger affected by foreign shocks during crisis times. The key driver of the elevated spillovers are the ECB credibility shocks, reflecting the impact of the European sovereign debt crisis on US expectations. Overall, the main result of strong international components in short horizon inflation expectations and rather local components in long horizon inflation expectations can be confirmed for the crisis period.

5 Conclusion

In this paper we explore the role of foreign and domestic structural shocks in determining US and EA inflation expectations. On the basis of shocks derived from a parsimonious structural vector autoregression, we propose a multi-step procedure to assess the economic nature and geographical source of structural shocks. We demonstrate that there is a significant difference between the structural drivers of short and long horizon expectations. Short horizon inflation expectations are closely linked to actual inflation and mainly respond to demand, supply and monetary policy shocks. Long horizon expectations are mostly determined by the credibility of central banks' inflation targets. Besides the economic content, we provide evidence that the structural shocks of the SVAR model separate into US specific, EA specific and global shocks. We find that cross country transmissions account for up to 73.7% of the variability of short horizon inflation expectations. This finding is consistent with previous literature on actual inflation rates, thus, further confirms a close link between short horizon inflation expectations and actual inflation rates. In contrast, long horizon inflation expectations are only explained by around 10% by cross country transmissions.

To our knowledge, theoretical models that rationalize underpinnings and implications of joint inflation expectations dynamics across countries and expectations horizons have not been formulated yet. We are confident that our results provide a good starting point for exploring the phenomenon of global expectations at short horizons but local expectations at long horizons.

The results in this paper provide further arguments for the discussion about globalized economies and a related weakening of inflation control by national monetary policy. Interestingly, the decrease in cross country transmissions during crisis times dampens global effects and may enhance the effectiveness of local measures. Finally, our results add empirical facts to the conventional wisdom that long horizon inflation expectations are most directly controllable by central banks via announcing and strengthening the credibility of an inflation target. Long horizon expectations appear as a convenient targeting variable of forward guiding strategies.

References

- Autrup, S.L. and Grothe, M. (2014), *Economic surprises and inflation expectations: Has anchoring of expectations survived the crisis?*, Working Paper Series 1731, European Central Bank.
- Bayoumi, T. and Swiston, A. (2010), *The Ties that Bind: Measuring International*

Bond Spillovers Using Inflation-Indexed Bond Yields, IMF Staff Papers No. 2 (57).

- Beechey, M. J., Johannsen, B. K. and Levin, A. (2011), *Are long-run inflation expectations anchored more firmly in the Euro Area than in the United States?*, American Economic Journal: Macroeconomics 3, 104-129.
- Bentolila, S., Dolado, J. and Jimeno, J.F. (2008), *Does immigration affect the phillips curve? Some evidence from Spain*, European Economic Review 52, 1398-1423.
- Bomfim, A.N. and Rudebusch, G.D. (2000), *Opportunistic and deliberate disinflation under imperfect credibility*, Journal of Money, Credit and Banking 32, 707-721.
- Castelnuovo, E. (2010), *Tracking U.S. inflation expectations with domestic and global indicators*, Journal of International Money and Finance 29, 1340-1356.
- Chen, L., Lesmond, D.A. and Wei, J. (2007), *Corporate yield spreads and bond liquidity*, Journal of Finance 62, 119-149.
- Christensen, J.H.E. and Gillan, J. (2012), *Do Fed TIPS purchases affect market liquidity?*, FRBSF Economic Letter 2012-07.
- Christensen, J.H.E., Lopez, J.A. and Rudebusch, G.D. (2010), *Inflation Expectations and Risk Premiums in an Arbitrage-Free Model of Nominal and Real Bond Yields*, Journal of Money, Credit and Banking 42, 143-178.
- Ciccarelli, M. and Mojon, B. (2010), *Global inflation*, Review of Economics and Statistics 92, 524-535.
- Clarida, R., Galí, J. and Gertler, M. (2000), *Monetary policy rules and macroeconomic stability: Evidence and some theory*, Quarterly Journal of Economics 115(1), 147-180.
- Darby, M.R. and Lothian, J.R. (1983), *The international transmission of inflation*, University of Chicago Press.
- Ehrmann, M., Fratzscher, M. and Rigobon, R. (2011), *Stock, bonds, money markets and exchange rates: Measuring international financial transmission*, Journal of Applied Econometrics 26, 948-974.
- Ehrmann, M., Pfajfar, D. and Santoro, E. (2014), *Consumer attitudes and the epidemiology of inflation expectations*, Bank of Canada Working Paper 2014-28.
- FED (2012), *Press release*, Release Date: December 12, 2012, [http : //www.federalreserve.gov/newsevents/press/monetary/20121212a.htm](http://www.federalreserve.gov/newsevents/press/monetary/20121212a.htm).
- Gürkaynak, R.S., Swanson, E.T. and Levin, A. (2005), *The sensitivity of long-term interest rates to economic news: Evidence and implications for macroeconomic models*, American Economic Review 95, Macroeconomics, 425-436.
- Gürkaynak, R.S., Swanson, E.T. and Levin, A. (2010a), *The TIPS yield curve and inflation compensation*, American Economic Journal: Macroeconomics, 70-92.

