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Conference Paper

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Beiträge zur Jahrestagung des Vereins für Socialpolitik 2013: Wettbewerbspolitik und Regulierung in einer globalen Wirtschaftsordnung - Session: Inflation Expectations, No. E21-V1

Provided in Cooperation with:

Verein für Socialpolitik / German Economic Association

Suggested Citation: Lemke, Wolfgang; Strohsal, Till (2013) : What Can Break-Even Inflation Rates Tell Us about the Anchoring of Inflation Expectations in the Euro Area?, Beiträge zur Jahrestagung des Vereins für Socialpolitik 2013: Wettbewerbspolitik und Regulierung in einer globalen Wirtschaftsordnung - Session: Inflation Expectations, No. E21-V1, ZBW - Deutsche Zentralbibliothek für Wirtschaftswissenschaften, Leibniz-Informationszentrum Wirtschaft, Kiel und Hamburg

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What Can Break-Even Inflation Rates Tell Us about the Anchoring of Inflation Expectations in the Euro Area? *

Wolfgang Lemke[†] Till Strohsal[‡]

14 February 2013

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Abstract

We assess whether euro area inflation expectations, as measured by break-even inflation rates (BEIRs), have remained anchored during the financial crisis. Since autumn 2008, the volatility of BEIRs has increased considerably. We treat observed BEIRs as a sum of ‘genuine BEIRs’ and additional ‘noise’ components, the latter picking up influences related to market illiquidity or demand-supply imbalances, but not reflecting genuine inflation expectations and inflation risk premia. We estimate a bivariate VAR with short-term and long-term BEIRs, allowing for measurement noise in both. Anchoring of inflation expectations is analyzed by means of the pass-through of shocks from shorter to longer-term expectations. We find that, according to the pass-through results, inflation expectations remained well-anchored during the crisis period. Moreover, measurement noise accounts for up to 30% of the increase in volatility of BEIRs.

Keywords: Inflation expectations, anchoring, pass-through, break-even inflation rates, state space model

JEL classification: E31, E52, C32

*This paper should not be reported as representing the views of the European Central Bank (ECB). The views expressed are those of the authors and do not necessarily reflect those of the ECB. Of course, all remaining errors are our own. Financial support by the Deutsche Forschungsgemeinschaft (DFG) through CRC 649 “Economic Risk” is gratefully acknowledged.

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1 Introduction

Well-anchored long-term inflation expectations reflect the public's perception of a credible central bank commitment to price stability.¹ Inflation expectations can be inferred from surveys or, more indirectly, from financial market information.² The latter approach comes with the advantage that inflation expectations are available at a high frequency. Typically, inflation expectations are derived from the inflation swap market or from the difference between nominal and inflation-linked ('real') government bond yields of the same maturity, where this measure is referred to as 'break-even inflation rate' (BEIR). BEIRs reflect market participants' inflation expectations and so-called inflation risk premia, which are associated with the uncertainty about future inflation rates and their correlation with the economic environment. However, apart from this narrow interpretation, measured BEIRs may reflect additional factors related to certain supply-demand constellations of the markets for inflation-linked or nominal bonds, the influence of (relative) market illiquidity or other forms of market tensions.³

In the euro area, BEIRs have become considerably more volatile since the intensification of the financial crisis in autumn 2008. This invites two questions: first, to what extent can the higher volatility be interpreted as an increase in the volatility of inflation expectations and inflation risk premia in the narrow sense, and to what extent does it rather reflect other factors such as market tensions as mentioned before? Secondly, and more importantly, are inflation expectations during the period of elevated volatility still anchored?

In order to measure the degree of anchoring, one approach is to estimate the reaction of a measure of inflation expectations (market- or survey-based) to macroeconomic news or to movements in the current rate of actual inflation.⁴ An insignificant response of the long-term inflation expectations measure indicates anchored expectations. A closely related approach is the pass-through criterion that analyzes how changes in expectations over shorter horizons affect those over longer horizons.⁵ Given long-term inflation expectations are well-anchored, they should not be driven systematically by revisions of shorter-term expectations.

The pass-through approach in the literature has typically used a univariate linear model

¹See ECB (2012).

²See ECB (2011a) for a summary of indicators of inflation expectations in the euro area.

³See ECB (2012) and ECB (2009) for a discussion of various market factors on euro area BEIRs and their interpretation.

⁴See, e.g. Levin et al. (2004), Cuijsen and Demertzis (2007), Gürkaynak et al. (2010)

⁵See, e.g., Jochmann et al. (2010), Gefang et al. (2011)

specified in first differences. Actually, this framework relies on two assumptions. Firstly, the univariate approach considers shorter-term expectations as exogenous. Hence, such regressions are potentially subject to an endogeneity bias and conclusions might therefore be misleading. Secondly, the differencing implies a unit root in the level of expectations so that a change in shorter-term expectations has an infinitely long lasting effect. The idea of inflation expectations being anchored at some target, however, is somewhat at odds with the concept of nonstationarity.

