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THE EVOLUTION OF THE ANCHORING OF INFLATION EXPECTATIONS

by Ines Buono* and Sara Formai*

Abstract

We investigate the degree of anchoring in inflation expectations for different advanced economies using data from professional forecasters' surveys. We define expectations as anchored when movements in short-run expectations do not trigger movements in expectations at longer horizons. Using time-varying parameter regressions, we provide evidence that anchoring has varied noticeably across economies and over time. In particular, we find that starting from the second half of 2008, inflation expectations in the euro area, unlike in the US and in the UK, have shown signs of a de-anchoring.

JEL Classification: E31, E52, D84, C14.
Keywords: anchoring, inflation expectations, nonparametric estimation.

Contents

1. Introduction 5
2. Data 9
3. Empirical strategy and results 11
   3.1 Time-invariant coefficients 11
   3.2 Time-varying coefficients 12
4. Robustness 16
5. Concluding remarks 17
Tables and figures 19
Appendix 24
References 26

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1 Introduction

Anchoring inflation expectations is a cornerstone of central banks' monetary policy strategy. Once incorporated into agents' wage- and price-setting decisions, expectations drive actual inflation and, if disconnected from the central banks' inflation target, may act as an impediment to the correct transmission of monetary policy.

In the last thirty years, the literature on inflation expectations has focused mainly on the effects of central banks' communication strategies, the general consensus being that inflation targeting regimes help expectations to remain well anchored (see Levin, Natalucci and Piger (2004), among others). However, recent evidence from several advanced economies has shifted the debate towards the risks entailed by actual inflation rates being persistently below the central banks' target. The resulting low inflation expectations may lead to destabilizing dynamics and self-fulfilling liquidity traps. This is a concern especially when monetary policy is constrained by the zero lower bound (ZLB): lower inflation expectations result in higher real interest rates, causing a monetary tightening that central banks are unable to counter by lowering nominal rates. If a low-inflation scenario gets entrenched in medium- and long-run expectations, then central banks may find it hard to bring inflation back to target.

At the current juncture, this concern is especially felt in the euro area, where inflation started falling in late 2012, dropping into negative territory at the end of 2014, and is still struggling to rebound. In part as a result of the recent decline in oil prices, inflation has fallen considerably, and has been hovering not far above zero even when using the definition that excludes volatile components. As ECB President Mario Draghi explained in a recent speech, "*[...] a prolonged period of low or even negative inflation rates might destabilize inflation expectations. And we know from international experience this change can happen quite quickly, especially if the objective of monetary policy is not clear. Thus, we have to judge carefully how an apparently temporary shock is spreading through the economy and affecting expectations.*"[1]

What is happening in the euro area is reminiscent of Japan's experience during its 'lost decade'. Starting from the early 1990s, Japan experienced a sharp fall in inflation, which slid into negative territory in 1995 and remained there, with few exceptions, for almost 20 years.

[1]Speech by Mario Draghi, President of the ECB, at the ECB Forum on Central Banking, Sintra, 26 May 2014.

One possible explanation for these events hinges on the de-anchoring of inflation expectations, which may explain why the Japanese authorities, stuck in a liquidity trap, were unsuccessful in their battle against deflation (see Piazza (2015)).

The aim of this note is to analyze the anchoring of inflation expectations for the four largest advanced economies from 1990 to March 2015. Using *Consensus Economics* surveys of professional forecasters, we study the inflation expectation pass-through, defined as the link between short- and long-term inflation expectations, given respectively by one-year-ahead and five-years-ahead expectations. If expectations are well anchored, then changes in short-run expectations should not be correlated with changes at longer horizons.[2] Our contribution to the literature is two-fold: first, we use the same data source for all the economies in the analysis, providing comparable results across countries. Second, we investigate whether the intensity of the inflation pass-through has changed over time in different economies. A common technique to assess the constancy of the parameters in a model is to compute estimates over a rolling window of fixed size through the sample. If the parameters are constant over the entire sample, then the estimates over the rolling window should not be too different; if the parameters change at some point during the sample, then the rolling estimates should capture this instability. Moving averages typically use equal weights for the observations in the subsamples used in each regression. As an alternative, we also consider estimations based on both exponential and Gaussian kernel functions that assign decreasing weights for observations more distant in time.

Acknowledging the relevance of inflation expectations for monetary policy implies abandoning the assumption of rational expectations and is theoretically justified by models that assume imperfect knowledge and adaptive learning (as in Orphanides and Williams (2004) and Orphanides and Williams (2007)), where agents recursively update their beliefs on the underlying inflation process. As Beechey, Johannsen and Levin (2011) suggest, inflation expectations are firmly anchored if long-run expectations are stable over time, exhibit little cross-sectional dispersion, and are insensitive to macroeconomic news.

Our working hypothesis is that any macroeconomic news affecting wages and prices is incorporated into short-run inflation expectations. For instance, a negative oil price shock should affect short-run inflation expectations since lower oil prices trigger a decrease in inflation that

[2] A limitation of this approach is that it does not allow to establish whether, if anchored, expectations are indeed anchored at the central bank's target or, more generally, whether they are anchored at a constant level.

persists for some time. In other words, the change in short-run inflation expectations is a proxy for macroeconomic news relevant for inflation. Since data on short- and long-run expectations are provided in the same release of the *Consensus* survey, they can be assumed to be both conditional on the same information set. This means that macro news are taken into account and incorporated into short-run expectations when long-run ones are disclosed. Thus, by regressing long-run (five-years-ahead) on short-run (one-year-ahead) expectations, we are taking into account all macroeconomic surprises that may be relevant to the forecast formation process. If monetary policy is credible and expectations are well-anchored, long-run expectations should be insensitive to these shocks and exhibit no co-movement with short-run expectations.

The notion of anchoring is conceptually linked to that of inflation persistence, defined as the tendency of inflation to converge slowly towards its long-run value following a shock.[3] Persistence was a key topic in the literature ten years ago, while now the focus has shifted to anchoring, although it is hard to clearly disentangle the two concepts, both theoretically and empirically. Since the anchoring of expectations may affect the properties of inflation dynamics, it can be interpreted as one of the factors causing persistence. This, however, may also depend on other factors, such as the degree of persistence of the marginal costs and the output gap, as well as on the intrinsic dependence of inflation on its own past values.

