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TAIL COMOVEMENT IN OPTION-IMPLIED INFLATION EXPECTATIONS AS AN INDICATOR OF ANCHORING

by Sara Cecchetti, Filippo Natoli and Laura Sigalotti*

Abstract

We analyse the degree of anchoring of inflation expectations in the euro area. Using a new estimation technique, we look at the tail co-movement between the moments of short- and long-term distributions of inflation expectations, where those distributions are estimated from daily quotes of inflation derivatives. We find that, since mid-2014, negative tail events impacting short-term inflation expectations have been increasingly channelled to long-term views, igniting both downward revisions in expectations and upward changes in uncertainty; instead, positive short-term tail events have left long-term moments mostly unaffected. This asymmetric behaviour may signal a disanchoring from below of long-term inflation expectations.

JEL Classification: C14, C58, E31, E44, G13.
Keywords: disanchoring, inflation swaps, inflation options, option-implied density, tail comovement.

Contents

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1. Introduction [1]

Headline inflation in the euro area has been falling since 2012 and became negative at the end of 2014. The 5y5y forward inflation swap, a measure of medium- to long-term inflation expectations, has fallen well below 2 per cent since September 2014; on September 4, the phrase *Inflation expectations for the euro area over the medium to long term continue to be firmly anchored* was abandoned for the first time in the ECB's monetary policy statements. The sovereign debt purchase programme (QE) announced in January 2015 aims not only at tackling the fall in actual inflation, but also at countering the undershooting of medium-term beliefs. With nominal interest rates at the zero lower bound, decreasing long-term inflation expectations tend to raise real rates, thus tightening financial conditions. Moreover, market agents pricing a persistent departure of inflation from the 2% target could reveal a loss of credibility of the monetary authority; consumers might be induced to postpone consumption and investments, leading to a deflationary spiral that might become entrenched.

In this paper we analyse whether there has been a downward disanchoring of long-term inflation expectations in the euro area. After reviewing available contributions, we discuss how a disanchoring could be identified and propose a new technique for estimating it by looking at daily movements in inflation swaps and inflation options. Only a handful of papers address this topic for the euro area in the period after the global financial crisis. An earlier study by Autrup and Grothe (2014) found that inflation expectations were firmly anchored, with the sample ending in 2012. As far as we know, the only paper that studies market-based inflation expectations and includes the last three years of data has, on that specific question, mixed results (Scharnagl and Stapf (2015)). We focus on signals that might point tol a possible transition from anchored to unanchored expectations during the most recent period.

Our approach is based on two key assumptions. First of all, when expectations are firmly anchored to the central bank's target, short- and long-term views should not comove. This means that shocks to short-term inflation expectations are not transmitted to long-term beliefs; on the contrary, a response of long-term expectations to actual inflation readings and to macroeconomic surprises, implying a positive and strong correlation with short-term expectations, can be interpreted as a signal of disanchoring. We can reasonably assume that, in an early phase of disanchoring, only sizable shocks producing unusual upswings

[1] We would like to thank Pietro Catte, Paolo Del Giovane, Andrea Finicelli, Giuseppe Grande, Marco Taboga, Jean-Paul Renne, Jelena Stapf and the participants to the March 2015 Bank of Italy workshop on *Low inflation and its implications for monetary policy* for their helpful comments and suggestions. The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Italy. All the remaining errors are ours. E-mail: sara.cecchetti@bancaditalia.it, filippo.natoli@bancaditalia.it, laura.sigalotti@bancaditalia.it.

or downswings in short-term expectations induce changes in long-term views. This could imply a *tail* reaction in long-term expectations as a result of a strong variation of short-term expectations. A standard correlation indicator, such as the Pearson's correlation coefficient, however, does not distinguish between positive and negative or core and tail comovements: for example, an increase in average correlations between short- and long-run mean expectations could be driven by comovements in the upper tail with constant or decreasing lower tail correlations. This case is not coherent with a possible scenario of disanchoring from below, provided that lower-than-expected macro data should first imply correlations in the left tails; therefore, using average correlations as the unique tool to make insights on the level of anchoring could be, at this stage, quite misleading.[2]

Secondly, many commentators pointed out that the anchoring of inflation expectations does not only require the containment of the level of expectations but, in general, stable market beliefs about future inflation (e.g., Gurkaynak et al. (2010), among others). This is highly relevant because unanchored expectations from one side (above the target) or from the other one (below it) imply an asymmetric attitude of market investors towards future inflation outcomes. Coherently with this view, inflation targeting should help anchor market perceptions of the entire distribution of future long-run outcomes. We therefore claim that disanchoring occurs not only when average medium- to long- term expectations are significantly far away from the central bank's target, but also when the *moments* of the distribution of long-term expectations are responsive to shocks impacting short-term beliefs.

Putting these two elements together, we investigate possible signs of disanchoring by estimating the comovement in the tails of short and long-term moments of the distributions of inflation expectations at different maturities. An increased comovement in the tails could signal increased sensitivity of long-term beliefs to economic news, per se a warning of possible disanchoring; moreover, an asymmetric response to extreme shocks (i.e., increasing comovement in one tail and constant or decreasing comovement in the other one) might suggest that the balance of risk is tilted to one side, or that long-term uncertainty only reacts to positive or negative news.

The whole investigation is carried out in two steps. In the first step, we derive the option-implied probability distributions of future inflation 1, 2, 3, 5, 7 and 10 years ahead from quoted inflation caps and floors, on a daily basis in the period between 2009 and 2015. To achieve this, we employ the newly developed semi-nonparametric technique presented in Taboga (2015). In this way, we are able to recover the term structures of the mean, standard deviation and skewness of inflation expectations, in addition to the time series of deflation

[2] These limitations of the Pearson's correlation coefficient should be taken into account also in investigating a possible disanchoring from above.

and high inflation probabilities and quantiles of the distributions. Secondly, we focus on the series of daily changes in average expectations and option-implied standard deviations: by selecting short and long maturities for each moment, we measure the bivariate comovement in the tails of the empirical distributions, in order to gauge the resilience of long-term beliefs and uncertainty to sizable shocks. We compute both the tail correlation measures based on copulas and the TailCor estimators of Ricci and Veredas (2013). Concerning mean inflation expectations, note that spot and forward inflation swap rates are used in place of option-implied values in order to avoid a spurious dependence of long-term expectations from short-term ones.

We find that the correlation in the tails of mean expectations and uncertainties has substantially increased since mid-2014. Interestingly, the comovement in short- and long-term mean expectations has increased only in the lower tail, meaning that negative shocks have been channeled to long-term average beliefs while positive ones are not. This result is robust to the specification of short-term expectations as the 1y spot and 1y1y forward inflation swap rate (in addition to the 5-year spot), while medium-to-long term expectations are always proxied by the 5y5y forward. Concerning option-implied estimates, we compare the higher moments of the 1, 2 and 3-year distributions with the 10-year ones. Short and long-term standard deviations comove only in the upper tail, signalling that uncertainty is more responsive to upward than downward short-term variations. In general, the effect of extreme shocks on the inflation expectations of market operators has recently become asymmetric: first of all, agents tend to react much more to downward than upward revisions of short-term inflation expectations; secondly, increases of uncertainty around short-term expectations tend to be associated with increases in long-term uncertainty. Both results are robust to different rolling windows and TailCor parameterizations. We can conclude that, in light of the whole investigation carried out, some signs of disanchoring have emerged during the last months.

Our first result is in line with Ehrmann (2014), who studies the stability of long-term beliefs in a panel of countries before and after inflation-targeting: under persistently low inflation, he finds that a sign of disanchoring from below with respect to a target is that inflation expectations get revised down in response to lower-than-expected inflation but do not respond to higher-than-expected outturns. The most common method used to assess the degree of anchoring in one economy involves testing the sensitivity of inflation expectations to surprises in macro news (the *news-regression* approach of Gurkaynak et al. (2005)). Using this technique, Ehrmann et al. (2011) analyse inflation expectations in some euro area countries from 1993 to 2008 using bond data, finding that the level of anchoring increased after the adoption of the single currency in 1999. Making a cross-country comparison, Beechey et al.

(2011) conclude that, over the same time span, expectations were more firmly anchored in the euro area than in the US. The method adopted here shares the same principle underlying the news-regression approach: being unusual and unexpected, large surprises in macro news can reasonably be labelled as tail events. While news regressions heavily depend on how the surprise component of announcements is estimated, our approach is totally market-based, so it is free from identification issues.[3] The link between tail comovements and anchoring is first investigated in Antunes (2015), where coefficients of upper tail dependence between daily revisions of short and long term inflation swaps are constructed using different types of bivariate copulas. That paper is especially concerned on the comparison of different parameterizations, and does not focus on the issue of possible asymmetric dependence in the tails; moreover, no investigations of the higher moments of the distribution are conducted.

We contribute to the literature in several ways, in addition to the newly designed technique to evaluating anchoring. First of all, this is the only paper that is able to draw conclusions on the issue of whether a disanchoring in the euro area has occurred in the last 3 years, when actual inflation drifted downward and deflation started to become a serious concern. Secondly, apart from Scharnagl and Stapf (2015), no other paper provides estimates of option-implied distributions of future inflation for the euro area; finally, no other applications of the copula and TailCor methods exist with option-implied data as far as we know.

This paper is organized as follows. In Section 2, we describe the dataset and the derivation of option-implied distribution of future inflation; we compute the time series of inflation and deflation probabilities, quantiles, means, standard deviations and skewnesses for the available maturities. Section 3 briefly explains the techniques employed to estimate the tail comovement and comments on the main results and Section 4 concludes.

2. Option-implied distributions of future inflation

The market for inflation-linked derivatives has witnessed a considerable development in the past few years; in particular, the most popular inflation derivatives include inflation swaps and inflation options (caps and floors).

An inflation cap is a derivative contract in which the holder has the right to receive compensation payments at the end of each period in which the inflation rate exceeds an agreed-upon strike rate. The contract involves no obligations when the realized inflation is below the strike. In exchange for the contingent future payment, the holder pays a

[3] The identification of the surprise effect in some specific news has recently been questioned: for instance, focusing on the 53 QE announcements made by the FOMC in the United States considered in the literature, Thornton (2014) finds that none of them meets the strict requirements for identification, and just 11 meet some of the requirements.

price (option premium) upfront; the quoted price reflects market beliefs on the probability distribution of future inflation. Analogously, a floor is a derivative contract which gives the holder the right to receive payments at the end of each period in which the inflation rate falls below the predetermined strike. A zero-coupon inflation option consists of a single compensation payment at maturity, while a year-on-year inflation option includes intermediate payments depending on the level of the inflation rate in each year of the reference period.

Bloomberg provides quotes for both zero-coupon and year-on-year options on euro area inflation. Our methodology is based on the extraction of probability distributions from quotes of European options on inflation, which is easier to perform on zero-coupon derivatives. The realized inflation rate is euro area HICPxT, lagged by three months in order to be known at the maturity date of the option. Unlike the case of stock options, the price of the underlying asset (inflation rate over the maturity of the option) is not observed daily; the fixed leg of a (zero-coupon) inflation swap contract over the same horizon, which is traded daily, is taken as a proxy. We use daily closing quotes of zero-coupon inflation options from the first available trading day, i.e. 5 October, 2009, until 18 February 2015. We consider options with maturity equal to 1, 2, 3, 5, 7 and 10 years. Our sample includes caps with strike rates ranging from 1 to 6% and floors with strikes between -2 and 3%.[4] Zero-coupon inflation swap rates for the same maturities are also taken from Bloomberg.

The degree of liquidity of this class of options is not easy to assess; according to Smith (2012) euro area inflation option markets are more liquid than the UK and US ones. Scharnagl and Stapf (2015), whose analysis is based on the same class of derivative contracts, account for liquidity factors by calculating put-call parities and comparing the evolution of option-implied expected inflation with the fix leg of inflation swaps, for different time horizons; they conclude that information embedded in options data is meaningful in describing aggregate beliefs of market participants on future inflation. In addition, it is worth mentioning that we adopt an estimation methodology (based on Taboga (2015)) which is robust to outliers: pricing errors due to low liquidity, especially in times of financial distress, should not have a significant impact on the results.