This paper also uses a pass-through approach to measure the anchoring of inflation expectations, but deviates from the single-equation regression set-up in two respects. First, we endogenize the dynamics of both short-term and long-term inflation expectations by specifying their joint dynamics as a bivariate VAR. Second, we allow for ‘measurement noise’ for short-term and long-term BEIRs: the idea is to separate other short-lived influences from the interplay of underlying or ‘genuine’ inflation expectations and associated inflation risk premia.⁶

We estimate the VAR with measurement errors in a state space framework using Maximum Likelihood based on the Kalman filter. Our sample contains weekly data of euro area BEIRs from 2004 until the beginning of 2012 with a sample split at Lehman’s bankruptcy. The estimated model enables us to distinguish between shocks that actually reflect revisions of expectations and shocks that simply represent measurement error. We can filter out ‘genuine’ BEIRs from their observed noise-contaminated counterparts and analyze the pass-through from short-term to long-term filtered BEIRs before and during the crisis.⁷

We find that euro area inflation expectations remained well anchored also during the years of the global financial crisis. This finding is conservative in the sense that ignoring observation noise would lead to overestimating the degree of anchoring. Finally, up to 30% of the increase in the variance of observed BEIRs since the Lehman crash is attributed to measurement noise.

The rest of the paper proceeds as follows. Section 2 presents some descriptive statistics on euro area break-even inflation rates. Section 3 introduces the econometric methodology. Results are discussed in Section 4, and Section 5 concludes.

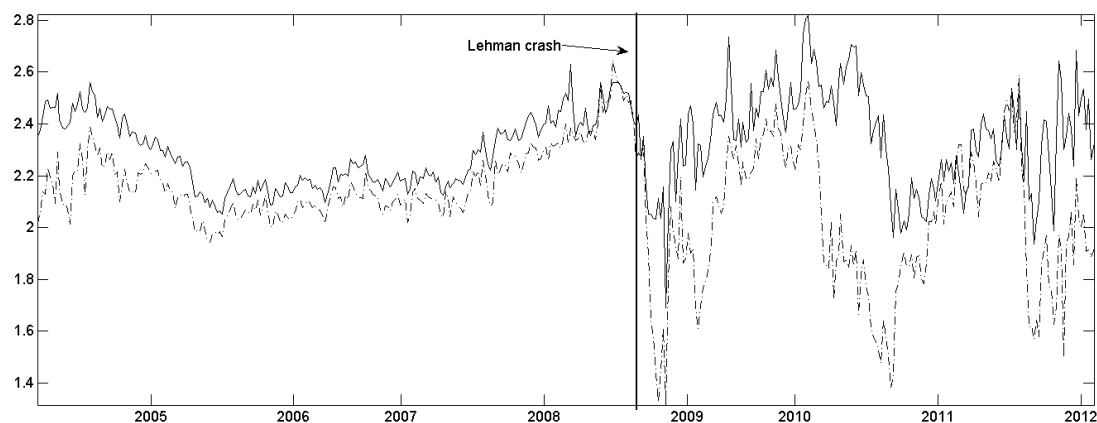
⁶However, we do not attempt to separate inflation expectations and inflation risk premia. In fact, for simplicity, we will often simply refer to the ‘genuine’ BEIRs as ‘genuine’ or ‘underlying’ inflation expectations, without acknowledging explicitly the presence of inflation risk premia.

⁷Note that we do not aim to explain the level of BEIRs, but only the relative dynamics of short- and long-term BEIRs. For an analysis of the driving forces of BEIRs, see, e.g. Ciccarelli and Garcia (2009).

2 Break-Even Inflation Rates in the Euro Area

Break-Even inflation rates refer to the difference between the yield of a nominal bond and the yield of a real bond. In the analysis below we use forward rates, where the m -year forward break-even rate beginning n years hence measures the rate that can be locked-in at t for the period $t+n$ until $t+n+m$. Our dataset consists of weekly observations in the sample period 02/02/2004 until 05/03/2012. The seasonally adjusted constant maturity series of BEIRs are based on ECB calculations according to Ejsing et al. (2007).⁸ We investigate the relationship between inflation expectations for two different horizons: the one-year forward rate four years ahead (BEI1F4) and the five-year forward rate five years ahead (BEI5F5). A four-year horizon ensures that BEIRs are generally based on liquid markets and its is often considered a proxy for shorter-term inflation expectations.⁹ The BEI5F5 represents a widely used market based measure, which is often referred to in official ECB publications to capture investors' long-term inflation expectations.¹⁰ Although BEIRs are not a pure measure of inflation expectations but rather include an inflation risk premium, we will use the terms 'BEIRs' and 'inflation expectations' synonymously in the remainder of the paper.

Figure 1: Observed Break-Even Inflation Rates



Notes: The solid line represents longer-term BEIRs (BEI5F5) whereas the dashed line refers to shorter-term BEIRs (BEI1F4). The sample period covers 02/02/2004 until 05/03/2012.

⁸See also ECB (2011b).

⁹See, e.g., Gürkaynak et al. (2010), Jochmann et al. (2010) and Gefang et al. (2011).