The literature has proposed several approaches to analyze the extent of anchoring, depending on how the phenomenon is defined and on the data used. Dovern, Fritsche and Slacalek (2012) rely on *Consensus* surveys to assess the disagreement among forecasters, assuming that if expectations are well-anchored, then their mean across forecasters should be stable around a given level, and the cross-sectional dispersion should be small. They find that having an independent central bank improves anchoring. An alternative definition, related to the one we use, implies that expectations are anchored if - at longer horizons - they are not affected by macroeconomic news. Levin et al. (2004), using high-frequency financial-market-based measures of inflation expectations, found that long-run expectations are less sensitive to macro news than short-run ones, and that they are more firmly anchored in inflation-targeting countries. A similar result is found by Beechey et al. (2011), who provide evidence that in the period 2003-

[3] This is the definition proposed by the Inflation Persistence Network, a research team consisting of economists from the European Central Bank and the national central banks of the Eurosystem (see Altissimo, Ehrmann and Smets (2006)).

2007 inflation expectations were more anchored in the euro area than in the United States. In a recent contribution, Miccoli and Neri (2015) used monthly market-based data for the euro area to measure expectations at different horizons and found that inflation surprises, defined as the difference between realized inflation and the median of analysts' expectations immediately before the release of the data, have a significant impact on inflation expectations even at the medium-term horizon.

Few papers focus on the definition of anchoring as inflation pass-through, i.e. the link between short-term and long-term expectations. By using micro panel data for US consumer inflation expectations from the Michigan survey, Drager and Lamla (2013) find that long-run inflation expectations have become more anchored over the last two decades, identifying as a turning point the preemptive tightening adopted by former Fed Chairman Alan Greenspan after 1996. Gefang, Koop and Potter (2008) used US and UK financial-market-based data from 2003 to 2008 and found, for both countries, evidence of a 'contained' pass-through of inflation expectations, meaning that medium-term expectations do respond to movements in actual inflation, but remain within a narrow range around the central bank's policy target.[4] More recently, Cecchetti, Natoli and Sigalotti (2015) have applied different methodologies to detect tail co-movements in financial-market-based inflation expectations at different horizons in the euro area, from the end of 2009 to February 2015. They find that, since mid-2014, negative tail events impacting short-term inflation expectations have spurred downward revisions also in long-term ones.[5]

The empirical results available in the literature are not fully comparable across economies and time, as the different analyses do not use a common definition of anchoring. Our work is, to our knowledge, the first to tackle this issue for the US, Japan, the UK and the euro area by exploiting data collected consistently by the same source, the *Consensus* survey, and estimating parameters that are allowed to vary over time. The remainder of this note is organized as follows. In Section 2 we describe our data set. Section 3 explains the empirical strategy and presents

[4]Gefang et al. (2008) define inflation expectations as being 'anchored' when the pass-through is small and constant over time, 'contained' when it is smaller than one and varying over time, and 'unmoored' when it is close or equal to one. See also Jochmann, Koop and Potter (2010).

[5]Recent internal analyses conducted at the BIS and the ECB have contributed to the debate. In BIS (2015) the sensitivity of long-term inflation expectations to actual inflation is explored using panel data for advanced economies. On average, anchoring appears to have weakened in 2010-2014 compared with 2000-2007. Using market-based data and time-varying parameter models, ECB (2015a) and ECB (2015b) found emerging risks of a de-anchoring in the Eurozone, but not in the US.

the main results. Section 4 deals with some robustness checks. Section 5 concludes.

2 Data

Our empirical analysis is based on semiannual data on expected inflation, taken from the *Consensus Economics* surveys released in April and October of each year.[6] Professionals are asked to provide their inflation forecasts for the current year and for each of the following five years, as well as the forecasts on average inflation between six and ten years ahead. While the data for the US, Japan and the UK cover the period from October 1989 to April 2015, the data for the euro area are collected starting from 1999. We have thus completed the series by aggregating the data for the four main euro area economies (Germany, France, Italy and Spain) - available from October 1989 - using weights based on GDP shares. As a result, our dataset contains around 50 semiannual observations for each economy.

We make two main adjustments to the *Consensus* raw data. First, semiannual surveys provide fixed-event forecasts, meaning that each expectation refers to a specific calendar year, regardless of the time in which the survey is collected (either April or October). This is a problem for the analysis since observations differ in the length of the forecasting horizon and, as a consequence, in the information set available to forecasters: expectations for a given year, when formulated in April, need a forecast of inflation for at least the subsequent 9 months; when formulated in October, instead, they require a forecast for only 3 months.[7] Following a common approach in the literature (see, for instance, Dovern et al. (2012)), we approximate the fixed-horizon forecast for the next twelve months as an average of the fixed-event forecast for the current and next calendar year weighted by their share in the forecasting horizon. The 12 month-ahead forecast for month m and year y is given by

$$\pi^e_{my,12} = \frac{12 - m + 1}{12}\pi^e_{my,y} + \frac{m-1}{12}\pi^e_{my,y+1}$$

where $\pi^e_{my,y}$ is the fixed-event forecasts for year y released in month m of year y. For instance, for April 2009 this is the average between the forecasts for 2009 and 2010, with a weight equal

[6] Short-term forecasts are available monthly, while longer-term forecasts are released only twice a year.

[7] For some economies the March and September CPI official data are released after the *Consensus* survey, for others before.

to 9/12 for the former and 3/12 for the latter, as 3 and 9 are the number of months respectively in 2009 and in 2010 that enter the forecasting horizon of 12 months starting in April 2009.

Our second adjustment to the Consensus data consists of 'cleaning' the forecasts for the UK and Japan of changes in VAT rates; we do this by considering both announcements and actual implementation.[8] Without such a correction, in fact, expectations at different horizons would not be comparable. Following the Bank of Japans approach,[9] as well as other studies for different countries (see, for instance, Carare and Danninger (2008), for Germany), we consider a pass-through of two thirds from an increase in the VAT to inflation, and adjust expectations accordingly.[10]

While semiannual *Consensus Economics* surveys have the advantage of providing homogeneous data for several advanced economies, a major drawback is the small number of observations compared, for example, with high-frequency data extracted from asset prices. However, the use of market-based data is not free from flaws. They are significantly affected by risk premia and market-specific liquidity compensation caused by inflation expectations and neither theory nor the available empirical evidence offer clear indications on how to measure and therefore adjust for these two components (see Miccoli, Natoli, Secchi, Sigalotti and Taboga (2015)).[11] This concern has become particularly relevant in the last few years, since the premia tend to be more volatile in periods of market distress.[12]

[8]For the euro area and the US, VAT regimes are at the state level, so any change should be weighted to assess the effects on inflation measured at the aggregate/national level.