The extraction of risk-neutral probability distributions from option quotes is based on the semi-nonparametric method developed in Taboga (2015). In what follows we briefly describe the estimation methodology; see Appendix A for a more detailed description. The probability distribution of future inflation is assumed to have a discrete support;[5] then,

[4] Cap strikes used are: $1.0, 1.5, 2.0, 2.5, 3.0, 3.5, 4.0, 4.5, 5.0, 5.5$ and 6%; the floor strikes are $-2.0, -1.5, -1.0, -0.5, 0.0, 0.5, 1.0, 1.5, 2.0, 2.5$ and 3%.

[5] Assuming that the probability distribution of future inflation is discrete does not reduce significantly the

in the absence of arbitrage opportunities there exists a finite set of positive state prices such that the price of any derivative contract on inflation can be expressed as a function of those state prices. Risk-neutral distributions can be simply obtained by rescaling once state prices are estimated. The method assumes that state prices are interpolated by a spline function, which is proved to be equivalent to a set of linear restrictions. The linearity of the problem allows to derive computationally inexpensive estimators. In particular, a least absolute deviations (LAD) estimator can be obtained through a linear programming problem. In addition, this methodology allows to incorporate unimodality restrictions on the estimators of state prices. Unimodality of risk-neutral distributions obtained from state prices is a very desirable property, and in the previous literature it was not dealt with.[6]

In addition to the computational convenience and the ability to avoid multimodal implied distributions, the advantages of this methodology include its robustness to outliers, which are known to contaminate data on option prices. Despite the lack of information on the liquidity of option quotes, the robustness of the methology supports confidence in the estimated state prices.

For each date in the sample and each maturity horizon, we extract the unimodal LAD estimator of state prices and derive the corresponding risk-neutral distribution, thus getting a time series of implied distributions. For instance, Figure 1 shows the time evolution of risk-neutral distributions of inflation on a 10 years horizon, as extracted from options data in the period September 2011 - February 2015. The plot highlights the tendency of the distributions to become more and more concentrated over time, as well as a shift of the mean towards lower inflation rates.

Throughout the paper, we need to bear in mind that probability distributions extracted from option quotes are risk-neutral by assumption, i.e. they are not adjusted for investors' risk preferences. Risk-neutral probabilities tend to assign more weight to outcomes investors are worried about. Risk-neutral distributions are often more dispersed than real-world ones as inferred by historical data, since they are likely to assign higher probability to extreme events such as deflation or very high inflation. Nevertheless, Bauer and Christensen (2014) point out that risk-neutral probabilities are useful for policy analysis, as policymakers are worried about extreme outcomes just like investors. As stated by Kocherlakota (2013), policy

scope of the methodology; in fact, continuous distributions can be arbitrarily approximated by discrete ones; moreover, most pricing algorithms require discretization at some stage; finally, market prices are inherently discrete.

[6] As explained in Taboga (2015), option-implied risk-neutral distributions are in fact often reported to be multi-modal. Despite the attempt to provide possible explanations for multimodality, it is still considered economically implausible by many economists, since it contradicts the principle that the more extreme an outcome is, the less likely it should be. Multimodality could be an artifact due to estimation procedures rather than an authentic feature of the data.

decision making should take into account the evolution of risk-neutral probabilities, since it reflects changes in market participants' views about future possible outcomes.

Figure 2 shows the mean of the option-implied distributions for maturities of 1, 2, 3, 5, 7 and 10 years. Inflation expectations, as proxied by the expected value of option-implied distributions, have been decreasing since 2012 for all maturities, with sharper falls for shorter horizons. The decline of inflation expectations has followed the fall of realized inflation, which was affected by the collapse of oil prices, weak aggregate demand and the appreciation of the euro. The contraction of investors' beliefs halted around mid January 2015 for all horizons.

Appendix A.2 proves that for a 1-year maturity the expected value of the implied distribution coincides with the fixed leg of an inflation swap having the same maturity. Comparing the time series of expected values derived from our estimates with the quoted inflation rates, we obtain a very accurate match. For maturities longer than 1 year, the quoted inflation swap rate must be equal to a nonlinear function of the implied distribution, and this is also true for our estimates. Figure 3 shows that the difference between the quoted inflation swap rate (red line) and the one implied by our probability distributions (blue line) is negligible for all maturities. Although the estimation methodology we adopt does not force this matching through a constraint, we still recover it in quoted prices: this confirms the robustness and reliability of the approach.

Figure 4 shows the standard deviation of option-implied distributions. The time series of implied volatilities gives insights on the degree of uncertainty in market expectations of future inflation: the higher the standard deviation, the more dispersed are investors' beliefs and/or the more difficult it is to forecast inflation. The figure shows that volatility has been decreasing since 2012 for all maturities, like option-implied means. In the context of long-term inflation expectations departing from the target, the attenuation of uncertainty around those expectations can be seen as an indicator of diminished credibility of monetary policy. In general, the lowering of volatility of inflation distributions is not univocally positive: if long-term expectations are departing from the target, its decrease indicates a higher concentration of beliefs around an undesirable outcome.

Figure 5 shows that the skewness of option-implied risk-neutral inflation expectations for short horizons (1, 2 and 3 years) became negative in the past few months after a gradual decline. For unimodal distributions such as the ones we estimated, a negative skewness indicates that the lower tail is fatter or longer than the upper tail; the recent developments of asymmetry indices for expectations up to 3 years point towards a predominance of the left tail, suggesting that market views are unbalanced towards negative inflation outcomes. Also for maturities of 5, 7 and 10 years the skewness of implied probability distributions has decreased from mid-2013 until January 2015, but it has remained positive.

Figures 2, 4 and 5 highlight that a notable change in financial market participants' inflation expectations has occurred, starting around mid January, 2015: few days before the ECB announced the details of the Quantitative Easing program (January 22) expected values of inflation went up, volatility had a small rebound and skewness increased, especially for longer maturities.

Figure 6 shows the risk-neutral probabilities of the average annual inflation rate over different time horizons falling below zero. Deflation probabilities at all maturities show a sharp increase in the last few months of 2014; for maturities up to 3 years, the rise started earlier – in the last quarter of 2013. Around mid January 2015 the increase halted and deflation probabilities decreased for all time horizons. Symmetric developments can be seen in Figure 7, which shows the risk-neutral probabilities that the average annual inflation rate at time different maturities falls between 1.5 and 2%.

Having estimated option-implied distributions, we can calculate confidence bands around the mean of expected future inflation. This allows to assess the significance of the decline in short- and long-term inflation expectations observed since 2012. Figure 8 shows the confidence bands for the expected value of option-implied probability distributions of future inflation, at the 10% level and time horizons of 1, 5, 7 and 10 years. The upper limit of the confidence band fell below 2% for maturities up to 7 years.[7]

All in all, the evidence presented in this section suggests that a deeper investigation of a potential disanchoring of long-term inflation expectations is needed. Section 3 deals with the study of tail comovement in short-and long term inflation expectations as an additional indicator of disanchoring.

3. Comovement between short and long-term moments

This section further investigates the unusually strong relation between short term and medium-to-long term inflation expectations that has been recently observed in the euro area. For example, the recent fall in inflation expectations at different horizons in the euro area has coincided quite strikingly with a sharp decline in commodity prices, especially oil. This close association looks unusual especially if we look at the fall in medium-to-long term inflation expectations. Even if the recent fall in oil prices were expected to be permanent, it should not have a permanent effect on inflation rates, and thus inflation expectations, over the medium

[7]In Figure 8 we compare the level of inflation expectations with the 2% reference level for illustrative purposes, even though the ECB policy objective entails the inflation rate being below, but close to, 2% over the medium term.

term.[8] Our assessment is that a high sensitivity of medium-to-long term inflation expectations to changes in short term expectations is an indicator of disanchoring. To analyse if there are actual signals of disanchoring in inflation expectations we thus look at the comovement between the moments of short-term inflation and long term inflation, using both a parametric approach, based on copula function, and a non-parametric approach, based on the so-called TailCor measure.[9]

The co-movement between two random variables X and Y can be studied in various ways. One standard measure is the **Pearson correlation**, defined as

$$\rho_{X,Y} = \frac{Cov(X,Y)}{\sigma_X \sigma_Y},$$

where $Cov(X,Y)$ is the covariance between the two variables, and σ_X and σ_Y are the related standard deviations. However, there are some important limits of such a measure: first of all, correlation can't be used to study cross dependence between variables, because zero cross correlation does not in general imply independence; moreover, it cannot be defined for certain distributions where the first two moments are not finite; even with data approximately normally distributed, it does not distinguish fat or long tails, nor correlation between positive and negative values. Finally, this measure is strongly influenced by outliers, which would compromise the interpretationn of the results.

Moreover, in the context of our assessment of potential signals of disanchoring of inflation expectations by looking at comovement between short term and long term inflation moments, looking at Pearson correlation may not be sufficient: in fact, it would provide only an average indication of co-movement without information on which variations (positive or negative) present a higher co-movement. For example, an increase in the correlation between short and long term means could be caused by an increased correlation in the right tail rather than in the left tail, and thus would not be in line with the hypothesis of downward disanchoring from the target.

These are the main reasons why in this paper we study the co-movement between two variables looking at the **conditional tail dependence**, that is defined as follows:[10]

Definition 1 (Conditional upper and lower tail dependence)**.** *Let X and Y be two random variables with marginal distributions $F_X(x)$ and $F_Y(y)$. Let x_k denote the quantile k of*

[8]For a discussion about indirect and second round effects of the drop in oil prices onto inflation, see Box 5 in ECB Economic Bulletin, Issue 1, 2015 and Box 3 in ECB Monthly Bulletin, December 2014.

[9]Even though further rigourous tests of causality are not conducted, we propose an interpretation of the relationship between short and long-term expectations in terms of shock transmission and persistence of macroeconomic news shocks.

[10]The coefficient of tail dependence was first introduced in the finance literature by Embrechts et al. (2003).

variable X, that is, the value of x that solves equation $F_X(x) = k$, and let y_k be the corresponding quantile for Y. The conditional upper tail dependence is defined as

$$\lambda_U = \lim_{k \to 1} Pr\{Y > y_k | X > x_k\};$$

the conditional lower tail dependence is defined as

$$\lambda_L = \lim_{k \to 0} Pr\{Y \leq y_k | X \leq x_k\}.$$

Basically the tail dependence looks at the probability that two variables co-move when relatively large changes occur in one of them.

3.1. Copula-based dependence measure

An easy way to compute tail dependence is by means of copula functions.[11] Copula functions are a special class of multivariate cumulative distribution functions which allows to separate the modeling of the marginal distributions from the dependence structure between the variables. Using a copula involves specifying any shape of marginal cumulative distribution functions of each random variable and the copula function that connects them and implies a certain shape for the dependence between them.[12]

The data for which we want to compute the tail dependence are the spot and forward inflation swap rates and the standard deviations derived from inflation options, at different horizons.[13] In order to fit copula functions to the data, they must be transformed into approximately uniformly distributed variables. In fact, our original data may not to lie on the $[0, 1]$ interval and may be temporally correlated. To clean up the data the necessary transformation involves three steps for each pair of variables:

1. First differencing the two variables to detrend them and make the series stationary;
2. Filtering the resulting variables through an AR(1) model for the conditional mean and a GARCH (1,1) specification for the variances, to eliminate persistence or heteroskedasticity;[14]

[11] Other measures of dependence which are copula-based are the bivariate measures Kendall's tau and Spearman's rho, which can be expressed in terms of the underlying copula alone (see, among others, Aas (2004)).

[12] See Appendix B for copula definition and main properties, and Nelsen (2006) for a detailed exposition of the theory and practical aspects of copulas.

[13] Without further assumptions on the inflation process, it is not possible to retrieve forward inflation densities (and their moments) from zero-coupon inflation options, the most liquid derivatives on inflation.

[14] A similar approach is adopted by Christoffersen et al. (2012)

3. Mapping the standardized daily revisions in inflation expectations into numbers between 0 and 1 through the computation of an empirical marginal cumulative distribution function $\tilde{F_X}$, so that $u_t = \tilde{F_X}(x_t)$, and $\tilde{F_Y}$, so that $v_t = \tilde{F_Y}(y_t)$: the variables u and v thus obtained are approximately uniformly distributed.