¹⁰See, e.g., ECB (2012)

Figure 1 shows the two series under investigation. The vertical line at 09/15/2008 refers to the bankruptcy of Lehman Brothers. What becomes immediately apparent is the pronounced co-movement and relatively low volatility of both series in the pre-crisis period. During the crisis, however, the BEIRs are less tightly connected with substantially increased volatility. Descriptive statistics of observed BEIRs are given in Table 1. We find means between 2.01 and 2.35 in the two sub-periods, respectively. The standard deviation of shorter-term rates more than doubled and increased significantly also for long-term expectations, when moving from the pre-crisis to the crisis sample. Furthermore, the correlation decreased from 0.85 to 0.51.

Table 1: Descriptive Statistics of Observed BEIRs

	pre Lehman 06-Feb-2004 to 12-Sep-2008			post Lehman 19-Sep-2008 to 09-Mar-2012		
	mean	std.	ρ	mean	std.	ρ
BEI1F4	2.175	0.134	0.852	2.010	0.283	0.513
BEI5F5	2.292	0.136		2.347	0.200	

Notes: The table shows empirical means and standard deviations of observed BEIRs. The empirical correlation is denoted by ρ .

The changing behavior of BEIRs over time leads to the natural question how this phenomenon should be interpreted with respect to the anchoring of inflation expectations. Moreover, why did the volatility of the expectations measure increase so strongly? In order to answer these questions, we will first review some of the ideas underlying the concept of the pass-through of inflation expectations and then introduce the empirical methodology.

3 Economic Motivation and Methodology

3.1 The Pass-Through Approach for Measuring the Anchoring of Inflation Expectations

Anchored inflation expectations should be characterized by market participants expecting inflation to eventually return to some implicit or explicit target. If that is the case, following Jochmann et al. (2010) and Faust and Henderson (2004), the process of actual

inflation can be modeled as

$$\pi_t = \theta\pi_{t-1} + (1 - \theta)\pi^* + u_t \quad , \quad (1)$$

with $|\theta| < 1$, π^* representing the inflation target and u_t as white noise. Thus, if market participants form conditional expectations at time t on inflation n periods ahead, we have

$$\pi_{t+n}^e = \pi^* + \theta^n(\pi_t - \pi^*) \quad (2)$$

for $n > 1$. In that framework, long-term inflation expectations are equal to the inflation target since $\theta^n \rightarrow 0$ as $n \rightarrow \infty$. If θ approaches 1, however, expectations are considered as unanchored because the target plays no role in the expectations formation process. As in the reduced form of the Stock and Watson (2007) model, inflation would follow a random walk and agents would expect it to remain at the same level as it is today.

According to equation (2), the idea of anchoring implies that longer-term expectations with horizon n_l are simply a multiple of shorter-term expectations with horizon n_s , namely

$$\pi_{t+n_l}^e = \theta^{n_l-n_s} \cdot \pi_{t+n_s}^e \quad . \quad (3)$$

From equation (3) it is often derived that the pass-through coefficient in the relation between the *first differences* of inflation expectations equals $\beta = \theta^{n_l-n_s}$, see Jochmann et al. (2010), Gefang et al. (2011).

$$\Delta\pi_{t+n_l}^e = \beta\Delta\pi_{t+n_s}^e + \varepsilon_t \quad (4)$$

For n_s fixed and $n_l \rightarrow \infty$ the regression coefficient goes to zero, which might be interpreted as a testable hypothesis of perfectly anchored inflation expectations.

On the other hand, what should be concluded, if β is found to be small, say 0.2? The empirical specification in (4) with ε_t white noise considers $\pi_{t+n_l}^e$ a random variable and imposes a unit root on the level of longer-term expectations. That is, a one unit change in shorter-term expectations would shift $\pi_{t+n_l}^e$ upwards by 0.2 units, and, ceteris paribus, longer-term expectations would stay at that higher level *forever*. We argue that this extreme persistence contradicts the basic idea of expectations being anchored at some target.

A further limitation of the univariate approach in (4) is given by the assumption that shorter-term expectations are exogenous, so that transmission is restricted to be only in one direction. This makes estimates of β potentially subject to endogeneity bias.

Finally, in practice, data from nominal and real yield curves might provide $n_l - n_s = 30$ years, if at all. Still, for a somewhat persistent inflation process with $\theta = 0.95$, we would have $\beta = 0.95^{30} = 0.21$. This numerical example shows that, empirically, it is not implausible to expect a pass-through coefficient estimate different from zero, even though expectations are anchored as $|\theta| < 0$.

Anchored long-term inflation expectations may therefore be influenced by short-term expectations for some periods. This effect should, however, vanish over time and $\pi_{t+n_l}^e$ should return to its long-run equilibrium: the inflation target. The faster the transmission from shorter to longer horizons fades out, the better the anchoring.