- Gürkaynak, R.S., Swanson, E.T. and Levin, A. (2010b), *Does Inflation Targeting Anchor Expectations? Evidence from the US, UK and Sweden*, Journal of the European Economic Association 8(6), 1208-1242.
- Henriksen, E., Kydland, F.E. and Sustek, R. (2013), *Globally correlated nominal fluctuations*, Journal of Monetary Economics 60, 613-631.
- Herwartz, H. and Lütkepohl, H. (2014), *Structural vector autoregressions with Markov switching: combining conventional with statistical identification of shocks*, Journal of Econometrics 183, 104-116.
- Hördahl, P. and Tristani, O. (2012), *Inflation risk premia in the term structure of interest rates*, Journal of the European Economic Association 10, 634-657.
- Krolzig H.M. (1997), *Markov-switching vector autoregressions: Modelling, statistical inference, and application to business cycle analysis*, Springer-verlag, Berlin.
- Lanne, M. and Lütkepohl, H. (2008), *Identifying monetary policy shocks via changes in volatility*, Journal of Money, Credit, and Banking 40, 1131-1149.
- Lanne, M., Lütkepohl, H. and Maciejowska K. (2010), *Structural vector autoregressions with Markov switching*, Journal of Economic Dynamics and Control 34, 121-131.
- Leduc, S., Sill, K. and Stark T. (2007), *Self-fulfilling expectations and the inflation of the 1970s: Evidence from the Livingston Survey*, Journal of Monetary Economics 54, 433-459.
- Lütkepohl, H. (2011), *Identifying structural vector autoregressions via changes in volatility*, DIW Discussion Papers 1259.
- Lütkepohl, H. and Netšunajev, A. (2014), *Disentangling demand and supply shocks in the crude oil market: How to check sign restrictions in structural VARs*, Journal of Applied Econometrics 29, 479-496.
- Melitz, M.J. and Ottaviano, G. (2008), *Market size, trade, and productivity*, Review of Economic Studies 75, 295-316.
- Mumtaz, H. and Surico, P. (2012), *Evolving international inflation dynamics: world and country-specific factors*, Journal of the European Economic Association 10, 716-734.
- Neely, C.J. and Rapach, D.E. (2011), *International comovements in inflation rates and country characteristics*, Journal of International Money and Finance 30, 1471-1490.
- Pflüger, C. and Viceira, L. (2012), *An empirical decomposition of risk and liquidity in nominal and inflation-linked government bonds*, preprint.
- Psaradakis, Z. and Spagnolo, N. (2003), *On the determination of the number of regimes in Markov-switching autoregressive models*, Journal of Time Series Analysis, 24, 237-252.

- Psaradakis, Z. and Spagnolo, N. (2006), *Joint determination of the state dimension and autoregressive order for models with Markov regime switching*, Journal of Time Series Analysis, 27, 753-766.
- Rigobon, R. (2003), *Identification through heteroskedasticity*, Review of Economics and Statistics 85, 777-792.
- Söderlind, P. (2011), *Inflation risk premia and survey evidence on macroeconomic uncertainty*, International Journal of Central Banking, 113-133.
- Springer (2012), *Is the European Central Bank losing credibility? A symposium of views*, International Economy Magazine Spring 2012 issue, 20-30.
- Strohsal, T. and Winkelmann, L. (2015), *Assessing the anchoring of inflation expectations*, Journal of International Money and Finance 50, 33-48.