The filtering procedure and the zero-one mapping described above are conducted not only for mean inflation expectations (proxied by inflation swap spot and forwards), but also for option-implied standard deviations. Summary statistics for levels, daily changes, filtered values and mapped values of each variable are reported in Table 1; variables are divided in two groups (Mean expectations and standard deviations), and groups are further divided into subgroups of different maturities. Considering the levels – line 1 of each subgroup –, the volatility of mean expectations decreases as the maturity increases; this is also true for the volatility of option-implied standard deviations. In general, the variability of mean expectations and standard deviations are quite comparable with each other: this provides further motivation for an investigation of tail comovements in higher moments of the distributions. Level variables, both means and standard deviations, are all strongly persistent: high autocorrelations in first differences are removed by the AR-GARCH filtering (see line 3 of each subgroup).

Once we have cleaned up and transformed the data as described above, we can estimate copulas in rolling windows, using the available data of the previous 200 calendar days. Different copulas allow for different dependence structures. We consider the **Student's t copula**, which belongs to the class of elliptical distributions, and displays symmetric tail dependence and potentially very heavy tails.[15] In general, for elliptical distributions, $\lambda_U(X,Y) = \lambda_L(X,Y)$, and in particular, for the Student's t-copula, the coefficients of lower and upper tail dependence are

$$\lambda_U(X,Y) = \lambda_L(X,Y) = 2t_{\nu+1}\Big(-\sqrt{\nu+1}\sqrt{\frac{1-\rho}{1+\rho}}\Big),$$

where $t_{\nu+1}$ denotes the distribution function of a univariate Students t-distribution with $\nu+1$ degrees of freedom and ρ is the linear correlation. The stronger the linear correlation ρ and the lower the degrees of freedom ν, the stronger is the tail dependence. However, the Students t-copula gives asymptotic dependence in the tail, even when ρ is negative and zero.

[15]See Appendix B for the definition of this copula. Different copulas that display different tail dependence could be also used: for example, the Gumbel copula displays upper tail dependence, while the rotated Gumbel copula displays lower tail dependence.

3.2. *TailCor dependence measure*

As an alternative to the parametric tail dependence measure implied by copulas, we consider the **TailCor** measure, a non parametric metric for tail correlations, introduced by Ricci and Veredas (2013). This measure can be implemented under mild assumptions and presents several advantages:

- it is not dependent on specific distributional assumptions;
- it allows to disentangle whether the evidence of tail correlations is caused by variables which are linearly correlated and/or nonlinearly correlated, in the sense that they are dependent only at the extremes;
- it is exact for any cut-off point of the tail;
- it can be computed for tails that are fatter, equal or thinner than those of the Gaussian distribution;
- it performs well also in small samples, without relying on asymptotic theory.

In the following we briefly explain the intuition underlying this dependence measure. The formal definition is provided in Appendix C, while technical details can be found in Ricci and Veredas (2013).

Let X_t $t = 1, \dots T$ be a random vector of size N at time t satisfying standard assumptions[16]. Let us explain the TailCor measure from a graphical point of view. The intuition underlying TailCor is that if two standardized random variables X_j and X_k are positively related (either linear and/or non-linearly), most of the times the pair of observations have the same sign. This means that if we look at the scatter plot of the random variables (see Figure 9 taken from Ricci and Veredas (2013)) most of the pairs of observations (depicted with dots) concentrates in the north-east and south-west quadrants. If we project all the pairs on the 45-degree line that diagonally crosses the quadrants, we get a new random variable $Z^{(jk)}$ (depicted with squares). Since the two random variables are positively related, the projected dots are dispersed all over the line. The degree of dispersion depends on the strength of the relationship between the two random variables: if the relation is strong, the cloud is stretched around the 45-degree line and the projected dots are very dispersed. If on the contrary the relation is weak, the cloud of dots is sparsed around the origin, without having a well defined direction.

Figure 10 illustrates an example of the dependence patterns in our data. It shows the scatterplots of the filtered innovations of the 1y1y and 5y5y forward rates in different time intervals. The left panel refers to a period (May 2011 - Feb 2012) in which there is little

[16] They are stationarity and mixing assumptions, see Ricci and Veredas (2013) for details.

tail dependence, whereas the right panel (corresponding to the period May 2014 - Feb 2015) shows dependence in both the upper and the lower tail.

The TailCor measure is equal - up to a normalization - to the difference between upper and lower tail quantiles of $Z^{(jk)}$, denoted by $IQR^{(jk)\xi}$. To separately focus in the tail of one side of the distribution, we also look at the **downside TailCor** (DownTailCor) and **upside TailCor** (UpTailCor) measures, respectively defined in terms of the difference between the median and the lower tail quantile, and the difference between the upper tail quantile and the median.[17] We compute the TailCor measures for the transformed uniformly distributed data.[18] Theoretically, the TailCor index takes values between 0 and infinity; however, the actual range of variation in most financial applications is very small. The fatter the tails of the bivariate distribution, the higher the exceedance of the upper bound over the value of square root of 2 (i.e., the largest value under a bivariate Gaussian).

3.3. Results

In this subsection we report results based on three categories of statistics of comovement: the Pearson's ρ correlation coefficient (average comovement); the coefficient of tail comovement estimated with the Student's t static bivariate copula and the TailCor index (two measures of average comovement in the tails); the UpTailCor and DownTailCor indices which track comovements over time in upper and lower tails, respectively.[19] The Student's t copula is preferred to other copula distributions because of its good fit to inflation swap data in terms of log-likelihood, AIC and BIC criteria (see Antunes (2015)). Every statistic is computed in-sample using rolling windows of 200 business days of observations (about 5 months); nonetheless, the conclusions we draw are robust to different window lengths.

Results on the comovement of mean expectations are depicted in Figures 11 to 14. Specifically, we examine comovements between medium-to-long term mean expectations (expectations 5 years ahead after 5 years) and those for shorter horizons, up to 5 years ahead. Mean expectations for 1, 2 and 3 years ahead are proxied by the 1-year spot inflation swap rate, the 1y1y forward and the 1y2y forward (i.e., expectations 1 year ahead after 2 years), respectively. Pearson's coefficient in Figures 11 shows a decline in the average correlation between 5y and 5y5y expectations during 2013 and the first half of 2014; a steady increase is

[17]The formal definition can be found in Appendix C.

[18]We use the Matlab code provided by the authors on their website. A sketch of the estimation procedure can be found in Appendix C.

[19]Results coming from the estimates based on the Student's-t copula and the TailCor index have almost the same interpretation in terms of dynamics, even though the TailCor index is the sum of the comovement in the two tails, while the upper tail index assumes symmetric marginal distributions, so the estimate of upper and lower tail comovement is symmetric.

then evident from end July 2014 up to levels close to 60 per cent (see picture on the lower-right corner). A closer look at the evolution of the ρ coefficients shows that the downward trend of comovements started later on for shorter term expectations: for instance, it started only at the beginning of 2014 for the 1y-5y5y correlation (see picture on the upper-left corner); the following increase is then driven by a rise in the correlation between the shortest end of the swap term structure (1y and 1y1y) and medium-to-long term expectations.

These results points to a recent increase in average correlations. To investigate possible further signs of disanchoring, we look at the path of the Student's t coefficient of tail comovement and the one of the TailCor index (Figures 12 and 13): both statistics suggest that the observed increase in the average correlation reflects, at least in part, an increased correlation in the tails. Here, too, the evidence in the 5y-5y5y comparison is found to be driven by the comovement of very short term expectations with medium-to-long term ones; the observed dependence starts to increase slightly before the rise in average correlations (two months before). These findings motivate a closer look at each tail separately, using the two variants of the TailCor: Figure 14 depicts the dynamics of the correlation in the upper tail (blue line) and lower tail (red line) proxied by the UpTailCor and DownTailCor. Interestingly, in the 5y vs 5y5y comparison, the DownTailCor index increases more than the UpTailCor since mid 2014: during the second part of the year, the difference in levels between the two indices becomes remarkable. This evidence implies that the correlation in the lower tail has increased more than the one in the upper tail, reaching its sample maximum, so negative events affecting short-term views have been transmitted to long-term expectations more than positive surprises. This stylized fact suggests that some disanchoring may be occurring, and further investigation is needed.

Figures 15 to 18 trace the comovement in option-implied standard deviations. Since forward second moments are not available, we proxy the short-end of the term structure with 1-year and 2-year spot values, and the long-end with 7- and 10-year standard deviations. The fact that spot long-term values are not independent from short-term ones entails some form of comovement being built-in, leading to potentially misleading interpretations of the level of the dependence; however, we are mostly interested in the evolution of the correlations over time rather than the level, so in this sense our investigation is totally meaningful. Based on the rolling estimates of the Pearson's ρ, there is no clear evidence of increased correlation between short- and long-run standard deviations during the last part of the sample (Figure 15). Interpreting the standard deviation as the uncertainty of market agents around their mean expectations, this means that the transmission of average uncertainty towards longer maturity has not increased through time.

However, the analysis of tail comovements suggests a different conclusion. While the dynamics of the TailCor index for short vs long-term uncertainties in the last period is not robust to the choice of the employed proxies (see Figure 17), the copula-based coefficient of tail dependence suggests a strong recent increase in tail comovements (e.g., 1y-10y and 2y-10y couples – upper-right and lower-right picture of Figure 16). Moreover, looking at UpTailCor and DownTailCor, we end up with opposite results with respect to the one obtained about mean expectations: the UpTailCor (proxying comovements in the upper tail) increases in the last part of the sample, while lower tail correlations (the DownTailCor) have mixed dynamics (see Figure 18); this divergence is clearly observed in the 1y vs. 10y and the 2y vs. 10y comparisons. The evidence from the term structure of option-implied standard deviations suggests a possible transmission from upper tail variations of short-term uncertainty to upper tail variations in long-term one; this means that positive shocks to uncertainty are persistent while shocks reducing uncertainty are short-lived.

To sum up, the joint reading of the results for average expectations and option-implied standard deviations gives a number of clear indications. From the second half of 2014, negative tail events like bad macro news or worse-than-expected data readings have been increasingly transmitted to long-term views, igniting downward revisions in average expectations and upward revisions in uncertainty. This strenghtened transmission has not concerned "positive" surprises to short-term expectations: facing those events, long-term expectations are found to be quite inelastic. We conclude that there actually are signs of disanchoring from below of long-term inflation expectations from the "below but close to" 2 percent target.

4. Conclusions

In this paper we propose a new method to detect possible signs of disanchoring of inflation expectations from the medium-to-long term objective of the monetary authority. Unlike the commonly used news-regression approach, our technique is totally market-based and does not require any identification of the surprise component incorporated in macro news or inflation readings. Applying the new estimation technique of Taboga (2015) to daily quotes of inflation caps and floors for the euro area, we are able to recover the term structures of means and standard deviations of inflation expectations from October 2009 up to present. We fully exploit the information coming from these estimates by looking at their historical behaviour and making comparisons between the long- and short-end of each term structure. To achieve this, we compute linear correlations and measures of tail comovement based on the theory of copulas and on the non-parametric TailCor indexes. We find that, from mid-2014, the tail comovement between short- and long-term mean expectations as well as that between the short- and long-term uncertainty around these expectations has gradually increased, signalling an increased transmission of extreme shocks hitting short-term beliefs to the distribution of long-term inflation expectations. Even more importantly, this transmission of tail shocks across the term structure is significantly asymmetric, given that long-term mean expectations have shown a tendency to follow strong downward revisions of short-term expectations while they have been quite inelastic to positive changes. Moreover, events raising uncertainty about short-term prospects of inflation are likely to also produce increases in long-term standard deviations, while declines in short-term uncertainty seem to be short-lived. The evidence based on the evolution of option-implied confidence bands around mean expectations is then confirmed by the analysis of comovements; this leads to conclude that some signs of disanchoring from below of long-term inflation expectations from the 2 per cent target are there and should not be overlooked.

While short- and long-term market-based measures of inflation expectations in the euro area have been unusually low in a historical and cross-country comparison already since 2013, in the last few months 5y5y forward inflation swap rates have also declined in the United States and, more recently, in the United Kingdom. Even though broader international trends may in part be responsible of this comovement, the underlying expected persistence of low inflation could be different in the three economies. A further avenue of research could be that of extending the estimation of option-implied distributions of future inflation to US and UK data, and assess tail comovements and spillovers across countries.

