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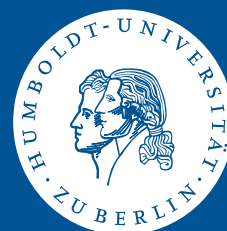


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Abstract

We quantify spillovers of inflation expectations between the United States (US) and Euro Area (EA) based on break-even inflation (BEI) rates. In contrast to previous studies, we model US and EA BEI rates jointly in a structural vector autoregressive (SVAR) model. The SVAR approach allows to identify US and EA specific inflation expectations shocks. By modeling the heteroscedasticity of the data, we are able to test the identifying restrictions of structural shocks and analyze time-varying spillovers. Adjusted for BEI risk premia, our main result suggests that spillovers of inflation expectations increase during times of macroeconomic stress. We document a significant impact of the European sovereign debt crisis on US expectations. The finding contributes to the discussion about a weakening of inflation control by national central banks and speaks in favor of internationally coordinated policy actions, especially during crisis times.

Keywords: International transmissions, break-even inflation, credibility of monetary policy, structural vector autoregressive (SVAR) analysis, identification through heteroskedasticity.

JEL classification: E31, F42, E52.

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1 Introduction

Private agents' expectations play a key role in modern monetary policy. In particular, inflation expectations have become a widely recognized variable, whose successful control is well known to facilitate greater stability of output, employment and prices, see among others Bomfin and Rudebusch (2000) and Orphanides and Williams (2005). For that reason, central banks like the Federal Reserve Bank (FED) and the European Central Bank (ECB) commit to anchored inflation expectations and have recently chosen inflation expectations as an important variable of their forward guidance strategy on the key policy rate, e.g. FED (2012) and ECB (2013).¹ Yet, in spite of their prominent role, the nature of inflation expectations and its transmission channels are not fully understood.

Empirical work on inflation expectations takes a country and central bank specific perspective. Cross country differences of anchored inflation expectations are usually considered in reduced form models where the countries are treated in isolation to each other, see e.g. Gürkaynak et al. (2010), Jochmann et al. (2010), Beechey et al. (2011), Cruikshank and Demertzis (2011) or Strohsal and Winkelmann (2012). Thus, expectation formation processes are assumed to be independent across countries and spillovers are neglected. However, recent empirical findings on the actual rate of inflation by e.g. Ciccarelli and Mojon (2010) or Mumtaz and Surico (2012) show that national inflation rates are strongly driven by a *global* inflation factor. The presence of global inflation suggests that also inflation expectations are connected across countries. Since little is known about the strength and scope of these linkages, it appears of crucial importance to reveal these spillovers.

In this paper, we quantify time-varying spillovers of inflation expectations between the United States (US) and Euro Area (EA). Our focus is on country specific inflation expectation shocks, originating in one country and maturity horizon and transmitting to foreign inflation expectations. To identify the country specific shocks, we employ the Markov switching structural vector autoregressive (MS-SVAR) model of Lanne et al. (2010). The idea is to use changes in the volatility of shocks to support the identification of the model. In comparison with GARCH type models applied in Normandin and Phaneuf (2004) or Bouakez and Normandin (2010), an attractive feature of the MS model is the possibility of economic interpretation of the volatility regimes. Furthermore, the MS model captures GARCH effects because a specific period may not be associated with a single volatility state but be a combination of different states. Utilizing information from changes in volatility appears promising since contemporaneous transmissions are highly relevant in our assessments and the identification via heteroscedasticity avoids setting ad hoc re-

¹In FED (2012) the FED defined anchored longer-term inflation expectations as well as medium term inflation expectations not larger than 2.5% as two explicit criteria to maintain their overnight interest rate at the zero lower bound.

restrictions on simultaneous effects.² As illustrated in Lütkepohl and Netšunajev (2013) we check identifying restrictions by formal statistical tests. The identified structural model provides the country specific inflation expectations shocks and allows the analysis of its cross country transmissions. Main results about time-varying spillovers are derived from a state dependent variance decomposition governed by the Markov switching structure.

Our multivariate setup can be considered as an extension of the single country *pass through* models of Jochmann et al. (2010), Gefang et al. (2012), Lemke and Strohsal (2013) or Dräger and Lamla (2013).³ As the previous studies, we utilize medium and long term expectations horizons and model them via forwards of five and ten year break-even inflation (BEI) rates.⁴ We are interested in both expectation horizons since they provide important information for the structural model. Medium term expectations carry information about country specific economic news, while long term inflation expectations are mainly affected by central banks' credibility in controlling inflation, compare Beechey et al. (2011), Bomfin and Rudebusch (2000) and Ciccarelli and García (2009). Our main goal is to derive the country specific shocks from US and EA BEI rates and to study their cross country spillovers. To account for global risk shocks that can not be assigned to a specific country, we include the VIX volatility index as a fifth variable. In general, global risk shocks may capture responses to oil or commodity price shocks.⁵ More specifically, the VIX is also known to explain market risk premia associated with BEI rates, see Söderlind (2011) or Christensen and Gillan (2012).⁶ Including the risk measure in the analysis of the structural model turned out to provide an effective and transparent approach to account for BEI risk premia. The identification of a risk shock allows a presentation of the main results in terms of risk adjusted BEI rates.

Empirical results are based on weekly data in a time period from 2004 to 2011. The detected Markov states enable us to endogenously distinguish between spillovers in non-crisis and crisis states. The key result of the paper is that spillovers of US and EA inflation expectations shocks play a significant role and increase during times of financial and macroeconomic stress. US inflation expectations are stronger affected by EA shocks than EA expectations by US shocks. In the crisis state 26% of the variance of US long

²E.g. avoids ad hoc exclusions of variables (Cholesky), sign or long-run restrictions.

³Pass through models are usually employed to assess the anchoring of inflation expectations. Inflation expectations are considered to be anchored if the long term expectations are not significantly affected by medium term expectations.

⁴A BEI rate is the spread between a nominal and inflation linked bond, also known as inflation compensation.

⁵Note that macroeconomic consequences of oil price shocks have usually shorter cycles than five years, see e.g. Lippi and Nobili (2012). Ciccarelli and García (2009) find that oil and commodity prices do not explain EA BEI rates.