3.2 Empirical Approach: VAR with measurement errors

In contrast to the univariate specification in (4), this paper proposes a bivariate VAR that allows for measurement errors. This setting represents a more flexible and dynamic framework that allows inflation expectations to be mean-reverting and accounts for endogeneity of both expectations horizons.¹¹ Moreover, the model can be used to filter ‘genuine’ break-even inflation rates from their noise-contaminated data counterparts. The model to be used in our analysis is given by:

$$\tilde{\pi}_t^e = \pi_t^e + \varepsilon_t \quad \text{with } \varepsilon_t \sim N(0, H) \quad , \quad (5)$$

$$\pi_t^e = c + \sum_{i=1}^p \kappa_i \pi_{t-i}^e + \eta_t \quad \text{with } \eta_t \sim N(0, Q) \quad . \quad (6)$$

In the system (5) to (6), $\tilde{\pi}_t^e$ represents a (2×1) vector of observed short- and long-term BEIRs, $[\tilde{\pi}_{s,t}^e \ \tilde{\pi}_{l,t}^e]'$, and π_t^e refers to the (2×1) vector of corresponding underlying ‘genuine’ BEIRs, $[\pi_{s,t}^e \ \pi_{l,t}^e]'$. Inflation expectations are stationary if all roots of the reverse characteristic polynomial lie outside the unit circle, i.e., $(I - \kappa_1 z - \dots - \kappa_p z^p) \neq 0$ for $|z| < 1$.

Transitory measurement error is denoted by ε_t and represents what we call *observation* or *measurement noise*. By contrast, η_t represents the longer lasting shocks, which we call *revision noise* or *innovations*. Both types of shocks are serially uncorrelated with non-diagonal covariance matrices H and Q , respectively. However, it is assumed that ε_t and η_t are uncorrelated, or, $\mathbb{E}[\varepsilon_t \eta_t'] = 0$. As shown in Appendix A, the system (5) to (6) is identified. The model is cast into state space form and parameters are estimated by (quasi) maximum likelihood based on the Kalman filter.

¹¹Using VAR approaches (but without accounting for measurement errors) is quite common in the literature for investigating interest pass-through in the context of monetary policy transmission. See, e.g., Saborowski and Weber (2013) and the references cited therein.

The dynamic empirical framework allows us to investigate two aspects of inflation expectations: the pass-through over time and a decomposition of shock volatility that is associated with revisions of expectations (revision noise) and with measurement noise, respectively.

We define the pass-through over time as the generalized impulse response function, GIRF, (see Pesaran and Shin 1998) of the underlying longer-term inflation expectations to shocks in shorter-term inflation expectation:

$$GI_l(n, \delta_s, I_{t-1}) = \mathbb{E}[\pi_{l,t+n}^e | \eta_{st} = \delta_s, I_{t-1}] - \mathbb{E}[\pi_{l,t+n}^e | I_{t-1}] \quad . \quad (7)$$

Under normality, the GIRF becomes

$$GI_l(n, \delta_s, I_{t-1}) = \sigma_{s,Q}^{-1} e_l' A_n Q e_s, \quad n = 0, 1, 2, \dots, \quad (8)$$

with selection vectors $e_l = [0 \ 1]'$, $e_s = [1 \ 0]'$ and A_n as the coefficient matrix at lag n in the infinite moving average representation of (6).

In contrast to the orthogonalized impulse response¹², the GIRF does not rely on the ordering of the variables. That is, we do not a priori rule out instantaneous effects in one or the other direction. As a standard measure to characterize impulse responses, half-lives are calculated in the later empirical analysis. They indicate the number of periods that is needed for longer-term expectations to absorb 50% of the initial impact of a shock which is passed-through by shorter-term expectations.

The variance of observed break-even rates is determined by two components, the variance of observation noise H and the unconditional variance of the state process (the ‘genuine’ BEIRs), $R(Q)$ which depends on the variance of revision noise and the autoregressive parameters:

$$\text{Var}[\tilde{\pi}_t^e] = R(Q) + H \quad . \quad (9)$$

The elements of $R(Q)$ are given by the entries of the vector

$$\text{vec}(\text{Var}[\pi_t^e]) = (I - \sum_{i=1}^p (\kappa_i \otimes \kappa_i))^{-1} \text{vec}(Q) \quad . \quad (10)$$