[9]See the 'Outlook for Economic Activity and Prices' published periodically by the Bank of Japan (BoJ). In the table with the inflation forecasts, the BoJ reports both the forecast for the CPI and that for the CPI excluding the effects of consumption tax hikes.

[10]See Table A1 in the Appendix for summary statistics.

[11]The information content of market-based data has been recently questioned by some members of the ECB Governing Council. In particular, it has been noted that the recent drop in euro-area inflation swap rates could reflect, in addition to changes in inflation expectations, also changes in inflation risk premia or other premia related to the illiquidity of inflation swaps.

[12]Fed Chairman Janet Yellen also provided some cautionary remarks on the use of market-based measures during the March 2015 press conference, pointing out that "*[...] they are also informative but can move around for reasons pertaining to liquidity in the treasury market and in the tips market and also because of changing perceptions of inflation risk.*"

3 Empirical strategy and results

3.1 Time-invariant coefficients

We analyze the co-movement between short- and long-run inflation expectations of professional forecasters: ideally, if inflation expectations are firmly anchored, a transitory shock should affect only short-run expectations.

We start by estimating the following time-invariant panel regression:

$$\Delta\pi^e_{t,h_l} = \alpha + \beta_1 \Delta\pi^e_{t,h_s} + \beta_c \Delta\pi^e_{t,h_s} * d_c + d_c + \epsilon_t \tag{1}$$

where π^e_{t,h_l} is the long-run forecast, i.e. the *Consensus* CPI inflation expectations from the survey at time t for the long-run forecast horizon h_l (where $l = 5, 4, 3, 2$ years ahead), π^e_{t,h_s} is the corresponding short-run forecast, namely the one-year-ahead forecast.[13] Expectations are taken in first differences as we want to capture the shocks that occurred between two consecutive surveys. In order to compare the estimates for different economies, we insert the interaction between the main dependent variable and a set of country dummies (d_c where c stands for United States, United Kingdom or euro area). Table 1 reports the descriptive statistics of the dependent and independent variables. The coefficients for the US, the UK and the euro area should be interpreted as deviations from those relative to Japan, chosen as the benchmark country.

The time-invariant inflation expectation pass-through is quite different across economies, at all horizons. For Japan the co-movement is always positive and statistically significant. A 1% increase in one-year-ahead inflation expectations is associated with an increase of 0.49% for two-years-ahead inflation expectations and of 0.27% for five-years-ahead expectations. Not surprisingly, the co-movement decreases with the forecasting horizon, since the persistence of shocks that affect inflation and its expectations tends to wane at longer horizons. According to the last row in Table 1, the regression shows a strong similarity between Japan and the euro area: their estimated pass-through coefficients are not statistically different at any horizon.

[13] We do not use the average expected inflation between six and ten years ahead as a measure of long-run inflation expectations since this forecast tends to be constant over time for all economies and it always approximately coincides with the central banks' monetary policy objective of each economy - a finding that is strikingly at odd with other measures available for the same forecast horizon, such as those derived from forward interest rates.

Conversely, for the US and the UK the estimated pass-through is statistically different from that of Japan at different horizons. In particular in Table 2 we present the Wald test for the hypothesis that the pass-through coefficients for each economy are equal to zero, i.e. H_0 : $\beta_1 + \beta_c = 0$. As expected we reject the null hypothesis at all horizons for the euro area, while for the US and the UK the coefficients are null for $h_l \geq 3$ and $h_l \geq 4$, respectively.

All the above tests are based on standard errors corrected for heteroskedasticity. Moreover residuals do not exhibit any autocorrelation, as shown by the Durbin-Watson tests reported in Table 3.[14]

The above estimates show the heterogeneity of the inflation pass-through across economies. However, coefficients may average out important variations over time. For instance, concerns of a possible de-anchoring of inflation expectations in the euro area are a recent issue, suggesting that the pass-through coefficient may have changed in the last part of the sample. In the next section we thus relax the hypothesis that coefficients are constant.

3.2 Time-varying coefficients

For each economy, we estimate the following regression:

$$\Delta\pi^e_{t,h_l} = \alpha_t + \beta_t \Delta\pi^e_{t,h_s} + \epsilon_t \tag{2}$$

where π^e_{t,h_l} is the 5-years-ahead inflation expectations, π^e_{t,h_s} is the 1-year-ahead inflation expectations, and, differently from regression (1), we allow the pass-through coefficient β_t to vary over time.

To this end, we first employ a rolling window methodology in which, given our sample of semiannual observations from $1990h1$ to $2015h1$, regression (2) is estimated with OLS for all the overlapping windows of n observations $[t - n + 1, t]$, with $t = 1990h1 + n - 1, ..., 2015h1$. This provides a sequence of estimated parameters $\{\beta_{1992h2}, ..., \beta_{2015h1}\}$. Figure 1 shows, for each of the four economies, the estimated parameters and the 95% confidence interval when $n = 20$.[15] The values reported in the chart for each date refer to the coefficient and confidence

[14]A Durbin alternative test that does not require strict exogeneity of the regressors and a Breusch-Godfrey test for higher order autocorrelation provide the same results.

[15]We excluded $2009h1$ from all the regressions, since the negative variation in the one-year-ahead expectations is much higher in absolute terms than the average (-1.9 for euro area; -2.4 for Japan; -0.4 for the UK and -3.0

intervals calculated over the 20-year rolling window ending on that date; the rolling windows are, therefore, backward-looking.

As Figure 1 shows, in the euro area the co-movement between short- and long-run expectations is positive at the beginning of the sample, then starts to decrease in the mid-2000s and approaches zero after 2006. However, the coefficient increases again starting in 2010 and is positive and statistically significant from $2013h2$, providing some evidence in favour of a recent de-anchoring of inflation expectations. The picture appears different for the US, where the co-movement, never statistically different from zero, decreases from 2007 and remains fairly constant in the last part of the sample. The pass-through coefficient, conversely, looks highly variable and in positive territory for Japan, whose graph is reported with a different y-scale. Although the magnitude of the coefficient is always high compared to that of the other economies, estimates are very imprecise, with wide confidence intervals that lie above zero only between $2007h2$ and $2011h1$.[16] Unfortunately this prevents us from drawing clear-cut conclusions. Finally, in the UK, although the co-movement shows an increase between 2004 and 2008, there is no evidence of a de-anchoring throughout the entire period.

Of course, rolling window results depend on the window length n. In particular, as n increases the graph becomes smoother. Nevertheless, the main trends and conclusions do not change when we vary n (see Graph A1 in the appendix where we compare results for n=10, 20 and 26).