⁶BEI rates are not a pure measure of inflation expectations but include liquidity as well as inflation risk premia, see Christensen et al. (2010) among others.

horizon inflation expectations can be explained by EA inflation expectations shocks, an increase by around 10 percentage points compared with the non-crisis state. In contrast, spillovers from the US to the long horizon EA expectations are much more stable across states and account for around 6% of the total variance only. The driving force behind the stronger spillovers from the EA to the US are mainly the EA ten year inflation expectations shocks. In the crisis states, the variance of EA inflation expectations shocks increase almost twice as much as the variance of US expectations shocks. Interpreting the ten year shocks as credibility shocks, we find that there is a significantly stronger increase in uncertainty about the ECB's credibility in crisis states than uncertainty about the Federal Reserve. While the European sovereign debt crisis appears to induce the uncertainty in the EA and transmits to US inflation expectations, uncertainty originating in the US during the crisis periods have only relatively small effects on EA inflation expectations.

The debate about increasing international components in domestic inflation rates and a related weakening of inflation control by national central banks (Galí, 2010, Henriksen et al., 2011, Neely and Rapach, 2011, or Mumtaz and Surico, 2012) can be extended to inflation expectations. Our finding that spillovers of inflation expectations increase for higher market uncertainty, rationalizes central banks' incentives to coordinate policy actions at an international level during times of market stress. It justifies discussion and advice between central banks about appropriate policy measures, especially during crisis times.

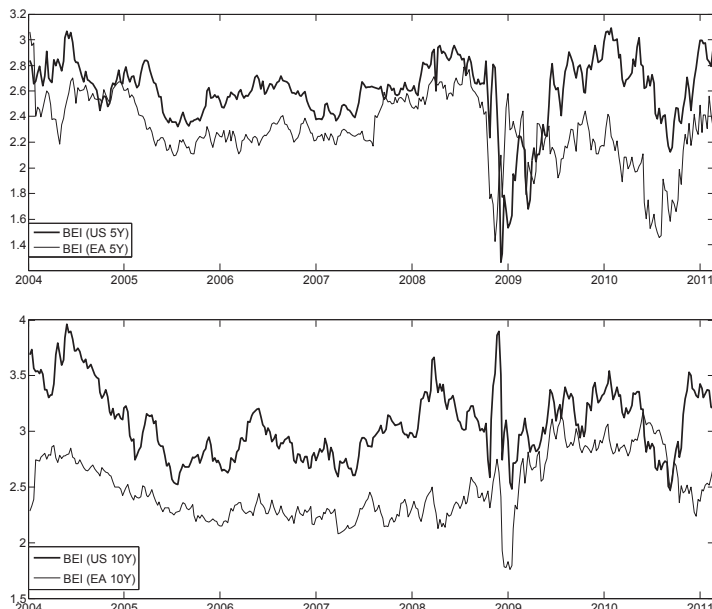
The rest of the paper is organized as follows. The upcoming Section 2 presents the inflation expectations and risk data. Section 3 introduces the Markov switching SVAR model. The main part is Section 4. We first describe the estimation and identification of our model and then present the results on the spillovers of US and EA inflation expectations shocks. Section 5 concludes.

2 Data

Inflation expectations are usually not directly observable. For that reason, empirical studies on inflation expectations rely on either measures based on inflation surveys or measures derived from financial market instruments. Financial market measures are most important, as they provide timely information about inflation expectations over a large number of expectations horizons, see Galati et al. (2011).

In this paper, we refer to the Fisher equation, thus, utilize the spread between yields of nominal and real (inflation linked) government bonds to measure inflation expectations. The spread is known as break even inflation (BEI) or inflation compensation. Following the pass through literature of e.g. Jochmann et al. (2010), we focus on medium and long

Figure 1: 5 and 10 year BEI rates.



Notes: Weekly averages (Monday to Friday) of daily one year forward break-even inflation (BEI) rates. BEI rates are derived from nominal and real Nelson-Siegel-Svensson yield curves, see Strohsal and Winkelmann (2012). 5 year expectation horizon (upper figure), 10 year expectation horizon (lower figure). 372 weekly observations.

term expectations horizons. Expectations are modeled by one-year forward BEI rates five and ten years ahead. The five-year horizon is meant to reflect today's expectations in five years for one year. The ten-year horizon provides a measure of today's expectations from year ten to year eleven.⁷ We study weekly data of the United States (US) and Euro Area (EA) in the time period from January 2004 to February 2011 providing 372 observations. Figure 1 illustrates the sample paths. The mean of the BEI rates is around 2.5 percent and slightly higher for the US ten year horizon. Conventional unit root tests suggest that the BEI rates are stationary. Despite the stationarity, a heteroscedastic pattern is clearly visible in the sample paths. We take advantage of the heteroscedasticity to identify inflation expectations shocks at the five and ten year expectations horizons, see Section 3. Including five and ten year expectations provides crucial information for our structural analysis. Standard macroeconomic models suggest that inflation expectations at a medium horizon just respond to macroeconomic shocks, see the discussion by Beechey et al. (2011) and translations to pass through regressions by Dräger and Lamla (2013). Thus, the five year BEI rates incorporate information regarding economic news effects just relevant

⁷The data is taken from Strohsal and Winkelmann (2012), where more details about the data are given.

for revisions at medium term expectations horizons. In contrast, long term expectations horizons are less sensitive to economic news, e.g. Gürkaynak et al. (2010) or Strohsal and Winkelmann (2013). Referring to Bomfin and Rudebusch (2000), revisions of inflation expectations ten years ahead comprise market perceptions regarding a central bank’s ability to stabilize prices. An important driver of ten year BEI rates is, therefore, the credibility of monetary policy.

With our structural model of the two largest economies in the world, we aim at identifying country specific (structural) inflation expectations shocks from the BEI rates. Since the dynamics of the BEI rates may not be fully explained by country specific inflation expectations shocks, we include the VIX volatility index to account for common global risk factors.⁸ On the one hand, the global risk shocks may capture commodity price shocks or inflation expectations shocks originating in other countries than the US or EA.⁹ On the other hand, the VIX is known to explain liquidity and inflation risk premia of US and EA BEI rates, see Söderlind (2011), Galati et al. (2011) or Christensen and Gillan (2012). Incorporating the VIX in the structural analysis is meant to support the identification of the US and EA specific inflation expectations shocks and to enable endogenous risk adjustments of the BEI rates.