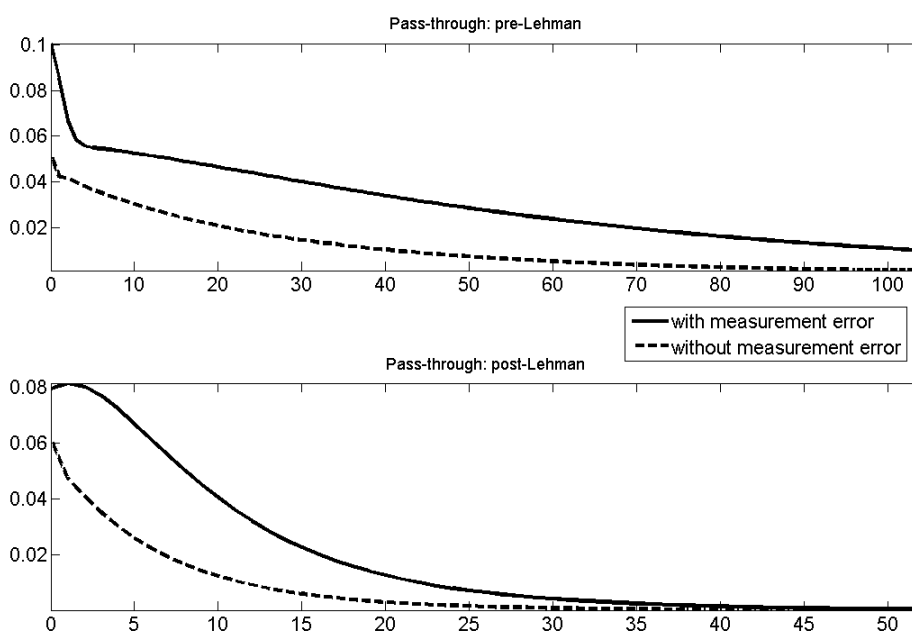
The decomposition (9) is later used to investigate to what extent the variability in observed BEIRs actually reflects revisions of inflation expectations as opposed to measurement noise.

¹²For a discussion, see Lütkepohl 2005)

4 Empirical Results

The following results are based on ML estimates of the state-space model given by (5) and (6). During model specification, the lag length p is successively increased until the residuals display no significant autocorrelation. For our data, the latter can statistically not be rejected for $p = 2$. Eigenvalues of the reverse characteristic polynomial turned out to be larger than one in modulus so that the vector process π_t^e is considered stationary.

Figure 2: Response of Long-Term BEIRs



Notes: The figure presents the response of long-term BEIRs to a 10 basis points shock in short-term BEIRs for two cases. Firstly, the solid line shows the response of 'genuine' BEIRs, i.e., BEIRs where measurement error is removed. Secondly, the dashed line refers to the estimates based on *observed* inflation expectations, i.e., BEIRs where measurement error is not removed. The response on the horizontal axes is measured in percentage points so that 0.1 corresponds to 10 basis points. The time on the horizontal axes is measured in weeks. The upper figure is based on the pre-Lehman sample. The lower figure is based on the post-Lehman sample with time only up to one year (52 weeks) since thereafter the response is essentially zero.

4.1 Results on the Pass-Through over Time

The upper and lower graphs in Figure 2 show the pass-through over time from shorter-term inflation expectations to longer-term expectations before and after the bankruptcy of Lehman. Apart from the pass-through of underlying BEIRs (i.e. the state variables), we also include the pass-through that results from *observed* BEIRs. More specifically, the latter is obtained by restricting the measurement variance H in equation (5) to zero. The size of the shock to shorter-term expectations is set equal to 10 basis points.

The pass-through of filtered inflation expectations in the pre-crisis period shows that at horizon zero the shock of 0.1 fully transmits to longer-term BEIRs. This result mirrors the high correlation of revision noise $\hat{\rho}_Q$, given in Table 2, which is accounted for in the GIRF formula (8). The pass-through exhibits a sharp decline in the following periods, however, so that at horizon 4 (one month) longer-term expectations have already absorbed 42% of the initial shock.

Table 2: Estimation Results: Revision Noise, Observation Noise and Correlation

pre Lehman	$\hat{\sigma}_H$	$\hat{\rho}_H$	$\hat{\sigma}_Q$	$\hat{\rho}_Q$	$\hat{\mathbb{E}}[\pi_t^e]$
BEI1F4	0.027 (0.003)	0.024 (0.550)	0.033 (0.005)	0.975 (0.007)	2.209
BEI5F5	0.010 (0.008)		0.034 (0.005)		2.314
post Lehman	$\hat{\sigma}_H$	$\hat{\rho}_H$	$\hat{\sigma}_Q$	$\hat{\rho}_Q$	$\hat{\mathbb{E}}[\pi_t^e]$
BEI1F4	0.077 (0.011)	0.674 (0.123)	0.065 (0.021)	0.707 (0.232)	1.995
BEI5F5	0.078 (0.013)		0.073 (0.020)		2.349

Notes: This table presents results from maximum likelihood estimation of the state-space model in (5) to (6). Standard errors based on the inverse hessian are given in parentheses. The estimated standard deviations of measurement error and revision noise are denoted by $\hat{\sigma}_H$ and $\hat{\sigma}_Q$. Analogously, $\hat{\rho}_H$ and $\hat{\rho}_Q$ refer to the correlations. Model implied unconditional means, $\hat{\mathbb{E}}[\pi_t^e]$, are given in the last column.