Technically, the rolling window estimation applies equal weights to all the past n observations $[t-n+1; t]$ and a weight equal to zero to all the remaining ones. Alternatively, one could apply a different weighting scheme: for instance, discounting the more distant past observations would produce a smoother path for β_t that could better capture gradual structural changes in the underlying relationship. To this end we implement an exponential weighting scheme within each window of $n = 20$, such that

for the US).

[16] As it is evident from the rolling window with n=10 (see Fig. A1 in the Appendix), the estimated coefficients for Japan exhibit a high variability which is partially due to the high variance of $\Delta\pi^e_{t,5y}$ (see Table A1 in the Appendix). In particular, three observations in the first part of the sample ($1997h1$, $1997h2$ and $1998h1$), marked by high negative correlation, are responsible not only for the low coefficients in the late 1990s, but also, being outliers, for their wide confidence intervals. Given the small size of the sample, a limited number of observations can strongly affect the results. This represents an issue only in the case of Japan. With high variability the median of the forecasters expectations would be more appropriate than the mean, but, unfortunately, it is not available in the semiannual *Consensus* releases.

$$\hat{\beta}_t = \left[\sum_{j=t-n}^{t} \omega_{j,t} x_j x_j' \right]^{-1} \left[\sum_{j=t-n}^{t} \omega_{j,t} x_j y_j \right], \tag{3}$$

where $\omega_{j,t}$ is the weight and x_j and y_j stand for generic observations of the dependent and the independent variable. In particular

$$\omega_{j,t} = cK \left(\frac{t-j}{H} \right), \tag{4}$$

where $K(z) = e^{(-z)}$ is the Exponential Kernel function normalized by the bandwidth H and c is an integration constant such that the weights within each window sum up to 1. Changes to H change the weighting scheme and the estimated parameters: the higher is H, the more uniform are the weights (when $H \to \infty$ the estimation becomes an unweighted OLS and corresponds to the rolling windows specification described above). This is shown in Figure 2, which reports the weights used in estimating β_t for $t = 1999h2$, $2004h2$, $2010h1$ and $2015h1$ when using different values of H. Following Giraitis, Kapetanios and Yates (2014), we set $H = \sqrt{T}$.

Figure 3 shows the results, which, as in the rolling windows analysis, start from $1999h2$. While the main insights of the analysis are confirmed for all the economies, some differences emerge for the euro area at the extremes of the sample. Exponential estimations provide positive and highly significant coefficients already starting from $2012h2$ while rolling windows results are less precise (wider confidence interval) with significance only from $2013h2$ (see Figure 1). This is consistent with the idea that the euro area is experiencing a movement toward de-anchoring only in the last part of the sample. This is better captured by the exponential weighting which discounts past observations when, presumably, the de-anchoring tendency was still not in place. Another difference emerges in the estimations up to 2005, when the rolling window captures a higher correlation between inflation expectations compared with the exponential scheme. Since in this case the rolling window gives greater weight to pre-euro observations (starting from 1990) compared with the exponential weight technique, this result suggests that the co-movement was higher in the pre-euro period and started to move toward zero around and after the introduction of the common currency. As both these estimation methods are backward-looking, we cannot directly test this hypothesis.

The discussion above points out that a constant weighting scheme may fail to capture in a timely manner an ongoing change in the underlying relationship, especially when dealing with low-frequency data. A further improvement in this direction may be to obtain each coefficient β_t using the observations in a neighbourhood of t, both in the near past and in the near future, with a scheme that assigns weights that symmetrically decrease with distance. Of course this is not feasible at the extremes of the sample, where the exponential scheme is the only viable option. We thus implement a Gaussian kernel-based nonparametric estimator whose properties are described in detail by Giraitis et al. (2014).[17] As we did before, the estimator is a weighted OLS, but the weights are now given by $\omega_{j,t} = cK[(t-j)/H]$, with $K = (1/\sqrt{2\pi})e^{\frac{-z^2}{2}}$, a Gaussian Kernel function normalized by the bandwidth H and where c is an integration constant. As Figure 4 shows, the Gaussian kernel function implies a scheme that is forward-looking at the beginning of the sample, backward-looking at the end and centered for the other observations. As for the exponential, the higher is H, the more uniform are the weights (when $H \to \infty$ the estimation becomes an unweighted OLS).[18] Finally the Gaussian kernel allows us to use the full sample of size T to estimate each β_t, which is particularly important given our small sample size. The estimator is thus given by $\hat{\beta}_t = [\sum_{j=1}^{T} \omega_{j,t} x_j x_j']^{-1} [\sum_{j=1}^{T} \omega_{j,t} x_j y_j]$.

Figure 5 shows the results, now starting from 1990. In interpreting the results, it is important to keep in mind that, except at the extremes of the sample, the rolling windows are now centered on the time reported on the horizontal axis, rather than backward-looking as in the previous exercises. As expected, the evolution over time of β_t is smoother for all economies and previous findings are generally confirmed. For the US we now observe a clear monotonically decreasing trend starting from a relatively high level of co-movement (0.2), although robust standard errors imply wide confidence intervals in the early part of the sample. For the UK and Japan the estimated pass-through is more variable over time and, especially for Japan, less precise. For the UK, it is never significantly different from zero, thus indicating firmly anchored expectations. For Japan, instead, between 1990 and 1992 the co-movement stands at very high levels, suggesting far from anchored expectations. Although estimates are not statistically different from zero from 1992 onwards, they increase again from 1999 to 2006 and then show a slow decreasing pattern until the very last part of the sample.[19]

[17] This estimator was recently used by Riggi and Venditti (2015) to estimate the time-varying parameters of a (backward-looking) Phillips curve.

[18] Also for this estimation we set $H = \sqrt{T}$.

[19] As already mentioned, results for Japan are more sensitive to a limited number of observations that behave

In the euro area the co-movement of expectations is significantly different from zero between 1990 and 2000, a result that could not emerge with the backward-looking sample used in the previous estimations. The co-movement then decreases starting from 1996, and becomes null in 2001, highlighting the key role played by the adoption of the common currency and the credibility of the ECB in keeping inflation expectations well anchored. The pass-through becomes again significantly different from zero starting from the second half of 2008. At the end of the sample, a 1 percentage point decrease in one-year-ahead is estimated to trigger a 0.2 percentage point decrease in long-run expectations: therefore, the reduction of one-year-ahead inflation expectations observed from the second half of 2012 implies a reduction by 0.3 percentage points of long-run expectations. Apart from the magnitude of point estimates, a matter of concern for the euro area is the evidence of an inflation pass-through increasing up to 2014h2 and stabilizing at positive values thereafter, which is something we do not find for the other economies in the sample.[20]

4 Robustness

We perform the robustness analysis using as a benchmark the Gaussian Kernel estimation.