3 The Markov switching structural vector autoregressive model

In the present paper we focus on the information available in BEI rates and VIX to study the spillovers of inflation expectations shocks across the US and EA. The idea is to use changes in the volatility of the BEI rates (apparent in Figure 1) and the VIX to support the identification of country specific inflation expectations shocks.

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 and allows to support imposed restrictions by formal statistical tests, compare Lütkepohl and Netšunajev (2013). The model deployed in the analysis is a structural vector autoregression with a time-varying variance of the structural shocks. The changes in the volatility are determined from the data by a Markov regime-switching (MS) mechanism. This ap-

⁸The VIX is computed by the Chicago Board of Options Exchange. It equals the square root of a variance swap written on a portfolio of out-of-the-money options on the S&P 500 index.

⁹Note that the dimension of our SVAR model with 5 variables is already relatively large. Thus, incorporating additional variables in the system to closer capture different sources of shocks may lead to numerical problems and harms the identification of the structural model.

proach was proposed by Lanne et al. (2010) and used by Herwartz and Lütkepohl (2011), Lütkepohl and Netšunajev (2013), and Netšunajev (2013) for different applications.

The reduced form of our $K = 5$ variable VAR model with p lags is

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

where ν is a constant intercept and the A_j s ($j = 1, \dots, p$) are 5×5 coefficient matrices. The reduced form errors U_t are assumed to depend on 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 $U_t | s_t \sim N(0, \Sigma_{s_t})$. The heteroscedastic nature, modeled via the state-dependence of the covariance matrix Σ , provides crucial information for the identification of the structural counterpart of (1). It is worth mentioning, that to make the identification work, other parameters than the residual covariance need to be state-independent, compare also e.g. Rigobon (2003).

The structural model is obtained by substituting U_t by $B\varepsilon_t$ in (1), where B is the matrix of contemporaneous effects of structural shocks ε_t on the variables Y_t . In absence of heteroscedasticity, B is the classic object for imposing just-identifying (short run) restrictions. Structural shocks in the model are orthogonal and are normalized such that they have unit conditional variance in the first state. The diagonal matrix Λ_i , $i = 2, \dots, M$ of the relative variances of structural shocks and the reduced form error covariance matrix Σ_{s_t} are conditioned on the same process s_t ¹⁰. The relation

$$\Sigma_1 = BB', \quad \Sigma_i = B\Lambda_i B', \quad i = 2, 3, \dots, M, \quad (2)$$

determines the matrix B . If B is constant across states and changes in the structural shocks' variance are distinct across variables, - up to sign - B is unique without imposing any further restrictions. Particularly in the case where $M > 2$, both conditions can be checked by formal statistical tests, see Lanne et al. (2010). A detailed discussion of tests for two and three state MS models can be found in Lanne et al. (2010) and Herwartz and Lütkepohl (2011). Apart from the statistical identification, exclusion restrictions on B are usually still necessary to obtain economically interpretable shocks. As demonstrated by Lütkepohl and Netšunajev (2013), one advantage of the identification through heteroscedasticity is that any restriction on B becomes over-identifying and, thus, testable.

Based on the conditional normality of the reduced form residuals, the likelihood function is set up and the model is estimated via maximum likelihood. Since the likelihood function is non-linear, numerical optimization methods are applied. The full algorithm and detailed discussion of related estimation problems can be found in Herwartz and Lütkepohl (2011).

¹⁰Note that we normalize variances of structural shocks in State 1 to be equal to one ($\Lambda_1 = I_K$)

Tests of over-identifying restrictions and confidence bands for impulse response functions are computed as suggested in Lütkepohl and Netšunajev (2013).

4 Estimation results

In this section we present the main results about inflation expectations spillovers between the US and EA. Results are based on the MS-SVAR model introduced in the previous section. We document the model selection procedure and how we achieve the identification of the country specific inflation expectations shocks. Finally, we show the main results derived from the state dependent variance decomposition.

4.1 Model specification and statistical identification

We order the variables in the vector Y_t in the following way:

$$Y_t = \left(\text{BEI}_t(\text{US } 5\text{Y}), \text{BEI}_t(\text{EA } 5\text{Y}), \text{BEI}_t(\text{US } 10\text{Y}), \text{BEI}_t(\text{EA } 10\text{Y}), \text{VIX}_t \right)'$$

Instead of focusing on the reduced form model (1), we are interested in structural shocks and their transmissions across the US and EA. While the BEI rates and their reduced form errors are correlated, the idea behind the structural shocks is that they occur in independence to each other and pin down the distinct sources that drive the variables in Y_t . For example, an inflation expectations shock in the US at a five year horizon does not predict the occurrence of any other structural shock within the system. However, the five year US shock may have a simultaneous or lagged effect on all inflation expectations measures and the risk measure. The contemporaneous effect of the structural shocks ε_t on the variables in Y_t are summarized in the matrix B .

$$B = \begin{pmatrix} b_{11} & b_{12} & b_{13} & b_{14} & b_{15} \\ b_{21} & b_{22} & b_{23} & b_{24} & b_{25} \\ b_{31} & b_{32} & b_{33} & b_{34} & b_{35} \\ b_{41} & b_{42} & b_{43} & b_{44} & b_{45} \\ b_{51} & b_{52} & b_{53} & b_{54} & b_{55} \end{pmatrix}, \quad \varepsilon_t = \begin{pmatrix} \varepsilon_t^{\pi^e(\text{US } 5\text{Y})} \\ \varepsilon_t^{\pi^e(\text{EA } 5\text{Y})} \\ \varepsilon_t^{\pi^e(\text{US } 10\text{Y})} \\ \varepsilon_t^{\pi^e(\text{EA } 10\text{Y})} \\ \varepsilon_t^{\text{Risk}} \end{pmatrix}$$

If the five year US inflation expectations shock $\varepsilon_t^{\pi^e(\text{US } 5\text{Y})}$ at time t is of size b_{11} , $\text{BEI}_t(\text{EA } 10\text{Y})$ responses by b_{41} .¹¹ The matrix of impact effects of shocks is of crucial importance since it determines impulse response functions and the variance decomposition, thus, feeds through

¹¹For the purpose of simpler interpretation we rescale the B matrix in such a way that the initial effect of shocks is unity.

Table 1: Markov switching- VAR model selection.