Appendix A. LAD method

A.1. *Extraction of option-implied distributions*

In what follows, a quick description of the estimation method to derive option-implied probability distribution functions elaborated in Taboga (2015). Let $\mathcal{I}$ be the stochastic value of the average annual inflation rate over a given time horizon. We assume that $\mathcal{I}$ has a discrete probability distribution with finite support, $R_{\mathcal{I}} = \{i_1, \ldots, i_n\}$. In the absence of arbitrage, there exists a n-dimensional vector of positive state prices $\pi = (\pi_1, \ldots, \pi_n)$ such that the price $\Pi(f)$ of any derivative contract on $\mathcal{I}$ having payoff $f = f(\mathcal{I})$ can be written as

$$\Pi(f) = \sum_{j=1}^{n} \pi_j f(i_j).$$

In particular, the price of a zero-coupon cap with strike k and maturity $T = M$ years is given by

$$\sum_{j=1}^{n} \pi_j((1+i_j)^M - (1+k)^M)^+,$$

whereas the price of a zero-coupon floor with strike k and maturity $T = M$ years equals

$$\sum_{j=1}^{n} \pi_j((1+k)^M - (1+i_j)^M)^+.$$

Suppose we have the market quotes of N_C caps with strikes $\{k_1^C, \ldots, k_{N_C}^C\}$ and N_P floors with strikes $\{k_1^P, \ldots, k_{N_P}^P\}$. Let C be the $N_C \times 1$ vector of cap quotes and P be the $N_P \times 1$ vector of put quotes. Let F_C and F_P be $N_C \times n$ and $N_P \times n$ matrices of payoffs, defined as

$$F_{C,ij} = (s_j - K_i^C)^+ \quad \text{and} \quad F_{P,ij} = (K_i^P - s_j)^+.$$

Having set

$$Y = \begin{bmatrix} C \\ P \end{bmatrix} \text{ and } X = \begin{bmatrix} F_C \\ F_P \end{bmatrix},$$

we can express the option prices as

$$Y = X\pi. \tag{1}$$

Since in practice market quotes encompass an error term,[20] the empirical version of Equation (1) is

$$Y^0 = X\pi + \varepsilon, \tag{2}$$

where Y^0 is the vector of observed market prices and ε is a vector of pricing errors. Our goal is to estimate the vector of positive state prices π given that we observe Y^0 and we know the payoff X; the risk neutral probability distribution of $\mathcal{I}$ is then obtained by rescaling,

$$d = \frac{\pi}{\sum_{j=1}^{n} \pi_j}. \tag{3}$$

State prices are parametrized using a spline curve; Taboga (2015) shows that this is equivalent to imposing a set of linear equality restrictions. With no loss of generality we assume that the support $R_{\mathcal{I}}$ of the distribution is equally spaced:

$$i_j = i_1 + (j-1)\delta, \quad \delta > 0 \text{ and } j = 1, \ldots, n.$$

Moreover, we assume that there exists a (piece-wise cubic and twice continuously differentiable) spline function $\overline{\pi} : [s_1, s_n] \mapsto \mathbb{R}_+$ which interpolates the state prices:

$$\overline{\pi}(s_j) = \pi_j, \quad , j = 1, \ldots, n.$$

The number of knot points of the spline function $\overline{\pi}$ is $N_T < n - 4$; the first four elements of $R_{\mathcal{I}}$ cannot be knot points. De Boors (1978)'s B-spline construction implies that the first derivative of $\overline{\pi}$ is piecewise quadratic, the second derivative is piecewise linear, the third is a stepwise constant function and the fourth is a function that is zero everywhere except at knot points. The latter condition translates into a linear constraint on the state prices associated with knot points:

$$ND\pi = 0, \tag{4}$$

[20]Pricing errors can arise for various reasons, including the bounce between bid and ask quotes, price discreteness and slate prices due to illiquidity.

where N is a $(n-4-N_T)\times(n-4)$ selection matrix whose rows are vectors of the euclidean basis of $\mathbb{R}^{n-4}$, $D = D_{n-3}D_{n-2}D_{n-1}D_n$ and D_k is the $(k-1)\times k$ first difference matrix

$$D_k = \begin{bmatrix} -1 & 1 & 0 & 0 & \cdots & 0 \\ 0 & -1 & 1 & 0 & \cdots & 0 \\ \vdots & & & & & \vdots \\ \vdots & & & & & \vdots \\ 0 & 0 & \cdots & 0 & -1 & 1 \end{bmatrix}. \tag{5}$$

The LAD estimator of state prices is based on Equation (2) and on the set of linear restrictions (4). An estimator minimizing absolute pricing errors is preferred to a least squares estimator because of its computational convenience and robustness to outliers, which are known to contaminate data on option prices. The LAD estimator $\hat{\pi}_{LAD}$ of the state prices is the solution of the minimization problem

$$\begin{aligned} \hat{\pi}_{LAD} \quad &= \quad \arg\min_{\pi} \sum_{i=1}^{N_C+N_P} w_i |Y_i^0 - X_i\pi| \\ &\quad\ \text{s.t.} \quad ND\pi = 0, \pi \geq 0 \end{aligned} \tag{6}$$

where Y_i^0 and X_i are the rows of Y^0 and X respectively and w_i are weights assigned to pricing errors. In our estimates we set $w_i = 1/\sqrt{Y_i^0}$: this choice applies a dampening factor to deeply out-of-the-money options, which tend to have larger pricing errors in percentage terms.

The minimization problem can be written as a linear programming (LP) problem:

$$\begin{aligned} &\min_z d^T z \\ &\text{s.t.} \quad Az = b, z \geq 0 \end{aligned} \tag{7}$$

where

$$d = \begin{bmatrix} w \\ w \\ 0 \end{bmatrix} \quad \text{and} \quad z = \begin{bmatrix} \varepsilon^+ \\ \varepsilon^- \\ \pi \end{bmatrix}$$

are $(2N_C + 2N_P + n)\times 1$ vectors, w is the $(N_C+N_P)\times 1$ vector of weights, ε^+ and ε^- are the positive and negative parts of the $(N_C+N_P)\times 1$ vector of pricing errors and

$$d = \begin{bmatrix} I & -I & X \\ 0 & 0 & ND \end{bmatrix} \quad \text{and} \quad b = \begin{bmatrix} Y^0 \\ 0 \end{bmatrix}.$$

The solution of the LP problem can be found by standard and computationally inexpensive LP algorithms. The LAD estimator $\hat{\pi}_{LAD}$ is then given by the last n components of the LP solution $\hat{z}$.

Once we have computed the LAD estimator $\hat{\pi}_{LAD}$, we can get a new estimator $\hat{\pi}_U$ fulfilling a unimodality condition. Since the risk-neutral distributions are obtained by rescaling the state prices, the unimodality of π implies the unimodality of ρ. Let

$$\varphi(\pi) = \arg\max_i \pi_i \quad \text{and} \quad g(\pi) = (g_1(\pi), \ldots, g_{n-1}(\pi)) \text{ s.t. } g_i(\pi) = \begin{cases} 1 & \text{if } i < \varphi(\pi) \\ -1 & \text{if } i \geq \varphi(\pi) \end{cases}$$

The set of vectors which satisfy unimodality is $U = \{\pi \in \mathbb{R}^n_+ : (D_n\pi) \circ g(\pi) \geq 0\}$, where D_n is the first-difference matrix defined in (5) and $\circ$ denotes the Hadamard or entrywise product. The unimodal LAD estimator $\hat{\pi}_U$ is then the solution of the minimization problem

$$\begin{aligned} &\min_{\pi} \sum_{i=1}^{N_C+N_P} w_i |Y_i^0 - X_i\pi| \\ &\text{s.t.} \quad ND\pi = 0, \pi \geq 0 \text{ and } \pi \in U. \end{aligned} \tag{8}$$

In order to compute $\hat{\pi}_U$, we solve problem (7) and derive $\hat{\pi}_{LAD}$, then we set $h = \varphi(\hat{\pi}_{LAD})$ and run a second LAD estimation imposing the unimodality condition through a set of $n-1$ additional linear inequality constraints: $\pi_{i-1} \leq \pi_i$ for $1 < i \leq h$ and $\pi_{i-1} \geq \pi_i$ for $h < i \leq n$.

A.2. Inflation swap rates in terms of state prices

Let s_M be the fixed leg of a zero-coupon inflation swap with maturity M years. Let $\mathcal{I}_M$ be the (stochastic) annual rate of inflation over the next M years. Taking expectations under the risk-neutral measure Q, the following condition must hold:

$$\mathbb{E}_0^Q[D_M((1+s_M)^M - (1+\mathcal{I}_M)^M)] = 0, \tag{9}$$

where D_M is the discount factor for the time interval $[0, M]$. Re-writing Equation (9) in terms of state prices and taking into account that $\mathbb{E}_0^Q[D_M] = \sum_j \pi_j$, we get

$$\Big(\sum_{j=1}^{n} \pi_j\Big)(1+s_M)^M - \sum_{j=1}^{n}(1+i_j)^M \pi_j = 0.$$

Since the risk neutral distribution d is given by $d_j = \pi_j / \sum_k \pi_k$, there follows that

$$s_M = \Big(\sum_{j=1}^{n} d_j (1+i_j)^M \Big)^{1/M} - 1. \tag{10}$$

For $M = 1$, this equivalence boils down to

$$s_1 = \sum_{j=1}^{n} d_j i_j \,;$$

the inflation swap rate equals the mean of the option-implied distribution d. For $M > 1$, Equation (10) states that the inflation swap rate is a nonlinear function of the probability distribution extracted from inflation options having the same maturity.

Appendix B. Copula functions

Definition 2 (Copula function). *A copula is an n-dimensional distribution function $C : [0,1]^n \to [0,1]$ of a random vector $(U_1, \ldots, U_n)$, where the marginal law of U_i is the uniform distribution on $[0,1]$ for all $i \in \{1, \ldots, n\}$.*

Copula functions are very popular in the study of multivariate distribution functions thanks to their role in imposing a dependence structure on predetermined marginal distributions. Their importance derives from *Sklar's theorem*, which proves that any multivariate distribution function can be characterized by a copula and that copula functions, together with univariate marginal distribution functions, can be used to construct multivariate distribution functions.

Theorem B.1 (Sklar's theorem). *Let H be an n-dimensional distribution function with marginals $F_1, \ldots, F_n$.*

Then an n-copula C exists such that, for each $\mathbf{x} \in \mathbf{R}^n$,

$$H(x_1, \ldots, x_n) = C(F_1(x_1), \ldots, F_n(x_n)).$$

If the marginals $F_1, \ldots, F_n$ are all continuous, then C is unique; otherwise C is univocally determined on $(RanF_1 \times RanF_2 \times RanF_n)$ (where $RanF_i$ denotes the rank of F_i). Conversely, if C is an n-copula and $F_1, \ldots, F_n$ are distribution functions, then the function H defined above is an n-dimensional distribution function with marginals $F_1, \ldots, F_n$.

The proof of this theorem can be found e.g. in Nelsen (2006).

The main feature of Sklar's theorem is that for continuous multivariate distribution functions, the univariate marginals and the multivariate dependence structure can be separated and the dependence structure can be represented by a copula.

Let F be an univariate distribution function. Let us recall that the generalized inverse of F is defined as $F^{-1}(t) = \inf\{x \in \mathbf{R} | F(x) \geq t\}$ for each t in $[0,1]$, with the usual convention that $\inf(\emptyset) = -\infty$.

An important corollary of Sklar's theorem, which is fundamental in the study of copulas and their applications, is the following:

Corollary 1. *Let H be an n-dimensional distribution function with continuous marginals $F_1, \ldots, F_n$ and copula C. Then for each $\mathbf{u} \in [0,1]^n$,*

$$C(u_1, \ldots, u_n) = H(F_1^{-1}(u_1), \ldots, F_n^{-1}(u_n)).$$

In the following we recall the **Student's t copula** that we use in the paper.

Definition 3 (Student's t copula)**.** *The* ***Student's t copula*** *can be written as*

$$C_{\rho,\nu}(u,v) = \int_{-\infty}^{t_\nu^{-1}(u)} \int_{-\infty}^{t_\nu^{-1}(v)} \frac{1}{2\pi(1-\rho^2)^{1/2}} \{1 + \frac{x^2 - 2\rho xy + y^2}{\nu(1-\rho^2)}\}^{-(\nu+2)/2} dsdt,$$

where ρ and ν are the parameters of the copula, and t_ν^{-1} is the inverse of the standard univariate Student's t-distribution with ν degrees of freedom, expectation 0 and variance $\frac{\nu}{\nu-2}$.