A BEI rates and inflation expectations

We adjust BEI rates by regressing them on measures of a liquidity premium. The liquidity adjusted BEI rate is then given by the residual of this regression. If $Liqu_t$ contains the liquidity measures and π_t^e represents the adjusted BEI rate, we have $\pi_t^e = BEI_t + \delta' Liqu_t$, where δ is a vector of coefficients. This is a common approach which is used, among others, by Chen et al. (2007) who estimate liquidity premia in corporate yield spreads and Gürkaynak et al. (2010a), Pflüger and Viceira (2012) or Autrup and Grothe (2014) who apply this approach to the yield spread of nominal and inflation-indexed government bonds.

Our choice of liquidity measures is mainly motivated by the discussion in Christensen and Gillan (2012). The first country-specific as well as horizon-specific measure is given by the spread between AAA rated corporate bond yields and nominal government bond yields.¹² The second liquidity measure captures the volatility processes in the markets. For the US, authors such as Söderlind (2011) or Christensen and Gillan (2012) propose to take the implied volatility of S&P 500 index options, i.e. the VIX index.¹³ In line with these papers, we utilize the VIX for US BEI rates and the implied volatility of EURO STOXX index options (VSTOXX) for EA BEI rates. The regression results are shown in Table 6.

We experiment with other variables like the GARCH variance but find that the dynamics of inflation expectations do not change much. Short horizon BEI rates are with R^2 s of 0.65 (EA) and 0.72 (US) stronger adjusted than long horizon BEI rates ($R^2 = 0.06$ for both US and EA BEI rates). The liquidity adjustments reduce the correlations between BEI rates.

B MS-SVAR estimation steps

The appendix describes the expectation maximization (EM) algorithm for the Markov switching SVAR model including parameter choices for the empirical application. The notation is based on Krolzig (1997) and Herwartz and Lütkepohl (2014).

• Definitions

¹²Since the credit risk component of AAA corporate bond yields is considered very small, the yield spread reflects the premium for the lower trading volume and larger bid-ask spreads in the corporate bond market versus the highly liquid Treasury bond market, see Christensen and Gillan (2012). AAA corporate bond yields for a three and ten year maturity horizon are obtained from the Macrobond database.

¹³The argument goes that future resale prices of a security and the liquidity premium that investors demand to guard against such risk move closely with the market volatility, see Christensen and Gillan (2012).

Table 6: Regressions for liquidity adjustments of BEI rates.

	EA		US	
	2Y	10Y	2Y	10Y
<u>AAA-spread:</u>				
EA(3Y), US (3Y)	-0.51 (0.07)	-	-0.36 (0.10)	-
EA(10Y), US(10Y)	-	-0.18 (0.08)	-	-0.11 (0.04)
<u>Volatility:</u>				
VSTOXX	-0.01 (0.00)	0.01 (0.00)	-	-
VIX	-	-	-0.07 (0.01)	0.01 (0.00)
R^2	0.65	0.06	0.72	0.06

Notes: AAA-spread is the difference between bond yields of triple A rated corporations and a government bond yield for the respective country and maturity horizon. VSTOXX and VIX are implied volatilities of options on the EURO STOXX 50 and S&P 500 index, respectively. Sample period Sep 2004 to March 2012. HAC standard errors of the estimated coefficients are given in parentheses.

- The baseline model is a VAR(p) of the form:

$$y_t = v + A_1 y_{t-1} + \dots + A_p y_{t-p} + u_t,$$

with $t = 1, \dots, T$ and y_t of dimension K .

- $\xi_t = \left(I(s_t = 1) \dots I(s_t = M) \right)'$,

$$E(\xi_t) = \left(Pr(s_t = 1) \dots Pr(s_t = M) \right)'$$

with states $s_t = 1, \dots, M$ and $I(\cdot)$ an indicator function which takes value 1 if the statement in the argument is true and 0 otherwise.

$\xi_{t|s} = E(\xi_t | Y_s) = \left(Pr(s_t = 1 | Y_s) \dots Pr(s_t = M | Y_s) \right)'$, with $Y_s = (y_1, \dots, y_s)$ and P the transition matrix, which yields $\xi_{t+1|t} = P \xi_{t|t}$, for $t = 0, 1, \dots, T-1$.