Model	$\log L_T$	AIC	SC
VAR(3) without MS	783.79	-1377.59	-1006.07
MS(2)-VAR(3)	1163.39	-2106.78	-1676.59
MS(3)-VAR(3)	1252.51	-2255.03	-1766.18*
MS(4)-VAR(3)	1290.15	-2300.30*	-1752.78

Notes: L_T is the value of the likelihood function, $AIC = -2 \log L_T + 2 \times \text{no of free parameters}$, $SC = -2 \log L_T + \log T \times \text{no of free parameters}$. Full sample Jan. 2004 - Jan. 2011 ($T = 372$ obs.).

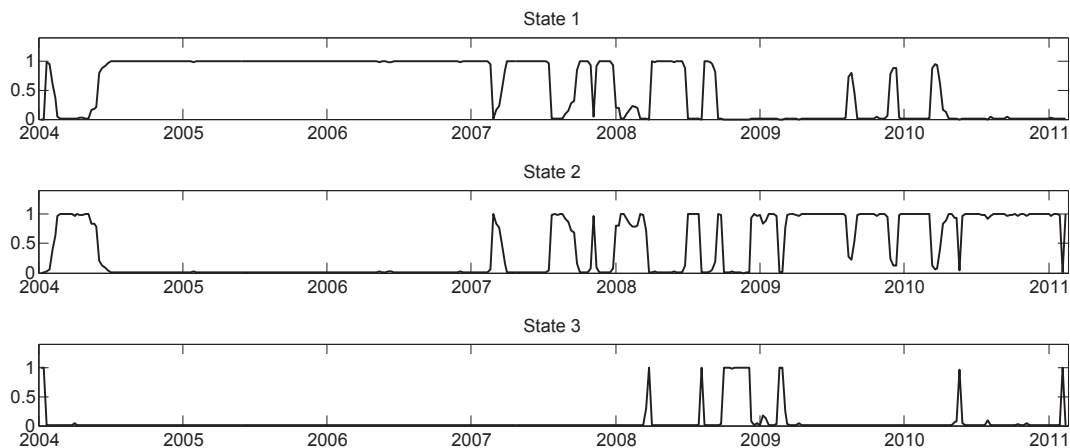
the whole system.

To specify an appropriate model for the identification of the structural matrices, we first choose the lag length of a standard VAR with constant parameters for the whole sample period from 2004 to 2011. We follow the suggestion of Schwarz criterion (SC) and continue with a VAR with three lags. We then implement the switching variance for different numbers of states M . Table 1 shows the log-likelihood and values of the Akaike Information Criterion (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 again preferred by the SC.

The estimated smoothed state probabilities of the MS(3)-VAR(3) 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 2007 is mainly associated with State 1, while State 2 and 3 dominate the second part of the sample. The period since 2007 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. Since crisis times include periods of relaxation, occasionally State 1 also materializes from 2007 onwards. State 3 captures the timing of key events like the failure of the home loan mortgage corporation Fannie Mae and Freddie Mac (July 2008), the investment banks Bear Stearns (March 2008) and Lehman Brothers (September 2008) as well as the downgrading of Greek government debt to junk bond status (April 2010). State 3 is also referred to as an "outbreak" state.

Given the indications in favor of a three state MS model, we estimate the MS-SVAR with three volatility regimes and three lags. Since we are interested in using heteroskedasticity for identification purposes, the data has to be informative on the relative variances of structural shocks. Put differently, the diagonal elements of the matrices Λ_2 and Λ_3 have to be distinct. In our model we find sufficient heterogeneity in variance, see the discussion

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



Notes: The states are characterized by three different volatility regimes. State 1 the lowest, State 3 the highest volatility. A state-invariant B matrix is imposed.

of LR-test results in the Appendix A. Furthermore, with a p -value of a LR-test of 0.44, we find that the decomposition shown in (2) is supported. Thus, the matrix B can be considered state-invariant in the three state model. Note that for the model with four MS states, preferred by the AIC (Table 1), these requirements for the statistical identification are not met. Although the MS(3)-SVAR(3) is statistically identified, restrictions on the matrix B are necessary to derive the US and EA specific inflation expectations shocks. In this context, any imposed restriction on B becomes over-identifying and testable, see Lütkepohl and Netšunajev (2013).

4.2 Identification of US and EA inflation expectations shocks

To derive the country and expectations horizon specific inflation expectations shocks, we translate and test restrictions advocated by the news and pass through regressions of e.g. Gürkaynak et al. (2010) and Jochmann et al. (2010). Our focus is on the following two zero restrictions on the elements of the B matrix.¹²

- **Restriction 1:** Medium horizon BEI rates are exogenous (e.g. Jochmann et al., 2010).
 \rightsquigarrow Inflation expectations shocks originating at the long (10Y) horizon do not affect the medium (5Y) horizon BEI rates contemporaneously.

$$H_0^1 : b_{13} = b_{14} = b_{23} = b_{24} = 0$$

¹²Restriction 1 and Restriction 2 are less restrictive than the modeling assumptions of the pass through and news regressions since we impose the restrictions on impact only such that the restricted shocks may have lagged effects on all variables in the SVAR system.

Table 2: Contemporaneous transmissions (Tests of zero restrictions on B).

No.	H_0	LR	df	p -value
R1	10Y IE shocks no effect on 5Y BEI	4.60	4	0.33
R2	R1 + No effect of 5Y IE shocks on foreign BEI	11.67	8	0.17
–	R1 + 5Y EA IE shocks no effect on US BEI	6.63	6	0.36
–	R1 + 5Y US IE shocks no effect on EA BEI	8.44	6	0.21

Note: Likelihood ratio tests are based on the MS(3)-SVAR(3) model. H_1 of all tests: state-invariant B . $LR = 2(\log L_T - \log L_T^r)$, where L_T^r denotes the maximum likelihood under H_0 and L_T denotes the maximum likelihood for the model under H_1 . $T = 372$ obs.

- Restriction 2: No cross country impact effects (e.g. Gürkaynak et al., 2010).
 \rightsquigarrow Inflation expectations shocks originating at the medium (5Y) horizon have no contemporaneous effects on foreign BEI rates.