Student's t copula allows for joint fat tails. Increasing the value of ν decreases the tendency to exhibit extreme co-movements. The Student's t-dependence structure supports joint extreme movements regardless of the marginal behaviour of the individual variables.

A copula which exhibits greater dependence in the positive tail is the **Gumbel copula**, defined as follows:

Definition 4 (Gumbel copula)**.** *The* ***Gumbel copula*** *is an asymmetric copula, given by*

$$C_\delta(u,v) = exp(-[(-\log u)^\delta + (-\log v)^\delta]^{1/\delta}),$$

where $0 < \delta \leq 1$ is a parameter controlling the dependence.

Perfect dependence is obtained if $\delta \to 1$, while $\delta = 0$ implies independence.

On the contrary, a copula characterized by a greater dependence in the negative tail is the **Rotated Gumbel copula**, that is equivalent to the Gumbel copula computed in $(1-u, 1-v)$.

Appendix C. TailCor measures

Let X_{jt} be the jth element of the random vector $\mathbf{X}_t$. Denote by Q_j^τ its τth quantile for $0 < \tau < 1$, and let $IQR_j^\tau = Q_j^\tau - Q_j^{1-\tau}$ be the τth interquantile range. Let Y_{jt} be the standardized version of X_{jt}:

$$Y_{jt} = \frac{X_{jt} - Q_j^{0.50}}{IQR_j^\tau}.$$

By standard trigonometric arguments, the projection of (Y_{jt}, Y_{kt}) onto the 45-degree line is

$$Z_t^{(jk)} = \frac{1}{\sqrt{2}}(Y_{jt} + Y_{kt}),$$

and the tail interquantile range is

$$IQR^{(jk)\xi} = Q^{(jk)\xi} - Q^{(jk)1-\xi},$$

where $Q^{(jk)\xi}$ is the ξth quantile of Z_t^{jk}. The larger ξ is, the further we explore the tails.

TailCor is then defined as follows (Ricci and Veredas (2013)):

Definition 5 (TailCor)**.** *Under technical assumptions, TailCor between X_{jt} and X_{kt} is*

$$TailCor^{(jk)\xi} := s_g(\xi, \tau) IQR^{(jk)\xi},$$

where $s_g(\xi, \tau)$ is a normalization such that under Gaussianity and linear uncorrelation $TailCor^{(jk)\xi} = 1$, the reference value.

A table with values of $s_g(\xi, \tau)$ for a grid of reasonable variables for τ and ξ can be found in Ricci and Veredas (2013), Appendix T.

When interest lies in the tail of one side of the distribution, downside TailCor and upside TailCor can be used:

Definition 6 (Downside TailCor)**.** *Downside TailCor is defined as*

$$TailCor^{(jk)\xi-} := s_g(\xi, \tau) IQR^{(jk)\xi-},$$

where $IQR^{(jk)\xi-} = Q^{(jk)0.50} - Q^{(jk)1-\xi}$.

Definition 7 (Upside TailCor)**.** *Upside TailCor is defined as*

$$TailCor^{(jk)\xi+} := s_g(\xi, \tau) IQR^{(jk)\xi+},$$

where $IQR^{(jk)\xi+} = Q^{(jk)\xi} - Q^{(jk)0.50}$.

The estimation procedure consists of four simple steps that can be followed under technical assumptions:

1. Standardize X_{jt} and X_{kt};
2. Estimate the IQR of the projection: $I\hat{Q}R_{\hat{Z},T}^{(jk)\xi}$;
3. Find the normalization $s_g(\xi, \tau)$ from the table;
4. Compute $Tai\hat{l}Cor_{\hat{Z},T}^{(jk)\xi} = s_g(\xi, \tau) I\hat{Q}R_{\hat{Z},T}^{(jk)\xi}$.

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Table 1: Summary Statistics

	Obs.	Average	Volatility	Min	Max	Autocorr
Mean expectation						
π_{1y}	1404	1.36	0.60	-0.77	2.68	1.00
$\Delta\pi_{1y}$	1403	-0.00	0.06	-0.52	0.45	-0.12
x_{1y}	1403	-0.00	1.00	-8.30	8.93	-0.01
u_{1y}	1403	0.50	0.29	0.00	1.00	-0.02
π_{1y1y}	1404	1.47	0.41	0.29	2.56	0.99
$\Delta\pi_{1y1y}$	1403	-0.00	0.06	-0.35	0.43	-0.33
x_{1y1y}	1403	-0.00	1.00	-4.43	11.10	-0.05
u_{1y}	1403	0.50	0.29	0.00	1.00	0.00
π_{1y2y}	1404	1.64	0.38	0.53	2.48	0.99
$\Delta\pi_{1y2y}$	1403	-0.00	0.06	-0.41	0.51	-0.40
x_{1y2y}	1403	-0.01	1.00	-8.70	6.93	-0.05
u_{1y1y}	1403	0.50	0.29	0.00	1.00	0.04
π_{5y}	1404	1.62	0.37	0.44	2.36	1.00
$\Delta\pi_{5y}$	1403	-0.00	0.03	-0.14	0.14	0.01
x_{5y}	1403	-0.01	1.00	-5.36	4.55	-0.00
u_{5y}	1403	0.50	0.29	0.00	1.00	0.02
π_{5y5y}	1404	2.24	0.22	1.48	2.79	0.99
$\Delta\pi_{5y5y}$	1403	-0.00	0.03	-0.19	0.15	-0.16
x_{5y5y}	1403	-0.00	1.00	-3.98	6.00	0.01
u_{5y5y}	1403	0.50	0.29	0.00	1.00	0.05
Option-implied standard deviations						
σ_{1y}	1404	1.26	1.01	0.66	11.69	0.61
$\Delta\sigma_{1y}$	1403	-0.00	0.89	-10.82	10.84	-0.20
x_{1y}^{sd}	1403	0.00	1.01	-6.47	6.45	-0.03
u_{1y}^{sd}	1403	0.50	0.29	0.00	1.00	0.01
σ_{2y}	1404	1.30	0.74	0.71	11.69	0.74
$\Delta\sigma_{2y}$	1403	-0.00	0.54	-10.05	10.31	-0.26
x_{2y}^{sd}	1403	-0.00	1.00	-10.51	6.38	-0.03
u_{2y}^{sd}	1403	0.50	0.29	0.00	1.00	-0.01
σ_{7y}	1404	1.46	0.39	0.57	7.76	0.80
$\Delta\sigma_{7y}$	1403	-0.00	0.24	-6.12	6.16	-0.01
x_{7y}^{sd}	1403	0.01	1.01	-7.18	14.18	0.02
u_{7y}^{sd}	1403	0.50	0.29	0.00	1.00	0.10
σ_{10y}	1404	1.51	0.30	0.57	2.09	0.98
$\Delta\sigma_{10y}$	1403	-0.00	0.06	-1.16	1.16	-0.21
x_{10y}^{sd}	1403	0.04	1.00	-11.67	8.63	-0.03
u_{10y}^{sd}	1403	0.50	0.29	0.00	1.00	0.06

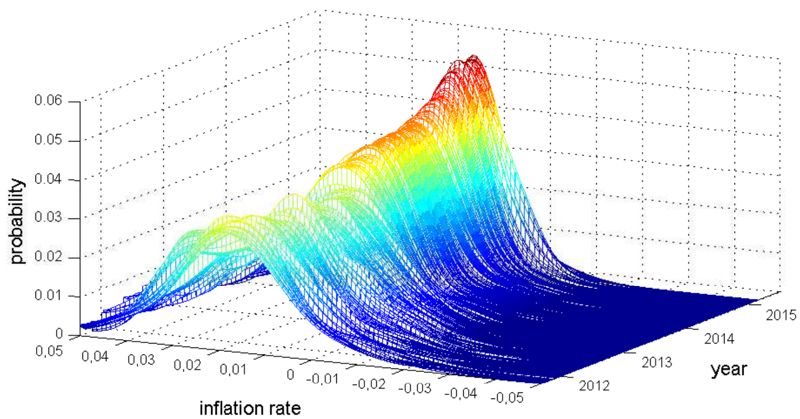

Figure 1. Option-implied risk-neutral distributions of annual euro area inflation over a 10 year horizon, extracted using the LAD estimator with unimodal restrictions from daily quoted between September 2011 and February 2015. The x-axis corresponds to the annual inflation rate (percentage points); the y-axis indicates the time interval (days).

Figure 2. Means of option-implied risk-neutral inflation distributions; percentage points. Option-implied distributions are extracted using the LAD estimator with unimodality restriction. Daily quotes of inflation caps and floors from October 2009 to February 2015 are taken from Bloomberg.

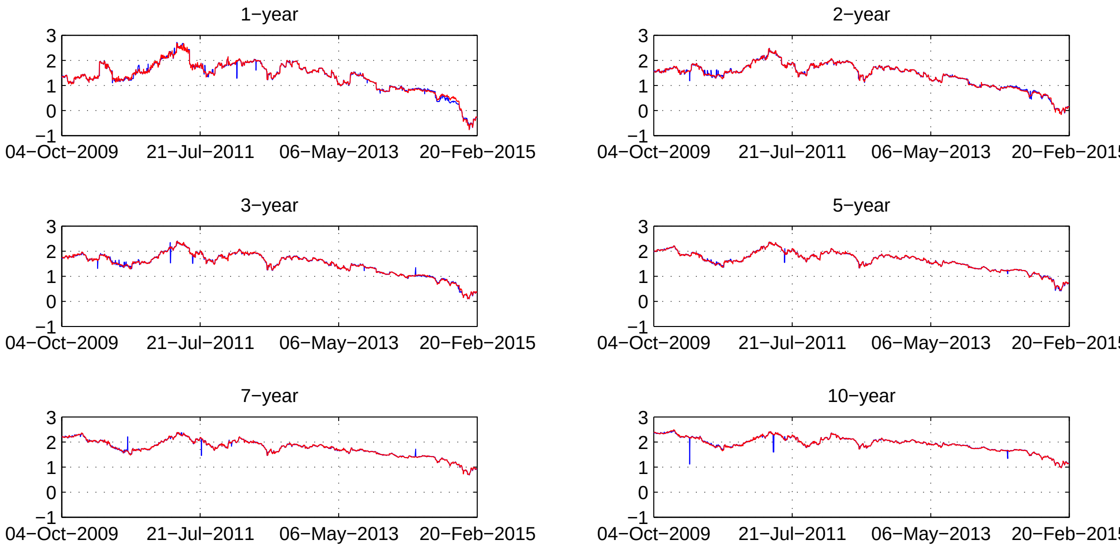

Figure 3. Market quotes of inflation swap rates (red line) at maturities 1, 2, 3, 5, 7 and 10 years and inflation swap rates implied by the probability distributions embedded in option prices (blue line) at the same maturities. Option-implied distributions are extracted using the LAD estimator with unimodal restrictions. Daily quotes of inflation swaps and inflation options from October 2009 to February 2009 are taken from Bloomberg.

Figure 4. Standard deviations of option-implied risk-neutral inflation distributions at maturities of 1, 2, 3 years (upper panel) and 5, 7, 10 years (lower panel); percentage points. Option-implied distributions are extracted using the LAD estimator with unimodality restriction. Daily quotes of inflation caps and floors from October 2009 to February 2015 are taken from Bloomberg.