As a first robustness check for the euro area, we provide estimations for a sample starting in 2000, to avoid mixing observations belonging to different monetary regimes.[21] Figure 6 compares these new results with those from the previous section. In both cases, the positive co-movement at the end of the sample is statistically different from zero. Also the result of anchored expectations during the first years of the introduction of the euro does not change.

Next, we add some macroeconomic variables as controls in regression (2). As already mentioned, our working hypothesis until now has been that shocks known to professional forecasters at the time of the survey are embedded in short-run inflation expectations. However, one may argue that the pattern of volatile variables, such as the oil prices, and of structural macro

differently from those in their neighbourhood. In particular, by removing $1997h1$, $1997h2$ and $1998h1$, which display high negative correlation between $\Delta\pi^e_{t,5y}$ and $\Delta\pi^e_{t,1y}$, the kink in $1998h1$ visible in Figure 5 disappears and the coefficients become positive and significant up to $2007h1$. Higher frequency data could be more suited to efficiently estimate Japan's anchoring, given their high variability over time.

[20] In the Appendix, we replicate Figure 5 for a different choice of the parameter H. The results are basically unchanged.

[21] Benati (2008) argues that the introduction of the single currency may have changed the statistical properties of the inflation process.

variables, like the unemployment rate, may affect long-run expectations independently from short-run ones. Following Miccoli and Neri (2015), we thus add two controls: ΔP_t^{oil}, the change in the monthly average of the Brent crude oil one-month-forward price, and u_{t-1}, the latest available data on the unemployment rate (both as deviations from their sample means). The left panels of Figure 7 show that, for both the US and the euro area, the pass-through coefficients $\hat{\beta}_t$ are virtually the same as in the regressions with and without controls; moreover, as shown in the right panels, the effect of both controls on long-run expectations is never significantly different from zero.[22]

As a final robustness check, we performed a panel kernel analysis, using a time-varying version of regression (1). The evolution of the co-movements is very similar to that obtained with the country-by-country regressions. Tests to compare the coefficients across economies, performed in particular in the last part of the sample, yield the expected results: while we can conclude that the euro-area coefficient is statistically different from those for the UK and the US, the evidence is not conclusive with respect to Japan given, again, the low precision of the estimates for this economy.

5 Concluding remarks

We contribute to the literature on inflation expectations by estimating the degree of anchoring for the US, the UK, Japan and the euro area, using data and methods that yield coefficients that are comparable across economies and time. We used *Consensus* survey data of professional forecasters' inflation expectations to estimate a time-varying parameter model that allows the anchoring coefficient to change over time. We define expectations as being anchored when there is no co-movement between short- and long-run inflation expectations.

Our findings, summarized in Figure 8, show that the extent of anchoring has varied substantially across economies and over time. Expectations in the US and in the UK appear firmly anchored, particularly in more recent years. In Japan, although the estimates are very imprecise, we find that the co-movement of expectations is stronger than in all other economies throughout the period: even if in the last few years it has decreased somewhat, its magnitude

[22]For the euro area we chose to use only the post-euro sample, since it makes little sense to look at the effect of an aggregate unemployment rate before it.

in 2015h1 was still around 0.2. Finally, for the euro area we find that there was high and statistically significant co-movement before the introduction of the euro, anchored expectations during the 2000s, but signs of a de-anchoring in the last few years.

In a panel version of the Gaussian Kernel estimations, we test for statistical differences between countries' coefficients and find that the co-movement in the euro area is, at least in the final part of the sample, statistically different from that of the UK and the US.[23] Our analysis thus suggests that in the aftermath of the crisis there have been clear signs of a de-anchoring in Japan and in the euro area, but not in UK and US. Whether the differences in the magnitude of the de-anchoring pressures are due to the size and the nature of the shocks that have affected the economies or to the reactivity of the central banks is a central issue for policy makers and needs to be analyzed in further research.

[23]Since the standard errors for estimates on Japan are very high, the differences between the co-movement of expectations for this country and those for the other three economies are always statistically insignificant.

Tables and figures

Table 1: Time-invariant pass-through for different forecast horizon. Panel regression.

	$\Delta\pi^e_{t,5y}$	$\Delta\pi^e_{t,4y}$	$\Delta\pi^e_{t,3y}$	$\Delta\pi^e_{t,2y}$
$\Delta\pi^e_{t,1y}$	0.27**	0.42***	0.51***	0.49***
	(0.13)	(0.12)	(0.16)	(0.10)
$\Delta\pi^e_{t,1y}*d_{US}$	-0.23	-0.36***	-0.43***	-0.27**
	(0.15)	(0.13)	(0.16)	(0.13)
$\Delta\pi^e_{t,1y}*d_{UK}$	-0.23	-0.32**	-0.34*	-0.13
	(0.15)	(0.15)	(0.17)	(0.13)
$\Delta\pi^e_{t,1y}*d_{euro}$	-0.12	-0.22	-0.22	-0.01
	(0.14)	(0.13)	(0.17)	(0.13)
R^2	0.06	0.12	0.14	0.36
N	193	195	195	195

Note: Robust s.e. in parenthesis. $^*p < 0.1$,$^{**}p < 0.05$, $^{***}p < 0.01$.

Table 2: Wald Test. Panel regression.

	$\Delta\pi^e_{t,5y}$	$\Delta\pi^e_{t,4y}$	$\Delta\pi^e_{t,3y}$	$\Delta\pi^e_{t,2y}$
$\Delta\pi^e_{t,1y} + \Delta\pi^e_{t,1y}*d_{US} = 0$	0.51	1.36	2.09	7.01***
	(0.48)	(0.24)	(0.15)	(0.01)
$\Delta\pi^e_{t,1y} + \Delta\pi^e_{t,1y}*d_{UK} = 0$	-0.46	1.25	4.86**	21.63***
	(0.50)	(0.26)	(0.03)	(0.00)
$\Delta\pi^e_{t,1y} + \Delta\pi^e_{t,1y}*d_{euro} = 0$	15.44***	11.60***	14.02***	33.49***
	(0.00)	(0.00)	(0.00)	(0.00)