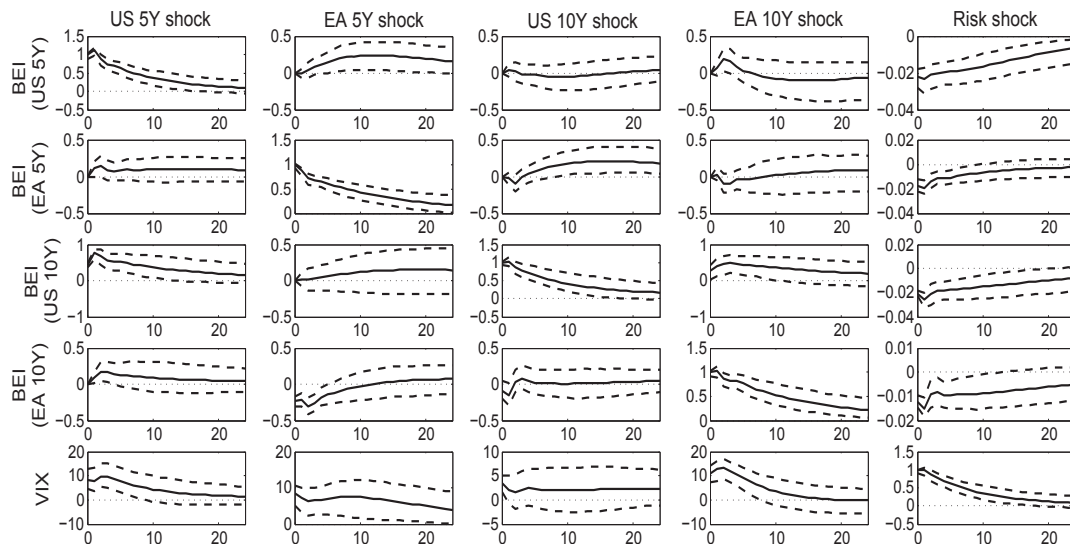
$$H_0^2 : b_{21} = b_{41} = b_{12} = b_{32} = 0$$

We start the testing of restrictions with imposing Restriction 1. Given that R1 is supported by the data we further test Restriction 1 and Restriction 2 jointly. Table 2 shows the LR-test statistics and corresponding p -values. This combination of restrictions is not rejected by the data either. However, we find that additional restrictions on B are not supported such that we proceed with the restricted (R2) MS(3)-SVAR(3). The corresponding restricted B matrix can be visualized as follows:

$$B = \begin{pmatrix} b_{11} & 0 & 0 & 0 & b_{15} \\ 0 & b_{22} & 0 & 0 & b_{25} \\ b_{31} & 0 & b_{33} & b_{34} & b_{35} \\ 0 & b_{42} & b_{43} & b_{44} & b_{45} \\ b_{51} & b_{52} & b_{53} & b_{54} & b_{55} \end{pmatrix}.$$

Note that the medium run shocks (first and second shock) are well separated by means of Restriction 2. On the contrary, long run shocks (third and fourth shock) are not separated by the Restriction 1. Restriction 1 separates medium run shocks from the long run shocks. The imposed restrictions on B carry over to the impulse response functions shown in Figure 3. Impulse responses discriminate the long run shocks originating in different countries and allow the labeling of the structural shocks. Time series of the five structural shocks and further discussion are provided in the Appendix B.

Figure 3: Impulse responses to inflation expectations and risk shocks.



Notes: MS(3)-SVAR(3) with R2 imposed (see Table 2). Impulse responses to a one unit structural shock with 95% bootstrap confidence bands based on 1000 replications. Full sample period (374 obs.).

US and EA five year inflation expectations shocks

As shown in the first two column of Figure 3, the country specific five year inflation expectations shocks have a significant contemporaneous and lagged impact on the level of domestic BEI rates. The US ten year BEI rate responds positively and strongly to a one unit US inflation expectation shock. On the contrary, the EA inflation expectation shock has a negative impact on the EA ten year BEI rate. This indicates that a ten year spot rate changes towards zero. We observe that the medium term shocks have a transitory effect for the ten year BEI rates for both countries. However, for the US the effect is much more persistent lasting for nearly 12 weeks, opposed by a 4 week long reaction in EA long term BEI rates, compare Strohsal and Winkelmann (2012). In the context of anchoring criteria as proposed by the pass through and news regressions (e.g. Lemke and Strohsal, 2013 and Beechey et al., 2011), we confirm a weak anchoring of inflation expectations in the US and a much stronger anchoring in the EA. Turning to cross country impact effects, recall that we find support of Restriction 2, meaning that US (EA) specific medium term shocks have no contemporaneous impact on the level of EA (US) BEI rates. Moreover, impulse responses of the shocks indicate that the reaction is mostly non-significant for the whole observed horizon.

Table 3: Relative variances of inflation expectations shocks in crisis states.

State	$\varepsilon^{\pi^e}(\text{US } 5\text{Y})$	$\varepsilon^{\pi^e}(\text{EA } 5\text{Y})$	$\varepsilon^{\pi^e}(\text{US } 10\text{Y})$	$\varepsilon^{\pi^e}(\text{EA } 10\text{Y})$	$\varepsilon^{\text{Risk}}$
2	4.96 (1.04)	6.92 (1.31)	2.23 (0.43)	4.42 (0.97)	7.72 (1.62)
3	48.6 (28.7)	43.69 (20.9)	17.4 (9.82)	28.2 (10.9)	78.0 (38.4)

Notes: Estimates of the restricted MS(3)-SVAR(3) model. Relative variances compared with the non-crisis state (variances in State 1 are normalized to one, see Equation (2)). Standard deviation in parentheses.

The response pattern of ten year BEI rates to the five year inflation expectations shocks display common characteristics as responses of BEI rates to US and EA specific economic news usually find in news regressions such as Galati et al. (2011) or Beechey et al. (2011). For that reason, we interpret the identified five year inflation expectations shocks as measures of US and EA specific macroeconomic news.

The state dependent relative variances of inflation expectations shocks are shown in Table 3. Variances of the five year shocks suggest that country specific uncertainty significantly increases during the crisis states. With around 6 and 45 times the variance in the non-crisis state (State 1), the increase in uncertainty has a similar order of magnitude in the US and EA. Compared with the ten year horizons, variances of five year inflation expectations shocks are much stronger affected throughout the crisis states.