- $\eta_t = \left(f(y_t | s_t = 1, Y_{t-1}) \dots f(y_t | s_t = M, Y_{t-1}) \right)'$,

where $f(\cdot)$ is the conditional distribution function:

$$f(y_t | s_t = m, Y_{t-1}) = (2\pi)^{-K/2} \det(\Sigma_m)^{-1/2} \exp(-0.5 u_t' \Sigma_m^{-1} u_t),$$

and covariance matrices have decomposition $\Sigma_1 = BB'$, $\Sigma_m = B\Lambda_m B'$ for $m = 2, \dots, M$.

Notation:

⊙ elementwise multiplication,

⊘ elementwise division,

⊗ Kronecker product,

I_K is a $K \times K$ dimensional identity matrix,

$1_M = (1, \dots, 1)'$ is a $M \times 1$ dimensional vector of ones,

$\theta = \text{vec}(v, A_1, A_2, \dots, A_p)$ is the parameter vector,

$Z'_{t-1} = (1, y'_{t-1}, y'_{t-2}, \dots, y'_{t-p})$ is the matrix of ones and lagged observations.

- **Initial values**

The following starting values are used for the iterations:

- $P = M^{-1} \mathbf{1}_M \mathbf{1}'_M$
- $\hat{\theta} = \text{vec}(\hat{v}, \hat{A}_1, \dots, \hat{A}_p) = \left[\sum_{t=1}^T Z_{t-1} Z'_{t-1} \otimes I_K \right]^{-1} \sum_{t=1}^T (Z_{t-1} \otimes I_K) y_t$
- $B = T^{-1} \left(\sum_{t=1}^T \hat{u}_t \hat{u}'_t \right)^{1/2} + B_0$, where $\hat{u}_t = y_t - (Z'_{t-1} \otimes I_K) \hat{\theta}$ and B_0 is a matrix of random numbers coming from standard normal distribution and scaled by a factor of 10^{-5} .
- $\Lambda_1 = I_K, \Lambda_m = c_m I_K, m = 2, \dots, M$ with $c_2 = 0.4, c_3 = 0.16$ for this application.
- $\xi_{0|0} = M^{-1} \mathbf{1}_M$

- **Expectation step**

For given $P, \theta, \Sigma_m, m = 1, 2, \dots, M$ and $\xi_0 = \xi_{0|0}$ the following parameters are computed:

- η_t for $t = 1, 2, \dots, T$,
- $\xi_{t|t} = \frac{\eta_t \odot P \xi_{t-1|t-1}}{\mathbf{1}'_M (\eta_t \odot P \xi_{t-1|t-1})}$, for $t = 1, 2, \dots, T$.
- $\xi_{t|T} = (P' (\xi_{t+1|T} \otimes P \xi_{t|t})) \odot \xi_{t|t}$, for $t = T-1, \dots, 0$.
- $\xi_{t|T}^{(2)} = \text{vec}(P') \odot ((\xi_{t+1|T} \otimes P \xi_{t|t}) \otimes \xi_{t|t})$, for $t = 1, \dots, T-1$.

- **Maximization step**

- Estimate P :

$$\text{vec}(\hat{P}') = \left(\sum_{t=0}^{T-1} \xi_{t|T}^{(2)} \right) \odot \left(\mathbf{1}_M \otimes (\mathbf{1}'_M \otimes I_M) \sum_{t=0}^{T-1} \xi_{t|T} \right)$$
- Estimate B and Λ_m :
Define $T_m = \sum_{t=1}^T \xi_{mt|T}$, where $\xi_{mt|T}$ denotes the m -th element of the vector $\xi_{t|T}$. Estimation of B and Λ_m is done by minimizing the likelihood function:
 - $l(B, \Lambda_2, \dots, \Lambda_M) = T \log \det(B) + \frac{1}{2} \left(B'^{-1} B^{-1} \sum_{t=1}^T \xi_{1t|T} \hat{u}_t \hat{u}'_t \right) + \sum_{m=2}^M \left[\frac{T_m}{2} \log \det(\Lambda_m) + \frac{1}{2} \text{tr} \left(B'^{-1} \Lambda_m^{-1} B^{-1} \sum_{t=1}^T \xi_{mt|T} \hat{u}_t \hat{u}'_t \right) \right]$.
- Then compute:

$$\hat{\Sigma}_1 = \hat{B} \hat{B}', \hat{\Sigma}_m = \hat{B} \hat{\Lambda}_m \hat{B}' \text{ for } m = 2, \dots, M$$