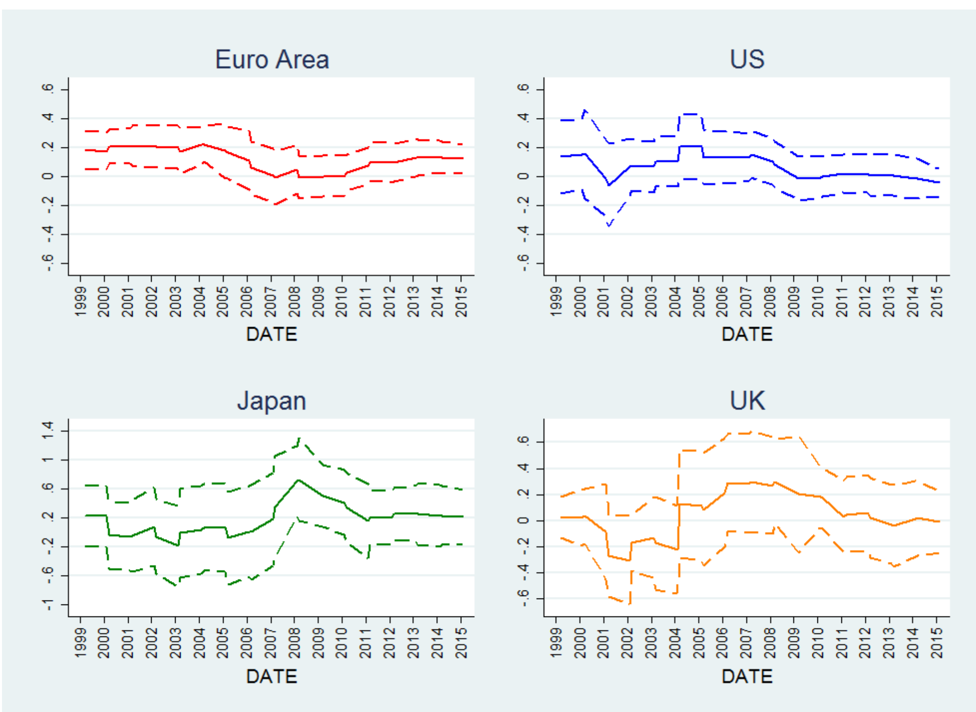
Note: F-stat and P-value in parenthesis. $^*p < 0.1$,$^{**}p < 0.05$, $^{***}p < 0.01$.

Table 3: DW test for residuals autocorrelation.

	$\Delta\pi^e_{t,5y}$	$\Delta\pi^e_{t,4y}$	$\Delta\pi^e_{t,3y}$	$\Delta\pi^e_{t,2y}$
Japan	2.0	2.0	2.4	2.1
US	2.2	2.2	2.3	2.1
UK	2.3	2.3	2.1	2.0
euro	2.1	1.8	2.0	1.8

Note: $d_u = 1.4$ and $d_l = 1.3$ critical values for $\alpha = 0.01$ Ho: No Autocorrelation. If $d < d_l$ reject Ho, if $d > d_u$ do not reject Ho, if $d_l < d < d_u$ test is not conclusive.

Figure 1: Inflation pass-through from one-year-ahead to five-year-ahead expectations: Rolling window estimation with n=20.



Note: Different y-scale for Japan.

Figure 2: Exponential weights at different observations, for varying H and n=20.

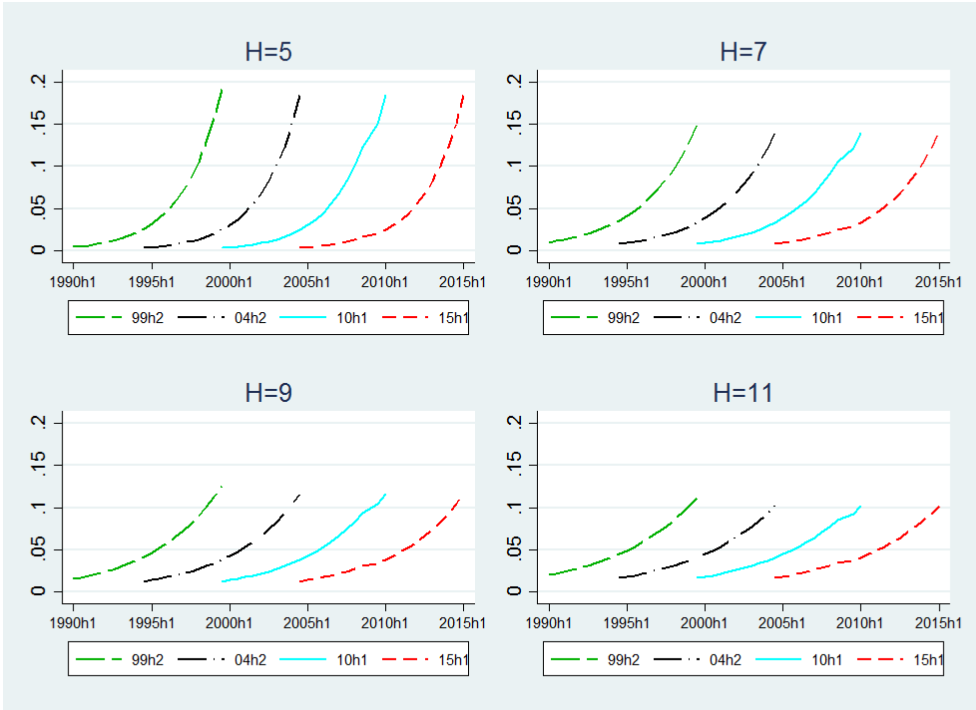
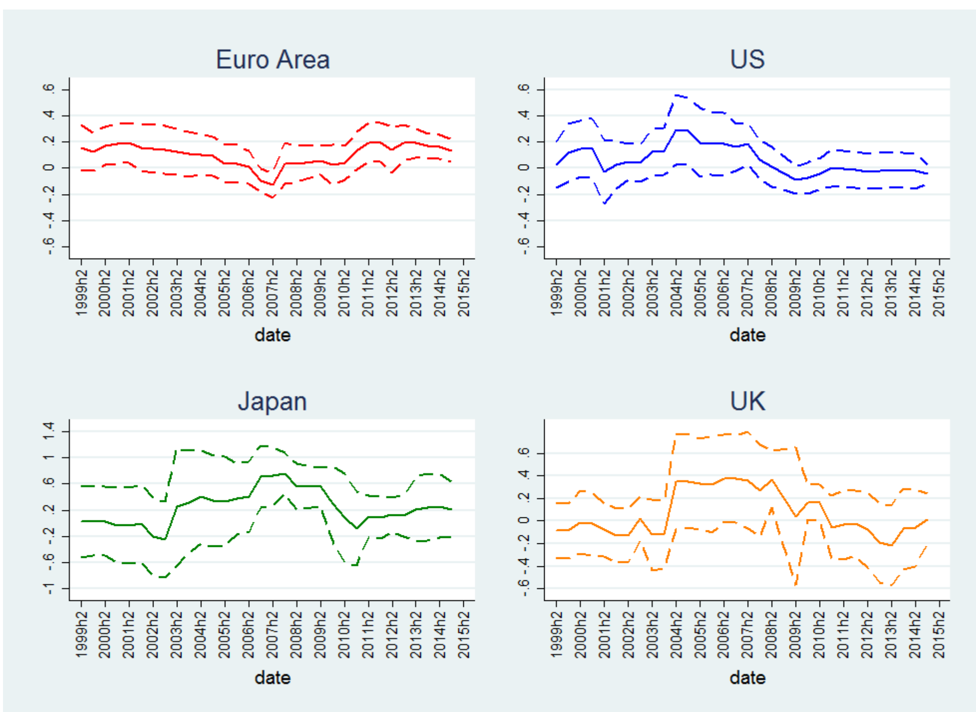