US and EA ten year inflation expectations shocks

Impulse responses shown in the third and fourth column of Figure 3 illustrate adjustments to the ten year inflation expectations shocks. The imposed Restriction 1 does not allow to discriminate the origin of the long run shocks. However, looking at the third and fourth column of Figure 3 it becomes clear that the third shock has to be the US shock. The fourth shock can not be the US one as the impact effect of that shock on the US ten year BEI rate is insignificant. Using similar arguments, the fourth shock is labeled as the EA ten year inflation expectations shock. As well as Restriction 2, Restriction 1 does not only hold on impact but for most of the lags. Thus, we do not find any significant effect of long term shocks to the medium term domestic and foreign BEI rates. The impulse responses indicate that in our setup the level effect of the ten year shocks across the US and EA are mostly non-significant. Referring to Bomfin and Rudebusch (2000), we interpret the ten year inflation expectations shocks to reflect adjustments of the market perceived credibility of the FED and ECB, respectively (see also Appendix B). The level of five year BEI rates is therefore not significantly affected by credibility shocks but mainly driven by domestic

economic news.

The state dependent relative variances of the ten year inflation expectations shocks, shown in Table 3, suggest that the FED's and ECB's credibility are significantly weaker during the crisis states. The uncertainty about the ECB is around 4 and 28 times larger than during the non-crisis state. In comparison with the FED, the ECB's credibility is almost twice as much affected throughout the crisis periods. This finding indicates the pronounced impact of the European sovereign debt crisis and highlights the US and EA specific nature of the identified inflation expectations shocks.

Risk shock

The last column of Figure 3 shows effects of the shock derived from the VIX. The VIX displays characteristics of a global component. At least on impact it is affected by all inflation expectations shocks. Furthermore, the risk shock has a significant effect on all BEI rates in the system. Impulse responses display a negative relation between the BEI rates and the risk shock. The negative impulse responses provide evidence against a commodity price shock or a common factor of inflation expectations shocks from other countries than the US or EA. The negative impulse responses are in line with findings of Söderlind (2011) and Galati et al. (2011) and speak in favor of a risk premium shock which materializes through the liquidity of inflation indexed bonds. For that reason, we refer to the risk shock as a measure of (global) BEI risk premia.

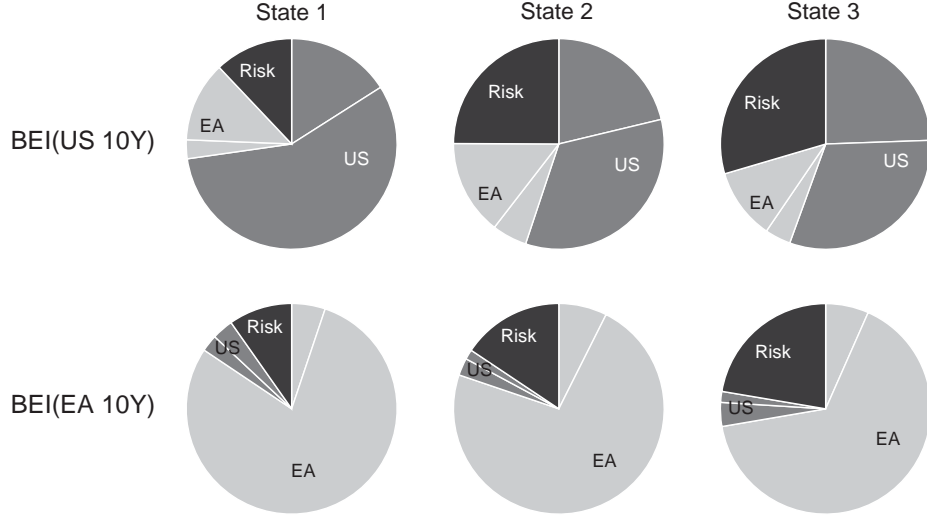
As shown in Table 3, the state dependent variance of the risk shock dominates the changes in variances of all other structural shocks. Along with impulse responses, this supports evidence of e.g. Hördahl and Tristani (2012) that risk premia play an important role for BEI rates.

So far, the impulse responses suggest that cross country effects play a minor role for the level of BEI rates. To quantify the role of spillovers in determining US and EA inflation expectations, we now derive the main results of the paper from the variance decomposition of the restricted MS(3)-SVAR(3) model. The state dependent nature of the inflation expectations and risk shocks allow the spillovers to vary across time.

4.3 Inflation expectations spillovers

Having identified the US and EA specific inflation expectations shocks, we now turn to assess their relative importance for US and EA inflation expectations. We first study the role of the common risk shock and then proceed with a risk adjusted analysis of inflation expectations spillovers. Following the pass through and news regression literature, the primary focus of the analysis is on explaining developments at the long (10Y) expectations

Figure 4: Variance decomposition of 10 year US and EA BEI rates.



Notes: Percentage of the variance explained by structural shocks in State 1 (non-crisis state), State 2 (crisis state) and State 3 (outbreak state). Equally shaded areas refer - clockwise - to the five year and ten year inflation expectation shock. Calculated from the restricted (R2) MS(3)-SVAR(3) forecast error variance decomposition at a two year forecast horizon.

horizons. The full set of results from the variance decomposition is shown in Figure C.1 in the Appendix C.

In order to get a first idea about the driving forces behind the 10 year US and EA BEI rates, Figure 4 illustrates the share of the total BEI variance explained by the structural shocks. We depict diagrams for each variance state separately such that the state dependent nature of the decompositions becomes apparent. In general, the figure shows that US and EA BEI rates are mainly driven by domestic structural shocks. However, also foreign inflation expectations shocks and the risk shock play a significant role. For the US BEI rate, the importance of the risk shock increases from 12 percent in the non-crisis states (State 1) to 30 percent in the outbreak state (State 3). Similarly, EA BEI rates are affected by 10 percent in the non-crisis state and up to 22 percent in the outbreak state. Since we identify the risk shock as measuring BEI risk premia, the variance decompositions indicate that BEI rates are less informative about inflation expectations during crisis periods.

Given the identification of the risk shock, the SVAR framework allows a risk adjusted analysis of spillovers. By setting risk shocks to zero, we consider variables driven by inflation expectations shocks only.¹³ The variance decomposition of the risk-free variables

¹³In the MA representation of the MS(3)-SVAR(3) the risk shock is set to zero while all other structural shocks remain unchanged.

Table 4: Inflation expectations spillovers.