- Estimates of the parameter vector θ are given by:

$$\hat{\theta} = \left[\sum_{m=1}^M \left(\sum_{t=1}^T \xi_{mt|T} Z_{t-1} Z'_{t-1} \right) \otimes \hat{\Sigma}_m^{-1} \right]^{-1} \sum_{t=1}^T \left(\sum_{m=1}^M \xi_{mt|T} Z_{t-1} \otimes \hat{\Sigma}_t^{-1} \right) y_t$$

- Initial regime probabilities are updated according to:

$$\xi_{0|0} = \xi_{0|T}$$

- **Convergence criteria**

Relative change in the value of the log-likelihood function is used as convergence criteria. The log-likelihood is evaluated for given $P, \theta, \Sigma_m, m = 1, 2, \dots, M$ and $\xi_{0|0}$ as follows. Compute:

- η_t for $t = 1, 2, \dots, T$,
 - $\xi_{t|t-1} = P\xi_{t-1|t-1}$, for $t = 1, 2, \dots, T$,
 - $\xi_{t|t} = \frac{\eta_t \odot P\xi_{t|t-1}}{\mathbb{1}'_M (\eta_t \odot P\xi_{t|t-1})}$, for $t = 1, 2, \dots, T$.
 - Then $\log L_T = \sum_{t=1}^T \log f(y_t | Y_{t-1})$,
- $$f(y_t | Y_{t-1}) = \sum_{m=1}^M Pr(s_t = m | Y_{t-1}) f(y_t | s_t = m, Y_{t-1}) = \xi'_{t|t-1} \eta_t.$$

Estimation of B , Λ_m and θ are iterated until convergence, i.e. relative change Δ in the log-likelihood is negligibly small (does not exceed tolerance value $\alpha = 10^{-9}$) for j -th and $(j - 1)$ -th rounds of iterations: $\Delta = \frac{\log L_T(j) - \log L_T(j-1)}{\log L_T(j-1)} < \alpha$.

- **Bootstrapping confidence bands for impulse responses**

Herwartz and Lütkepohl (2014) discuss a fixed design wild bootstrap procedure for constructing confidence intervals for impulse responses in the presently considered model class. The bootstrap samples are constructed as

$$y_t^* = \hat{v} + \hat{A}_1 y_{t-1} + \dots + \hat{A}_p y_{t-p} + u_t^*$$

where $u_t^* = \zeta_t \hat{u}_t$ and ζ_t is a random variable taking values 1 and -1 , each with probability 0.5. We bootstrap parameter estimates θ^* , B^* and Λ^* conditionally on the initially estimated transition probabilities.

C Regressions with macroeconomic surprise variables

To support the interpretation and labeling of structural shocks, we regress the shocks on the surprise component of macroeconomic data releases (including monetary policy announcements). The surprise component for each data release is computed as the released value less the mean of market expectations. Market expectations refer

to financial market experts (mostly bankers) asked the Friday before each data release. For each US data release around 50 experts contribute to the survey, while for EA releases the number is around 30. Data is provided by Bloomberg. Following Gürkaynak et al. (2005), we assume that the surprise component of macroeconomic data releases (i.e. forecast errors) provides a measure of demand, supply and monetary policy shocks.¹⁴ Weekly shocks are regressed on weekly surprise variables. Surprises are zero in cases of no releases. Regressions capture surprises about the following macro variables:

- US: Gross domestic product (GDP), industrial production (IP), urban consumer price index (CPI), unemployment rate (UEM), Output per hour (Productivity), trade balance of goods and services (Trade), consumer credit (CCredit), personal income (Income), federal funds target rate (MP).
- EA: Gross domestic product (GDP), industrial production (IP), harmonized consumer price index (CPI), unemployment rate (UEM), labor costs (Productivity), trade balance with non eurozone (Trade), consumption expenditure (CExp), government final consumption expenditure (GovC), ECB main refinancing rate (MP).

Regression results are provided in Table 7. The first two structural shocks $\varepsilon_{t,1}$ and $\varepsilon_{t,2}$ respond mainly to EA and US surprises, respectively. $\varepsilon_{t,3}$ and $\varepsilon_{t,4}$ are mostly invariant and $\varepsilon_{t,5}$ responses to a mixture of US and EA surprises. Labels are attached as discussed in Section 4, compare also results of F-tests reported in Figure 3.