When measurement noise is neglected, a different picture emerges: the impulse response suggests a weaker transmission of shocks to short-term BEIRs to long-term BEIRs. As can be seen from Table 2, there is no significant observation noise and accordingly no significant correlation in \hat{H} . A specification which ignores that the correlation of shorter- and longer-term inflation expectations is driven by both H and Q would therefore tend to underestimate the correlation of revision noise. In that sense, our approach represents a conservative procedure to assess the anchoring of inflation expectations. In

Table 3: Half-Lives: Filtered vs. Observed BEIRs

	State-Space	VAR
pre Lehman	15	15
post Lehman	12	5

Notes: This table shows half-lives to characterize the pass-through over time. The second column includes half-lives of the response from long-term BEIRs to a shocks in short-term BEIRs based on the state-space approach. The third column presents those, observed from a VAR, i.e. when measurement error is ingnored.

the pre-crisis sample, the effect of underestimating the correlation of η_t turns out to be pronounced. Setting $H = 0$, we estimate a shock correlation of 0.63, which leads to a lower instantaneous transmission of about 52% of the initial shock in the pre-crisis period.

In the crisis period we find no evidence for a deanchoring of filtered inflation expectations. On the contrary, correlation of revision noise declines so that the instantaneous pass-through is approximately 8 basis points only. The half-life of the initial shock equals 12 weeks (see Table 3). Moreover, the shape of the GIRF is different in the crisis period. We observe a smooth decreasing of the pass-through, being practically zero after one year. Considering half-lives, our estimates suggest that longer-term inflation expectations are even somewhat more firmly anchored during the crisis period.

Ignoring measurement error has the same effect as in the pre-crisis period even though the impact on the instantaneous pass-through is smaller. The underestimation of revision noise correlation is less pronounced since we find $\hat{\rho}_H$ and $\hat{\rho}_Q$ to be close together and to be significant.

4.2 Decomposing the Variance of BEIRs

In order to shed some light on the increased volatility of break-even rates, we split the total variance of the observed series into two parts: $R(Q)$ (variance related to revision noise) and H (variance of observation noise), see equation (9). In the pre-crisis period, we find that only 4% of the total variation of shorter-term expectations and 0.5% of longer-term expectations is related to measurement error.¹³ Therefore, in the pre-crisis

¹³Statistically, observation noise in longer-term inflation expectations is equal to zero and the 0.5% simply refer to the insignificant point estimate.

Table 4: Volatility Decomposition: Revision Noise and Observation Noise

	$\text{Var}[\tilde{\pi}_t^e]$	$R(Q)$	H	$\Delta H/\Delta \text{Var}[\tilde{\pi}_t^e]$
pre Lehman				
BEI1F4	$1.954 \cdot 10^{-2}$	$1.879 \cdot 10^{-2}$	$0.075 \cdot 10^{-2}$	
BEI5F5	$2.033 \cdot 10^{-2}$	$2.022 \cdot 10^{-2}$	$0.011 \cdot 10^{-2}$	
post Lehman				
BEI1F4	$7.693 \cdot 10^{-2}$	$7.086 \cdot 10^{-2}$	$0.607 \cdot 10^{-2}$	9.28%
BEI5F5	$3.956 \cdot 10^{-2}$	$3.343 \cdot 10^{-2}$	$0.614 \cdot 10^{-2}$	31.35%

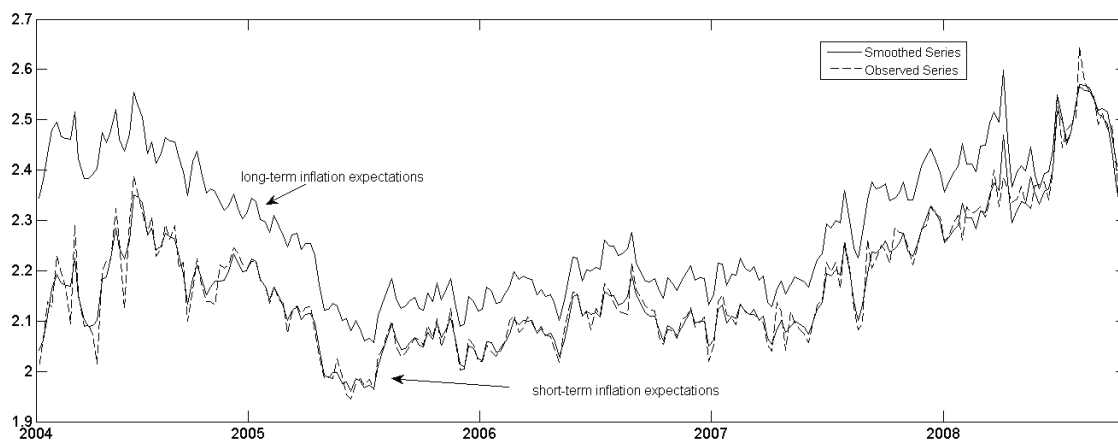
Notes: $\text{Var}[\tilde{\pi}_t^e] = R(Q) + H$ denotes the total volatility of inflation expectations decomposed into the part that is related to revision noise and the part that reflects observation noise. The last column shows the fraction of the increase in volatility, which is driven by observation noise.

period, the movement of BEIRs should mostly reflect revisions of inflation expectations. During the crisis period, however, the fraction of volatility related to observation noise increases. We find \hat{H} to account for approximately 8% in the variation of shorter-term expectations. In longer-term expectations, our estimates imply that actually 15% of the total variation is simply due to measurement noise. These numbers become even more meaningful in terms of changes. Approximately 10% and 30% of the increase in total volatility are related to observation noise in shorter- and longer-term expectations, respectively.