State	US 10Y			EA 10Y		
	1	2	3	1	2	3
$\varepsilon^{\pi^e}(\text{US } 5\text{Y})$	18.2	28.3	34.7	2.9	3.1	3.7
$\varepsilon^{\pi^e}(\text{EA } 5\text{Y})$	3.3	7.1	5.6	5.7	8.8	8.5
$\varepsilon^{\pi^e}(\text{US } 10\text{Y})$	64.5	45.1	44.2	3.6	1.8	2.2
$\varepsilon^{\pi^e}(\text{EA } 10\text{Y})$	14.0	19.4	15.5	87.8	86.3	84.6
Total spillovers	17.3	26.5	21.1	6.5	4.9	5.9

Note: Percentage of the risk adjusted BEI variance explained by the structural inflation expectations shocks (ε^{π^e}) in States 1 (lowest volatility, non-crisis state) to state 3 (highest volatility, crisis state). Spillovers refer to aggregated spillovers of foreign five and ten year inflation expectations shocks. Calculated from forecast error variance decomposition at 100 weeks horizon.

are referred to as the variance decomposition of US and EA inflation expectations.

Table 4 shows the variance decomposition of 10 year US and EA inflation expectations for the three volatility states. For most of the shocks, spillovers across the US and EA are larger during the crisis states (State 2 and 3) than during the non-crisis state (State 1).

For US inflation expectations, EA shocks at the ten year horizon are relatively more important than EA shocks at the five year horizon. Specifically, while EA economic news account for up to 7.1 percent, credibility shocks triggered by the ECB explain up to 19.4 percent of US inflation expectations. Thus, most of the spillover effects are not attributed to economic news but to external effects of the ECB's monetary policy. Besides the relative importance of shocks at medium and long term expectations horizons, total spillovers of EA shocks are with 17.3 percent the smallest in the non-crisis state (State 1) and increase to 26.1 percent in the crisis state (State 2). Since spillovers are the highest in State 2 and mainly work through ECB credibility shocks, we conclude that the European Sovereign debt crisis has a pronounced impact on US inflation expectations.

In contrast to US inflation expectations, EA inflation expectations are far less affected by US specific shocks. US five and ten year inflation expectations shocks explain similar fractions of EA inflation expectations. Specifically, US economic news and credibility shocks induced by the FED determine up to 3.7 percent of the variance of EA inflation expectations. Most interesting, while the impact of US economic news increases in the crisis states, the role of the FED's credibility shocks decreases down to 1.8 percent (State 2). Although uncertainty about the US monetary policy increases significantly in the crisis states (see the discussion of Table 3) this uncertainty does not transmit to the

EA. Total spillovers of US specific inflation expectations shocks are rather stable across the non-crisis (State 1) and crisis states (State 2 and 3). The findings can, again, be explained by the important role of spillovers induced by the European sovereign debt crisis. It further suggests that the impact of the global financial crisis and global recession on EA inflation expectations can neither be ascribed to US economic news nor the FED's credibility shocks. Consequently, the global financial crisis and global recession materialize through the domestic inflation expectations shocks.

Overall, the results show that spillovers of inflation expectation shocks increase during crisis periods. In particular the impact of the European sovereign debt crisis on US inflation expectations highlights the increasing relevance of external effects of local policy measures.

5 Conclusion

This paper contributes to the empirical literature on inflation expectations by deriving US and EA specific inflation expectations shocks and studying their cross country spillovers. We propose to employ a structural vector autoregressive (SVAR) model identified via heteroscedasticity to derive the country specific inflation expectations shocks from US and EA break-even inflation (BEI) rates. A recently proposed Markov switching SVAR model enables a data driven identification without setting ad hoc restrictions on contemporaneous effects. We particularly benefit from the model in identifying different types of shocks that explain US and EA BEI rates. Closely related to modeling assumptions and empirical findings of the previous literature, we identify inflation expectations shocks that reflect US and EA specific macroeconomic news as well as shocks that capture the FED's and ECB's credibility in controlling inflation. Besides the inflation expectations shocks, we identify a risk shock which captures characteristics of a BEI risk premium. The identification of the risk shock allows a simple and transparent adjustment of BEI rates for market risk premia. Given the Markov switching structure of the model, we endogenously distinguish spillovers in non-crisis (low volatility) and crisis (high volatility) states.

The key result of the paper is that spillovers of US and EA inflation expectations increase during times of financial and macroeconomic stress. Throughout the considered time period from 2004 to 2011, US inflation expectations are stronger affected by EA shocks than EA expectations by US shocks. The strongest spillovers, explaining nearly 20 percent of US inflation expectations in the crisis state, originate in the EA and reflect transmissions of credibility shocks triggered by the ECB's monetary policy. The finding reflects a highly relevant external effect of the European sovereign debt crisis on US inflation expectations. The present paper puts forward new evidence to the discussion about increasing interna-

tional components in domestic inflation rates. Spillovers of inflation expectations shocks appear a promising candidate to explain a global inflation factor. In particular, our results highlight that the conduct and credibility of monetary policy play an important role for the cross country transmissions.

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Table A.1: Tests for Equality of variances across states.

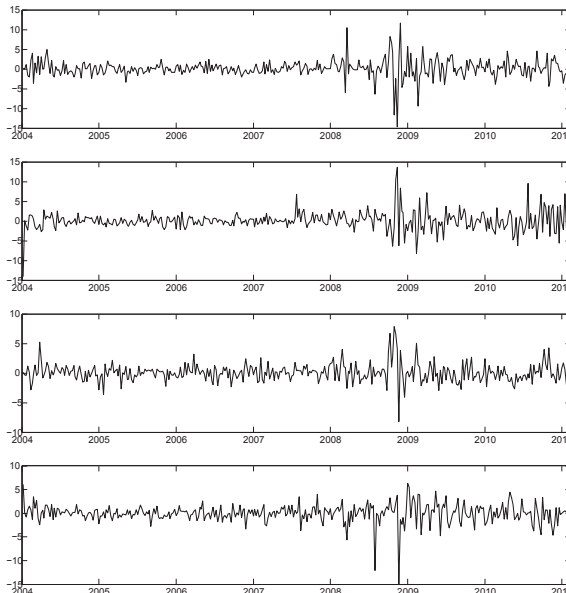
H_0	LR statistic	p -value
$\lambda_{21} = \lambda_{22}, \lambda_{31} = \lambda_{32}$	5.46	0.06
$\lambda_{21} = \lambda_{23}, \lambda_{31} = \lambda_{33}$	10.25	0.00
$\lambda_{21} = \lambda_{24}, \lambda_{31} = \lambda_{34}$	4.23	0.12
$\lambda_{21} = \lambda_{25}, \lambda_{31} = \lambda_{35}$	7.84	0.01
$\lambda_{22} = \lambda_{23}, \lambda_{32} = \lambda_{33}$	10.27	0.00
$\lambda_{22} = \lambda_{24}, \lambda_{32} = \lambda_{34}$	7.73	0.02
$\lambda_{22} = \lambda_{25}, \lambda_{32} = \lambda_{35}$	4.13	0.13
$\lambda_{23} = \lambda_{24}, \lambda_{33} = \lambda_{34}$	4.99	0.08
$\lambda_{23} = \lambda_{25}, \lambda_{33} = \lambda_{35}$	10.29	0.00
$\lambda_{24} = \lambda_{25}, \lambda_{34} = \lambda_{35}$	4.23	0.12