Figure 3: Inflation pass-through from one-year-ahead to five-year-ahead expectations: Exponential weighting estimation, H=7, n=20.



Note: Different y-scale for Japan.

Figure 4: Kernel weights at different observations, for varying H.

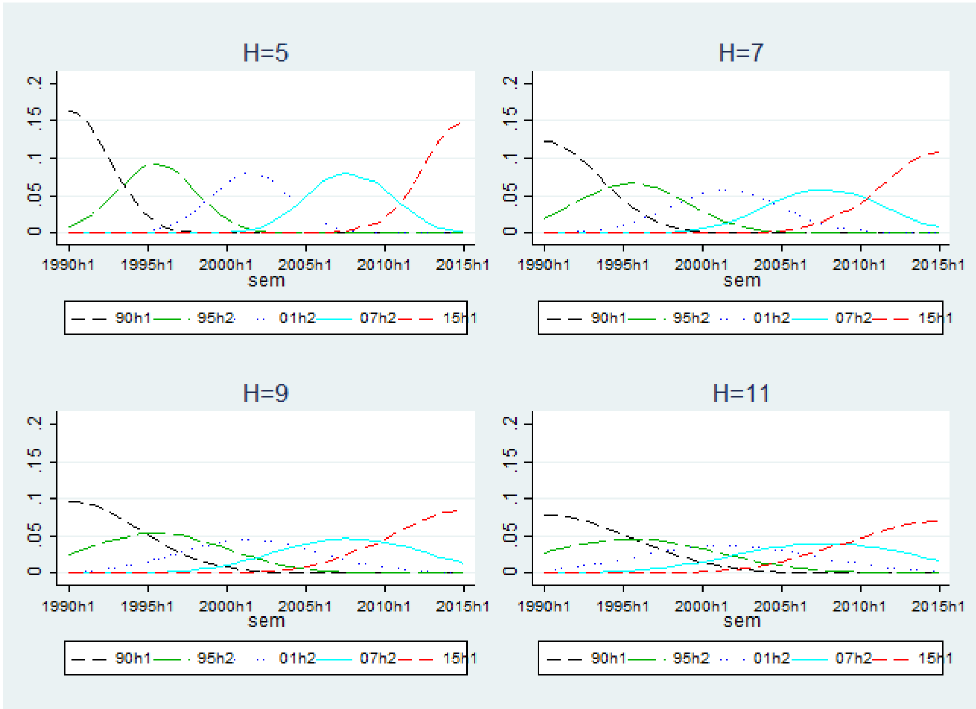


Figure 5: Inflation pass-through from one-year-ahead to five-years-ahead expectations: Gaussian weighting estimation, H=7, n=T.

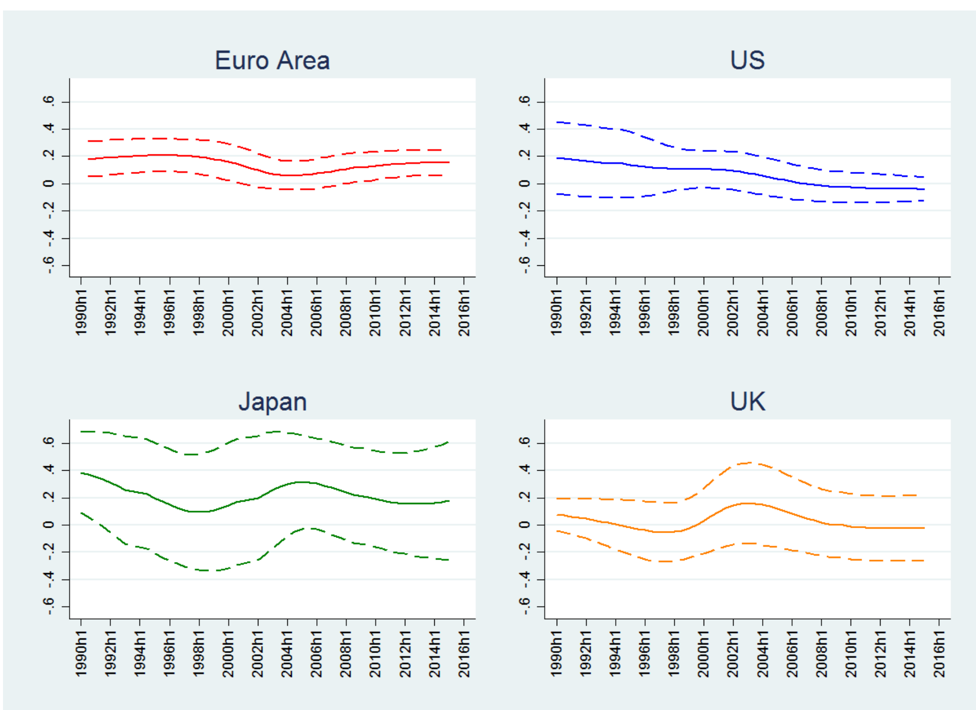


Figure 6: Inflation pass-through for the Euro Area: baseline estimation (red lines) versus sample excluding pre-euro observations (black lines).