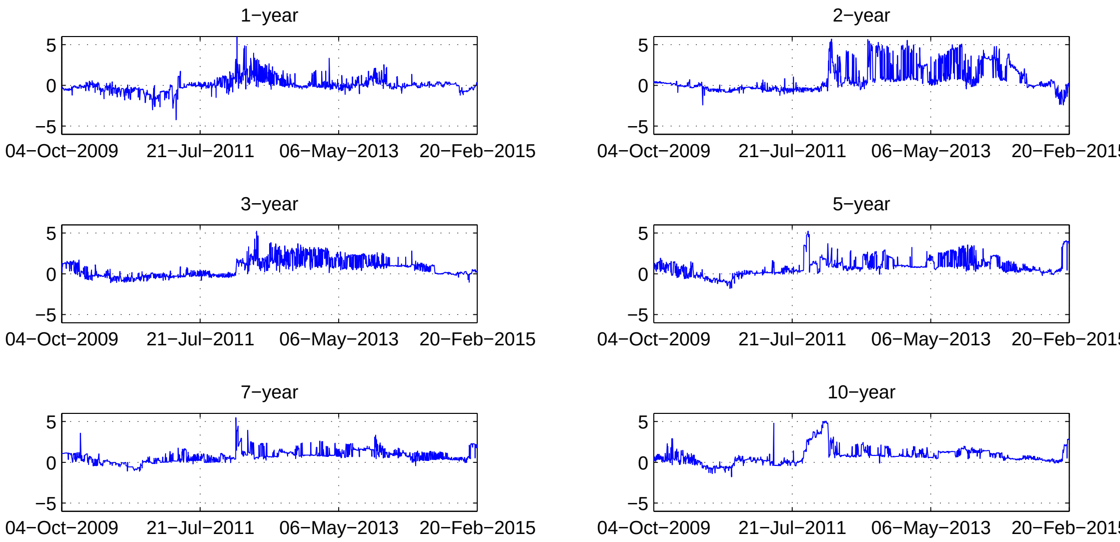

Figure 5. Skewnesses of option-implied risk-neutral inflation distributions at different maturities (1-year, 2-year, 3-year, 5-year, 7-year, 10-year); percentage points. Option-implied distributions are extracted using the LAD estimator with unimodality restriction. Daily quotes of inflation caps and floors from October 2009 to February 2015 are taken from Bloomberg.

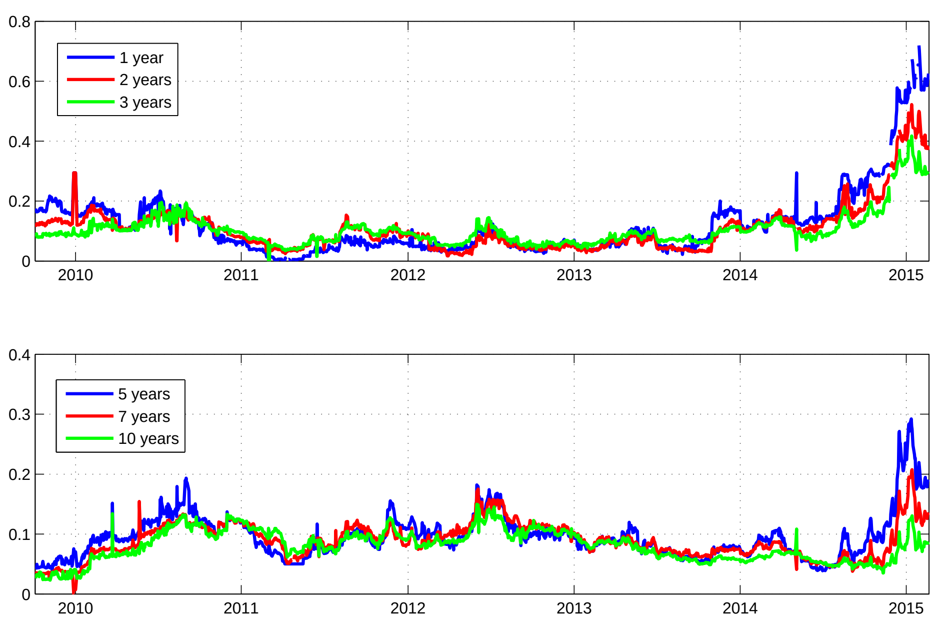

Figure 6. Risk-neutral probability that the average annual inflation rate at different maturities (1, 2, 3, 5, 7 and 10 years) is negative. Option-implied distributions are extracted using the LAD estimator with unimodality restriction. Daily quotes of inflation caps and floors from October 2009 to February 2015 are taken from Bloomberg.

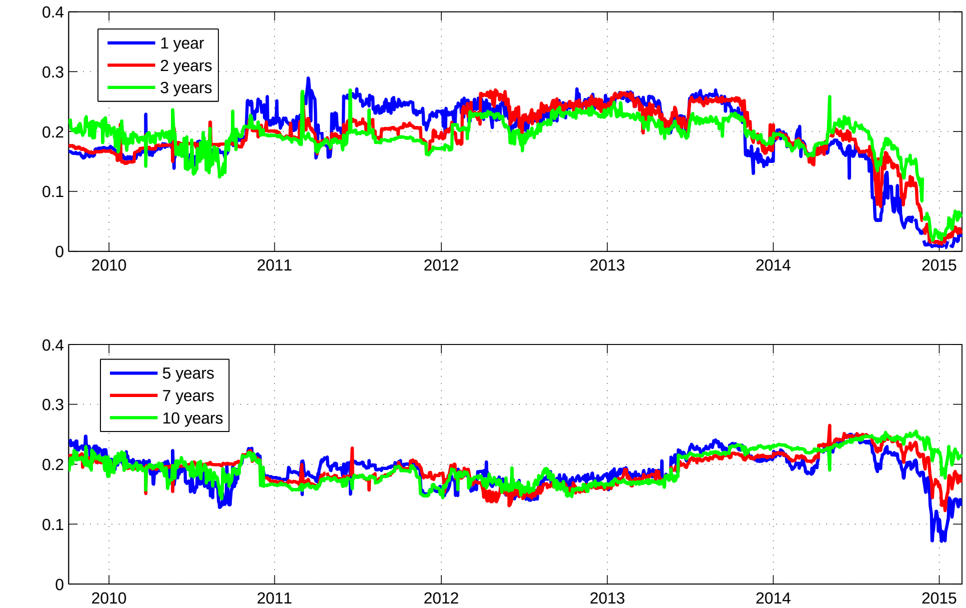

Figure 7. Risk-neutral probability that the average annual inflation rate at different maturities (1, 2, 3, 5, 7 and 10 years) falls between 1.5 and 2%. Option-implied distributions are extracted using the LAD estimator with unimodality restriction. Daily quotes of inflation caps and floors from October 2009 to February 2015 are taken from Bloomberg.

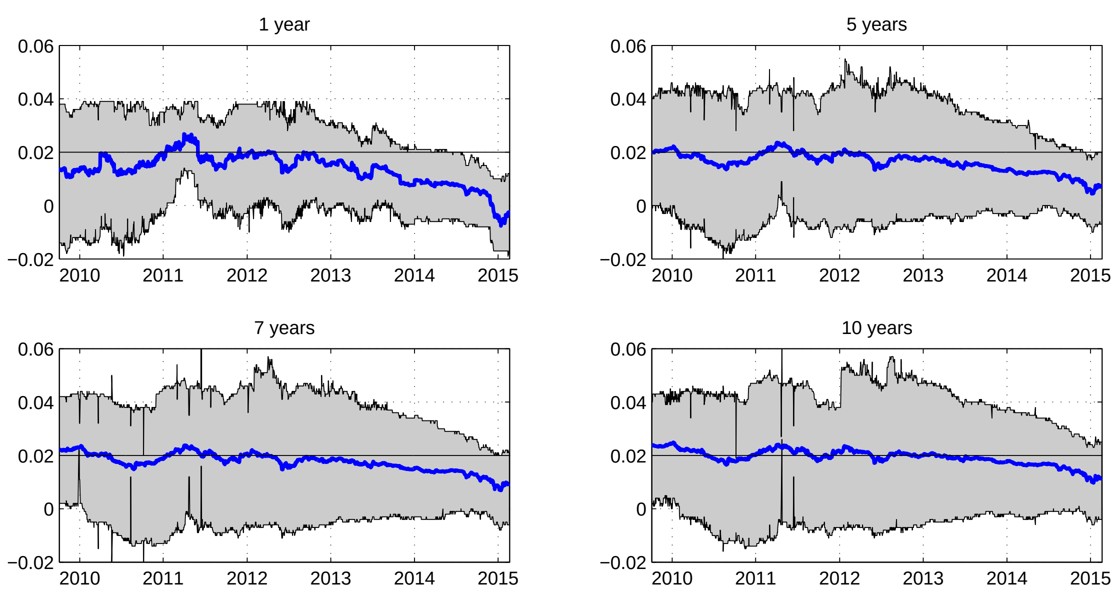

Figure 8. Confidence bands for the mean of the option-implied probability distributions of future inflation, at the 10% level and at maturities 1, 5, 7 and 10 years.

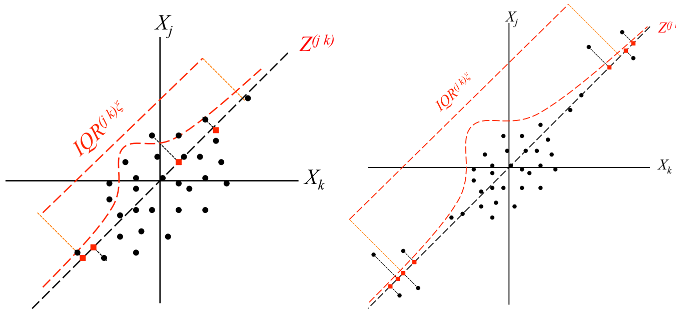

Figure 9. Source: Ricci and Veredas (2013), Figure 1 at page 34. Diagrammatic representation of TailCor. Scatter plots, along with the 45-degree line, where X_j and X_k are positively related (the pairs are depicted with circles). Left plot shows a linear relation while right plot shows a nonlinear relation. Projecting the observations onto the 45-degree line produces the random variable $Z^{(jk)}$, depicted with squares. For illustrative purposes the projection is shown only for the observations on the tails but in the estimation it is done for all the observations.

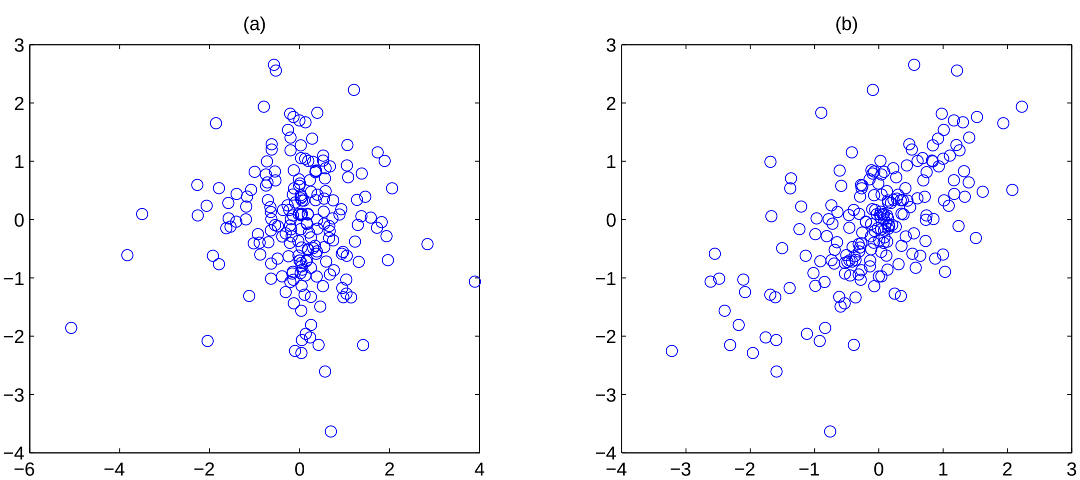

Figure 10. Scatterplots of the filtered innovations of the 1y1y and 5y5y forward rates in different time intervals: the left panel refers to the period May 2011 - Feb 2012 and shows little tail dependence, whereas the right panel refers to the period May 2014 - Feb 2015 and shows clear tail dependence in both the upper and the lower tail.