Shocks and regression results have the following interpretations: Significantly negative coefficient estimates indicate a reverse relation between the sign of ε_t and the respective proxies of shocks. E.g. a positive supply shock working through an unexpected increase in US productivity results in a negative $\varepsilon_{t,2}$, thus, as indicated by impulse responses in Figure 3, decreases US inflation expectations. In contrast, an unexpected increase in US consumer credit has mainly the effect of a demand shock and results in a positive $\varepsilon_{t,2}$ and increases inflation expectations. Monetary policy shocks triggered by an unexpected increase in the federal funds target rate (ECB main refinancing rate) relate to negative structural shocks thus decrease US (EA) inflation expectations. This finding is inline with impulse responses of US inflation expectations (Livingston Survey, eight-month forecast horizon) to monetary policy shocks studied by Leduc et al. (2007). The regressions provide evidence that the MS-SVAR model extracts US specific, EA specific and global shocks. Especially $\varepsilon_{t,1}$ and $\varepsilon_{t,2}$ combine demand, supply and monetary policy shocks.

¹⁴Note that surprise variables share common characteristics with the structural shocks since both are centered, uncorrelated and heteroscedastic.

Table 7: Regressions of structural shocks on macro surprises.

Macro surprise	Structural shocks of MS-SVAR model				
	$\varepsilon_{t,1}$	$\varepsilon_{t,2}$	$\varepsilon_{t,3}$	$\varepsilon_{t,4}$	$\varepsilon_{t,5}$
<u>US:</u>					
· MP	-12.5 (0.00)	-6.54 (0.00)	4.02 (0.10)	-1.79 (0.50)	-1.34 (0.00)
· Trade	0.08 (0.40)	0.04 (0.48)	0.06 (0.29)	0.08 (0.18)	0.01 (0.09)
· Productivity	0.67 (0.11)	-0.39 (0.08)	-0.29 (0.13)	-0.30 (0.47)	-0.03 (0.86)
· CPI	-0.42 (0.84)	1.83 (0.03)	-0.14 (0.93)	-1.60 (0.40)	0.50 (0.30)
· GDP	0.24 (0.61)	0.16 (0.50)	-0.12 (0.77)	-0.60 (0.24)	-0.35 (0.14)
· IP	0.55 (0.22)	0.74 (0.03)	-0.54 (0.24)	1.57 (0.34)	0.89 (0.05)
· UEM	-2.81 (0.42)	0.73 (0.28)	-0.70 (0.66)	0.57 (0.57)	-0.18 (0.78)
· CCredit	0.01 (0.91)	0.04 (0.09)	-0.06 (0.21)	0.01 (0.70)	-0.02 (0.26)
· Income	-0.35 (0.57)	0.47 (0.31)	0.74 (0.36)	0.11 (0.86)	0.24 (0.70)
<u>EA:</u>					
· MP	-2.57 (0.01)	-0.06 (0.86)	-1.86 (0.19)	-0.27 (0.70)	-0.05 (0.09)
· Trade	-0.02 (0.90)	0.02 (0.71)	-0.11 (0.33)	0.02 (0.84)	-0.01 (0.94)
· Productivity	-0.17 (0.85)	-0.87 (0.10)	1.58 (0.17)	1.06 (0.25)	-0.34 (0.29)
· CPI	-11.6 (0.01)	1.72 (0.38)	7.71 (0.13)	-1.47 (0.69)	-0.71 (0.77)
· GDP	-3.65 (0.10)	-3.65 (0.11)	1.15 (0.42)	0.61 (0.67)	-0.71 (0.33)
· IP	2.04 (0.01)	0.14 (0.53)	-0.09 (0.80)	-0.04 (0.90)	0.26 (0.02)
· UEM	0.67 (0.82)	0.87 (0.44)	-1.43 (0.62)	0.04 (0.98)	0.44 (0.72)
· CExp	3.31 (0.01)	-3.26 (0.14)	-1.12 (0.47)	-1.38 (0.11)	0.73 (0.44)
· GovC	4.58 (0.00)	-1.62 (0.12)	0.04 (0.96)	-1.41 (0.05)	-1.22 (0.39)
R^2	0.10	0.11	0.04	0.05	0.06

Notes: Variable explanation see text. Sample period 2004 to 2012 (388 Obs.). P-values based on HAC standard errors are given in parentheses.