In order to get a graphical impression of the underlying BEIRs, Figures 3 and 4 show the smoothed estimates $\mathbb{E}[\pi_t^e | I_T]$, in which $I_T = \{y_0, y_1, \dots, y_t, \dots, y_{T-1}, y_T\}$. The first Figure covers the pre-crisis period. In line with the estimates from Table 2, there is no visible difference between observed and smoothed longer-term inflation expectations. Smoothed shorter-term expectations, however, exhibit a lower volatility than the observed series, as, for instance, clearly noticeable in the first half of 2004.

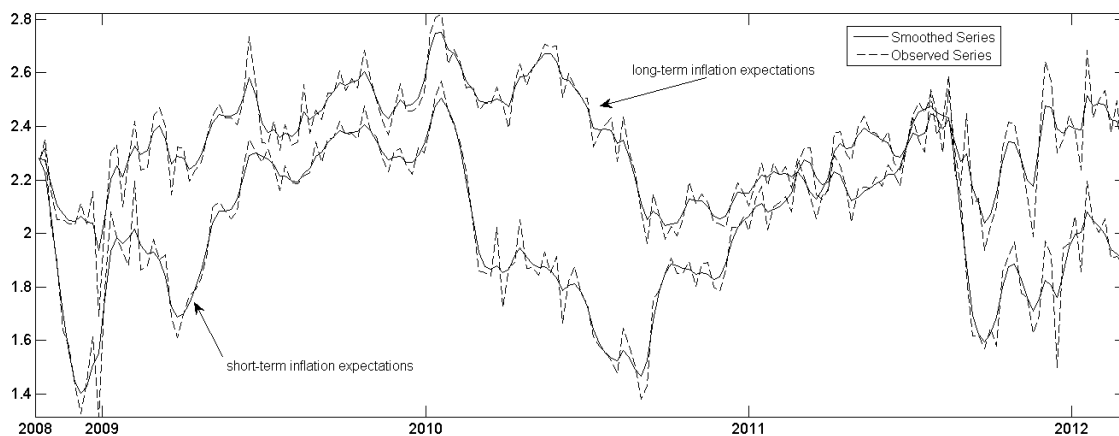
In the crisis period, the measurement error is evident in both series. A substantial fraction of the transitory fluctuations is removed throughout the whole post-Lehman period. Towards the end of 2011, for example, spikes of the smoothed series are approximately 20 basis points smaller than those of the observed series.

Figure 3: Smoothed BEIRs: Pre-Lehman Period



Notes: Solid lines represent observed inflation expectations. Dashed lines indicate Kalman-smoothed inflation expectations $\mathbb{E}[\pi_t^e | I_T]$.

Figure 4: Smoothed BEIRs: Post-Lehman Period



Notes: Solid lines represent observed inflation expectations. Dashed lines indicate filtered inflation expectations based on the Kalman smoother $\mathbb{E}[\pi_t^e | I_T]$.

5 Conclusion

We analyze the anchoring of long-term inflation expectations in the euro area by investigating the pass-through from shocks to shorter-term BEIRs to their longer-term counterparts. As the variance of BEIRs has strongly increased since the intensification of the global financial crisis in autumn 2008, we also investigate how much of this volatility increase can be attributed to actual revisions of expectations and how much does simply

reflect measurement noise. For the econometric analysis, we employ a VAR allowing for measurement errors. Our results imply no signs of a de-anchoring of euro area inflation expectations during the crisis. Moreover, up to 30% of the increase of volatility can be assigned to measurement error.

A Appendix: Model Identification

The state-space model in equations (5) to (6) for the preferred lag length has the following representation:

$$(I - \kappa_1 L - \kappa_2 L^2) \tilde{\pi}_t^e = \eta_t + (I - \kappa_1 L - \kappa_2 L^2) \varepsilon_t . \quad (11)$$

According to Granger's lemma (Granger and Newbold 1977), this is equivalent to an ARMA(1,2) model. The autoregressive coefficient matrices κ_1 and κ_2 in the reduced form (11) are the same as in the structural model (5) to (6), so they can be ignored and we may define the 'observables' $Z_t = (I - \kappa_1 L - \kappa_2 L^2) \tilde{\pi}_t^e$ as follows:

$$z_{1,t} = \eta_{1t} - \kappa_{1,11} \varepsilon_{1,t-1} - \kappa_{1,12} \varepsilon_{2,t-1} - \kappa_{2,11} \varepsilon_{1,t-2} - \kappa_{2,12} \varepsilon_{2,t-2} + \varepsilon_{1,t} \quad (12)$$

$$z_{2,t} = \eta_{2t} - \kappa_{1,21} \varepsilon_{1,t-1} - \kappa_{1,22} \varepsilon_{2,t-1} - \kappa_{2,21} \varepsilon_{1,t-2} - \kappa_{2,22} \varepsilon_{2,t-2} + \varepsilon_{2,t} . \quad (13)$$