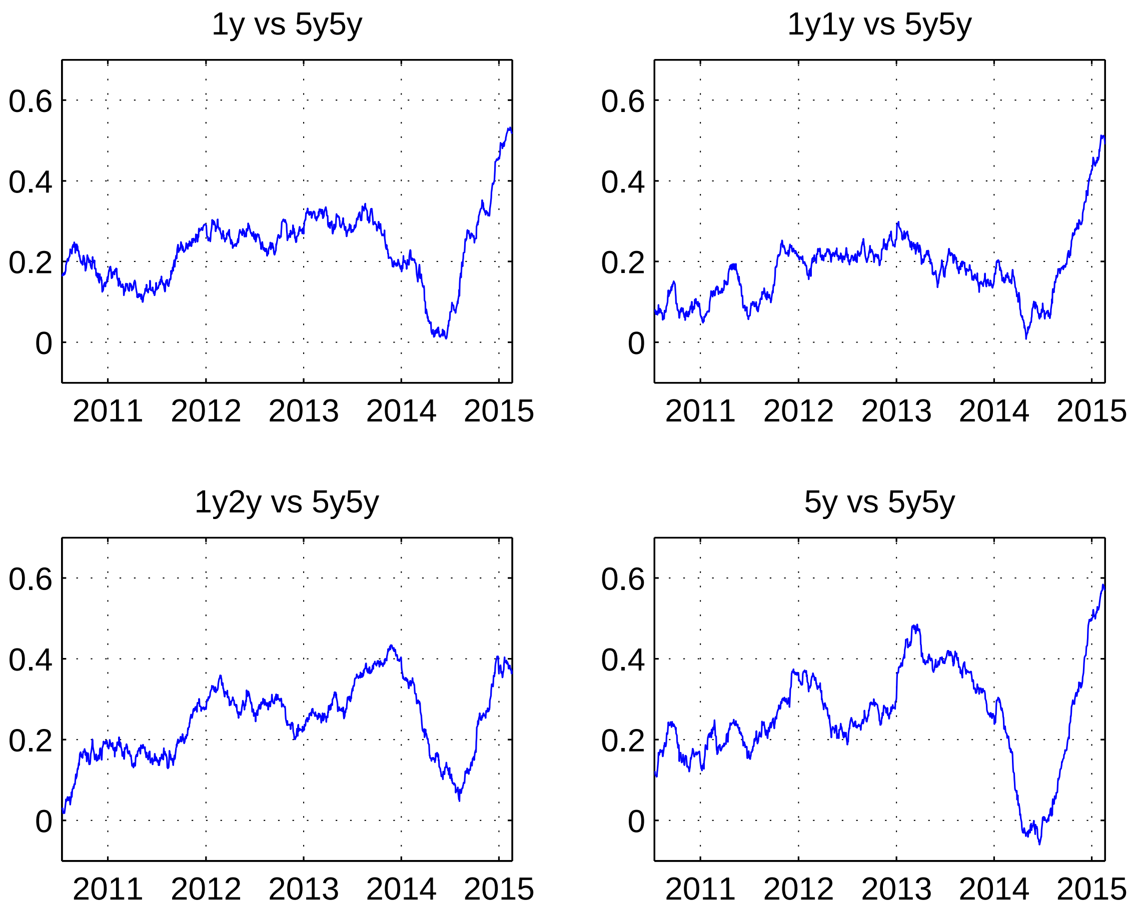

Figure 11. Pearson ρ correlation coefficient on short vs. medium-to-long term market-based inflation expectations. Short-term mean expectations are 1y ahead, 1y ahead after 1 year (1y1y forward inflation swap), 1 year ahead after 2 years (1y2y forward) and 5 years ahead, while medium-to-long term expectations are 5 years ahead after 5 years (5y5y forward). The coefficient is computed using 200 business days rolling windows. Sample: 5-Oct-2009 to 19-Feb-2015.

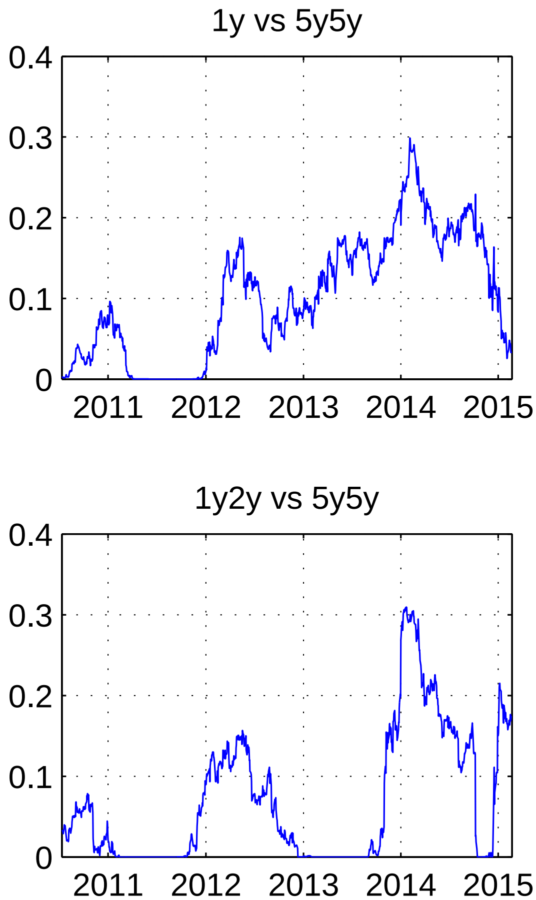

Figure 12. Index of tail-comovement using the Student's t copula on short- vs. medium-to-long term mean inflation expectations. The index ranges from 0 (no tail dependence) to 1. This index indicates the average comovement on both upper and lower tails. Short-term mean expectations are 1y ahead, 1y ahead after 1 year (1y1y forward inflation swap), 1 year ahead after 2 years (1y2y forward) and 5 years ahead, while medium-to-long term expectations are 5 years ahead after 5 years (5y5y forward). Values are computed using 200 business days rolling windows. Sample: 5-Oct-2009 to 19-Feb-2015.

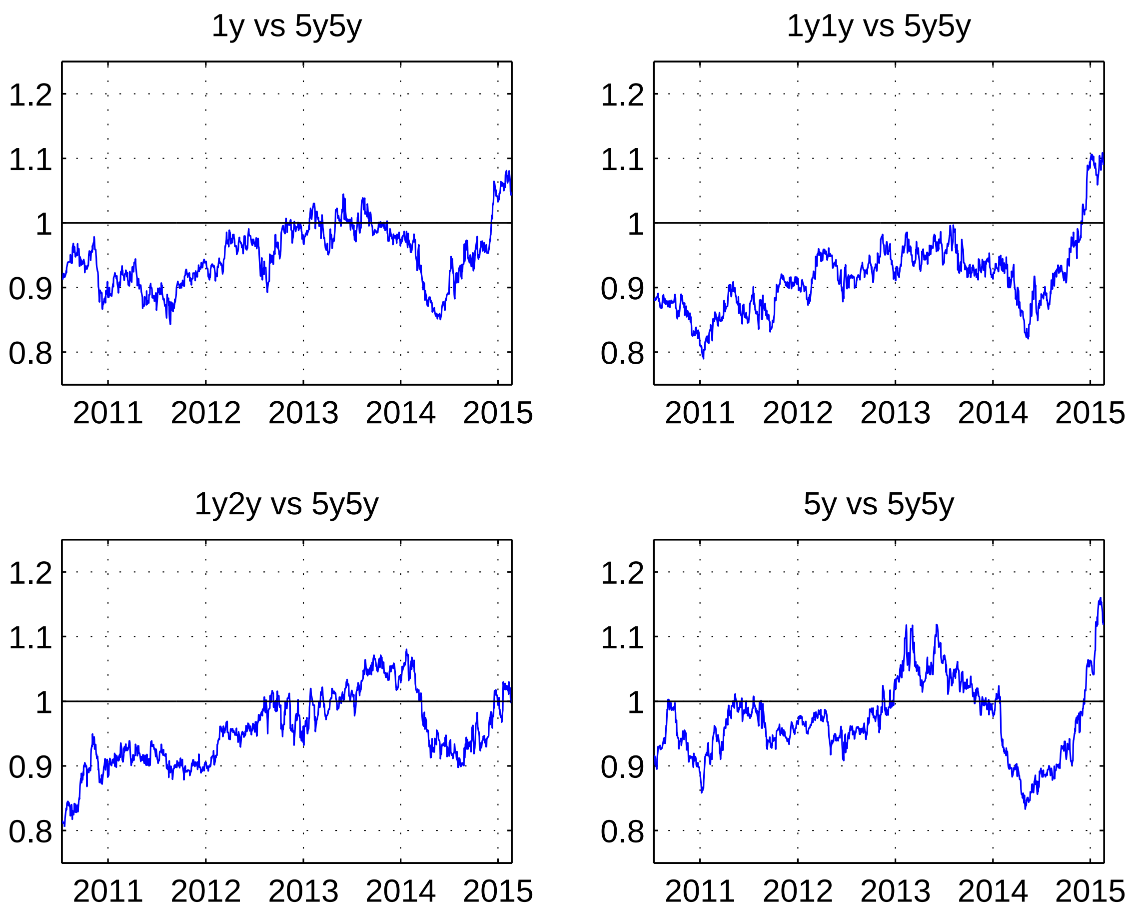

Figure 13. TailCor index computed on short- vs. medium-to-long term mean inflation expectations. It takes values between 0 and $+\infty$; under Gaussianity and uncorrelation, the index takes the value 1. This measure indicates the average comovement in both upper and lower tails. Short-term mean expectations are 1y ahead, 1y ahead after 1 year (1y1y forward inflation swap), 1 year ahead after 2 years (1y2y forward) and 5 years ahead, while medium-to-long term expectations are 5 years ahead after 5 years (5y5y forward). Values are computed using 200 business days rolling windows. Sample: 5-Oct-2009 to 19-Feb-2015.

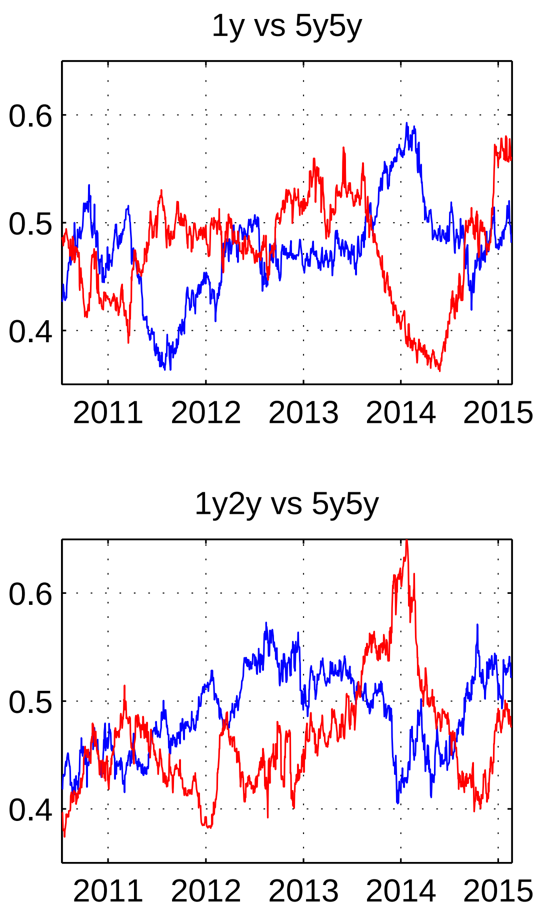

Figure 14. UpTailCor (blue line) and DownTailCor (red line) computed between short and medium-to-long term mean expectations. Short-term mean expectations are 1y ahead, 1y ahead after 1 year (1y1y forward inflation swap), 1 year ahead after 2 years (1y2y forward) and 5 years ahead, while medium-to-long term expectations are 5 years ahead after 5 years (5y5y forward). Values are computed using 200 business days rolling windows; $\xi = 0.85$, $\tau = 0.75$. Sample: 5-Oct-2009 to 19-Feb-2015.