Notes: Tests for Equality of λ_{ij} s of MS(3)-VAR(3). Model with State-invariant and unrestricted B .

A Statistical identification

Since we are interested in using changing variances of structural shocks for identification purposes, the main question of interest is whether we have sufficient heterogeneity to get identification. For our five dimensional system we have to check if 10 pairs of relative variances are distinct across three states. To test the requirement formally we use likelihood ratio tests. Tests statistics of the relevant hypotheses are presented in Table A.1. For the current model, the null hypotheses of pairwise equality is rejected at a 15% significance level for all pairs. Looking closer at the results three null hypotheses may not be rejected at a 10% significance level. They are all related to the structural shock ordered last. This may be an indication that the statistical separation of structural shocks in the model with state invariant B may not be ideal, namely the fifth shock may be partly a mixture of other shocks. This may have implications for the restrictions testing shown in Table 2. Specifically the actual number of degrees of freedom of the limiting χ^2 distribution for the test may be lower by one unit than the number of identifying restrictions imposed on the B . This implies that the p values will be smaller than the ones shown in Table 2. However even with lower degrees of freedom the p values can not drop below conventional significance level, leading us to reject restriction R1 or R2. For that reason we think that we have enough heterogeneity in variances of structural shocks to support identification discussed in Section 4.2.

Figure B.1: Inflation expectations shocks.



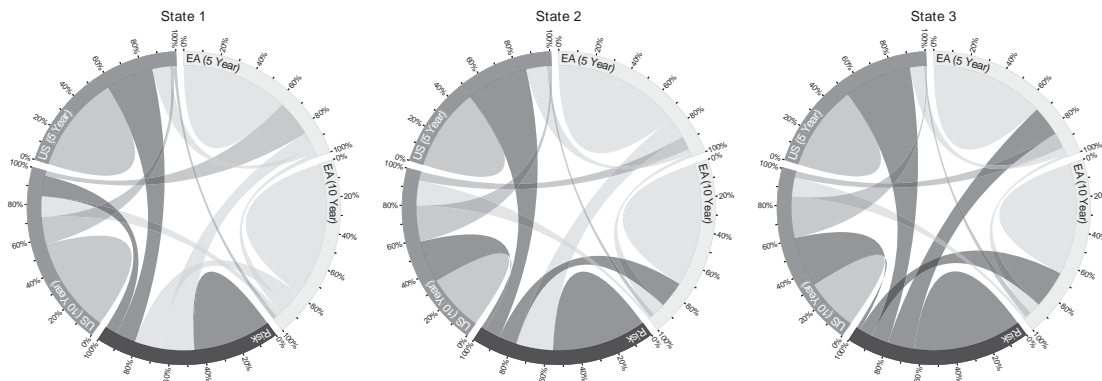
Notes: Structural shocks of the restricted (R2) MS(3)-SVAR(3). 1. US 5Y 2. EA 5Y 3. US 10Y 4. EA 10Y.

B Inflation expectations shocks

The time series of the US and EA specific inflation expectations shocks are shown in Figure B.1. The upper two graphs display the US and EA five year inflation expectations shocks. As discussed in the text (Section 2 and 4.2), the five year shocks are identified as US and EA specific economic news. News are usually defined as the deviation between a macroeconomic release (e.g. inflation or GDP) and markets' expectations prior to the release, i.e. a forecast error. Under rational expectations forecast errors should be uncorrelated, thus, share the characteristics of the structural shocks. The structural shock captures the full set of relevant news releases which are - most likely - not measurable through conventional news variables as employed by e.g. Gürkaynak et al. (2010) (the set of expectations about economic news releases is usually restricted). Since our data are weekly averages the shocks reflect averages of relevant economic news during the respective weeks.

The lower two graphs show the ten year inflation expectations shocks. These shocks are interpreted to reflect credibility shocks regarding markets perceptions about the FED's and ECB's ability to stabilize prices. The idea is that if a central bank is perfectly

Figure C.1: Variance decomposition.



Notes: The three chord diagrams display the forecast error variance decomposition (two year horizon) for the three volatility regimes (state 1 "non-crisis", state 2 and 3 "crisis"). Calculated from the restricted (R2) MS(3)-SVAR(3).

credible, the long term inflation expectations should be constant (anchored at an inflation target), see equation (4) of Bomfin and Rudebusch (2000). Since it is not only the BEI risk premium which causes the variation in the ten year BEI rates, we can conclude that inflation expectations are not perfectly anchored in the US and EA. In Section 4.2 and 4.3, we demonstrate that in particular domestic macroeconomic news appear to play a significant role. Moreover, the credibility is varying across time and driven by the uncorrelated credibility shocks – think of an MA process. An increasing variance of the ten year inflation expectations shocks (the credibility shocks) speaks in favor of a decreasing credibility.

C Forecast error variance decomposition

Figure C.1 depicts the full set of the variance decomposition referred to in Section 4.3. For each of the 5 variables (4 BEI rates+VIX), it presents the percentage of the variance explained by inflation expectations and risk shocks. The width of the connecting bands between two variables indicates the extend to which the structural shock affects the respective other variable. The colors of the bands refer to the color of the variable whose structural shock explains a larger portion of the variance. Non connected regions reflect the percentage explained by own structural shocks.

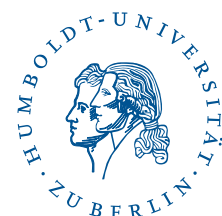
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