The coefficient matrix κ_1 , for instance, is defined as $\begin{pmatrix} \kappa_{1,11} & \kappa_{1,12} \\ \kappa_{1,21} & \kappa_{1,22} \end{pmatrix}$. Equations (12) and (13) give 11 moments, namely 2 variances and a covariance, $\mathbb{E}[z_{1,t}^2]$, $\mathbb{E}[z_{1,t} z_{2,t}]$ and $\mathbb{E}[z_{2,t}^2]$, and 8 auto- and cross-correlations, $\mathbb{E}[z_{1,t} z_{1,t-2}]$, $\mathbb{E}[z_{2,t} z_{2,t-2}]$, $\mathbb{E}[z_{1,t} z_{2,t-2}]$, $\mathbb{E}[z_{2,t} z_{1,t-2}]$, $\mathbb{E}[z_{1,t} z_{1,t-1}]$, $\mathbb{E}[z_{2,t} z_{2,t-1}]$, $\mathbb{E}[z_{1,t} z_{2,t-1}]$ and $\mathbb{E}[z_{2,t} z_{1,t-1}]$. All these moments depend on the 6 unknown parameters of the structural model, Q_{11} , Q_{12} , Q_{22} , H_{11} , H_{12} and H_{22} . The system of equations

$$A(Q_{11} \ Q_{12} \ Q_{22} \ H_{11} \ H_{12} \ H_{22})' = \mathbb{E}[\cdot] , \quad (14)$$

with A of dimension (11×6) and $\mathbb{E}[\cdot]$ as the vector of moments¹⁴ with dimension (11×1) , summarizes the conditions. A coefficient matrix A , given by

$$A = [A_1 \ | \ A_2] , \quad (15)$$

¹⁴Moments are ordered as in the above text.

where $A'_1 =$

$$\begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{pmatrix}$$

and $A_2 =$

$$\begin{pmatrix} 1 + \kappa_{1,11}^2 + \kappa_{2,11}^2 & 2(\kappa_{1,11}\kappa_{1,12} + \kappa_{2,11}\kappa_{2,12}) & \kappa_{1,12}^2 + \kappa_{2,12}^2 \\ \kappa_{1,11}\kappa_{1,21} + \kappa_{2,11}\kappa_{2,21} & \kappa_{1,11}\kappa_{1,22} + \kappa_{1,12}\kappa_{1,21} + \kappa_{2,11}\kappa_{2,22} + \kappa_{2,12}\kappa_{2,21} & \kappa_{1,12}\kappa_{1,22} + \kappa_{2,12}\kappa_{2,22} \\ \kappa_{1,21}^2 + \kappa_{2,21}^2 & 2(\kappa_{1,21}\kappa_{1,22} + \kappa_{2,21}\kappa_{2,22}) & 1 + \kappa_{1,22}^2 + \kappa_{2,22}^2 \\ -\kappa_{2,11} & -\kappa_{2,12} & 0 \\ 0 & -\kappa_{2,21} & -\kappa_{2,22} \\ 0 & -\kappa_{2,11} & -\kappa_{2,12} \\ -\kappa_{2,21} & -\kappa_{2,22} & 0 \\ -\kappa_{1,11}(1 - \kappa_{2,11}) & -(\kappa_{1,12} - \kappa_{1,11}\kappa_{2,12} - \kappa_{2,11}\kappa_{1,12}) & \kappa_{1,12}\kappa_{2,12} \\ \kappa_{1,21}\kappa_{2,21} & -(\kappa_{1,21} - \kappa_{1,22}\kappa_{2,21} - \kappa_{1,21}\kappa_{2,22}) & -\kappa_{1,22}(1 - \kappa_{2,22}) \\ \kappa_{1,21}\kappa_{2,11} & -(\kappa_{1,11} - \kappa_{1,22}\kappa_{2,11} - \kappa_{1,21}\kappa_{2,12}) & -(\kappa_{1,12} - \kappa_{1,22}\kappa_{2,12}) \\ -(\kappa_{1,21} - \kappa_{1,11}\kappa_{2,21}) & -(\kappa_{1,22} - \kappa_{1,11}\kappa_{2,22} - \kappa_{1,12}\kappa_{2,21}) & \kappa_{1,12}\kappa_{2,22} \end{pmatrix}$$

of full rank would ensure the sufficient condition for identification. One possibility to show that A has actually rank 6, is to consider the (6×6) leading principal minor. We can expand this determinant by the first column. Repeating this procedure yields

$$-\kappa_{2,11} \cdot \begin{vmatrix} \kappa_{1,11} & \kappa_{1,12} \\ \kappa_{1,21} & \kappa_{1,22} \end{vmatrix} = -\kappa_{2,11}(\kappa_{2,12}\kappa_{2,21} - \kappa_{2,11}\kappa_{2,22}) .$$

In general, this expression is different from zero and so A has full rank.

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