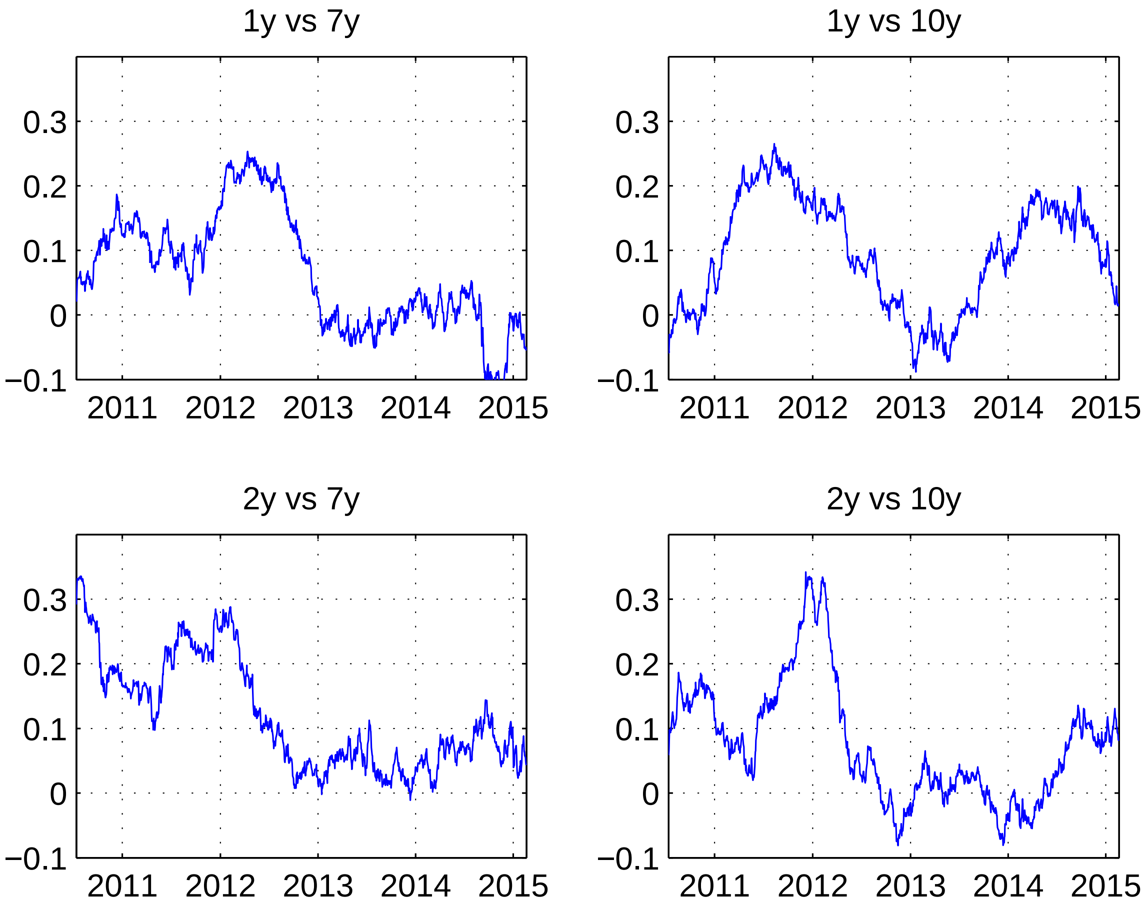

Figure 15. Pearson ρ correlation of short- vs. long-term standard deviations of option-implied distributions of future inflation. Short-term standard deviations are 1 or 2-year ahead, while long-term ones are 7 or 10 years ahead. The coefficient is computed using 200 business days rolling windows. Sample: 5-Oct-2009 to 19-Feb-2015.

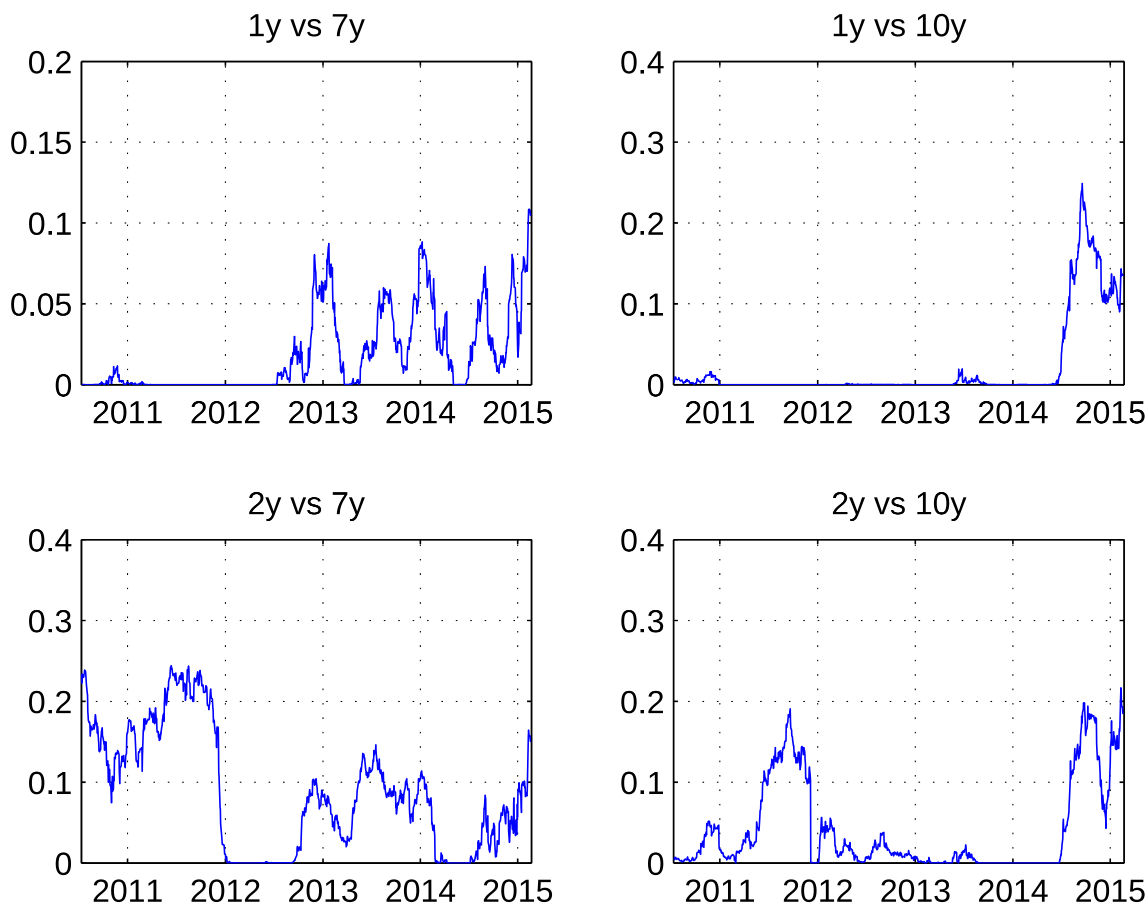

Figure 16. Index of tail-comovement using the Student's t copula on standard deviations of short vs. long-term option-implied distributions. The index ranges from 0 (no tail dependence) to 1. This index indicates the average comovement on both upper and lower tails. Short-term standard deviations are 1 or 2-year head, while long-term ones are 7 or 10 years ahead. Sample: 5-Oct-2009 to 19-Feb-2015.

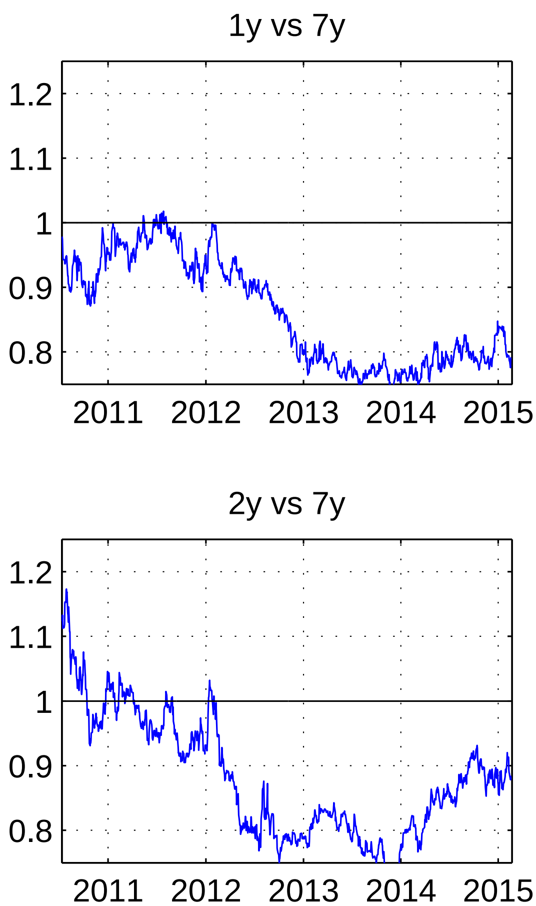

Figure 17. TailCor index computed on standard deviations of short vs. long-term option-implied distributions. It takes values between 0 and $+\infty$; under Gaussianity and uncorrelation, the index takes the value 1. This measure indicates the average comovement in both upper and lower tails. Short-term standard deviations are 1 or 2-year head, while long-term ones are 7 or 10 years ahead. Values are computed using 200 business days rolling windows. Sample: 5-Oct-2009 to 19-Feb-2015.

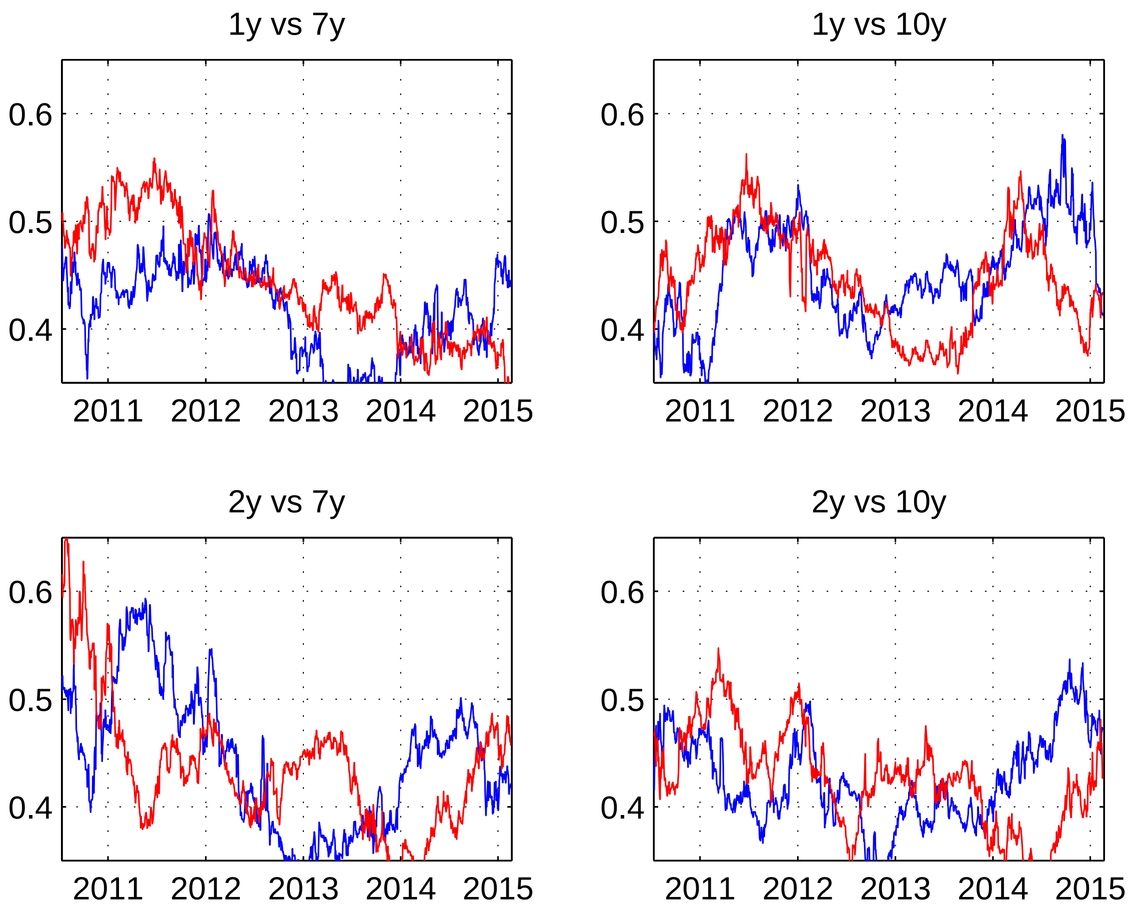

Figure 18. UpTailCor (blue line) and DownTailCor (red line) computed on standard deviations of short vs. long-term option-implied distributions. Short-term standard deviations are 1 or 2-year head, while long-term ones are 7 or 10 years ahead. Values are computed using 200 business days rolling windows; $\xi = 0.85$, $\tau = 0.75$. Sample: 5-Oct-2009 to 19-Feb-2015.

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