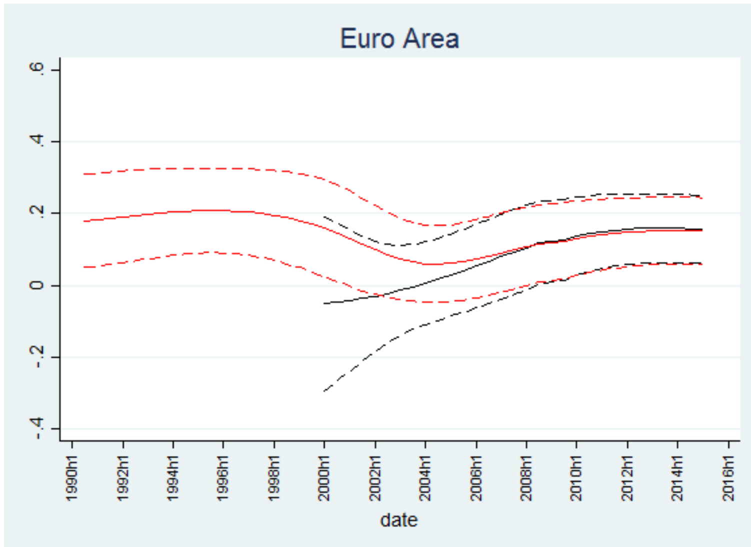


Figure 7: Inflation pass-through from one-year-ahead to five-year-ahead expectations, with and without controls (left panels) and controls' coefficients (right panels).

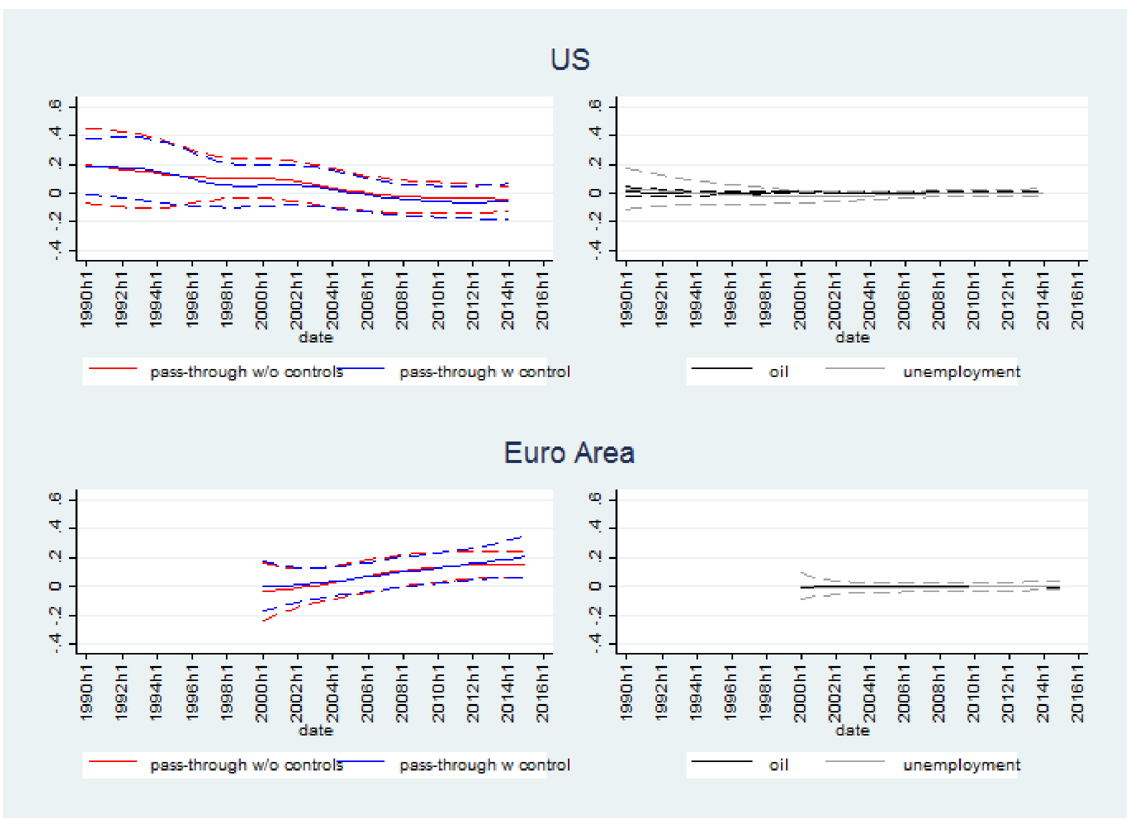
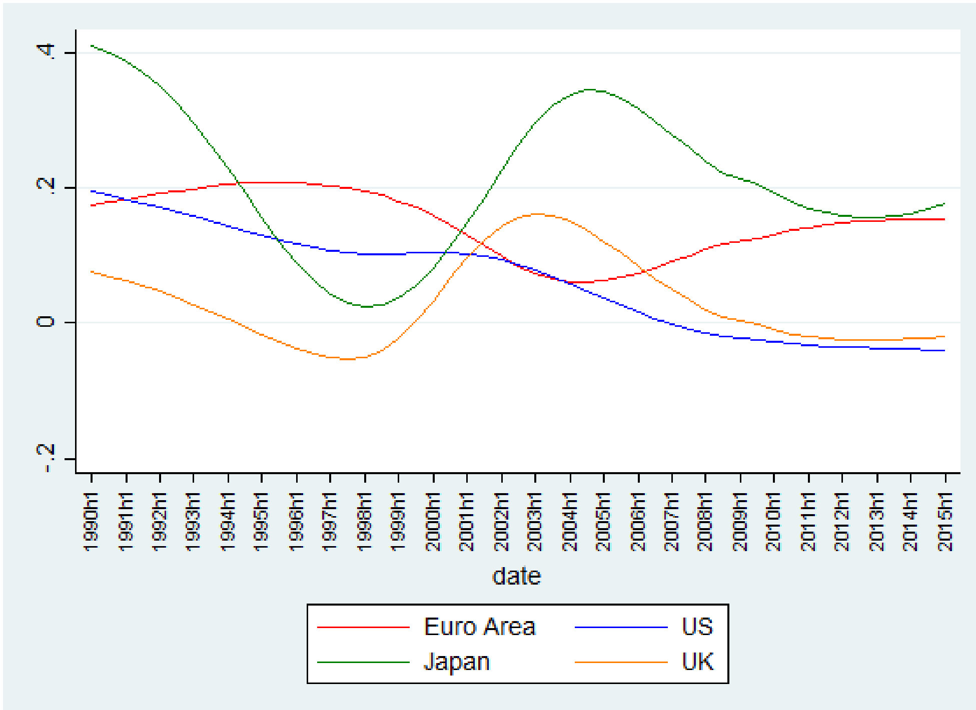


Figure 8: Inflation pass-through from one-year-ahead to five-years-ahead expectations, Gaussian weighting estimation, H=7, n=T.

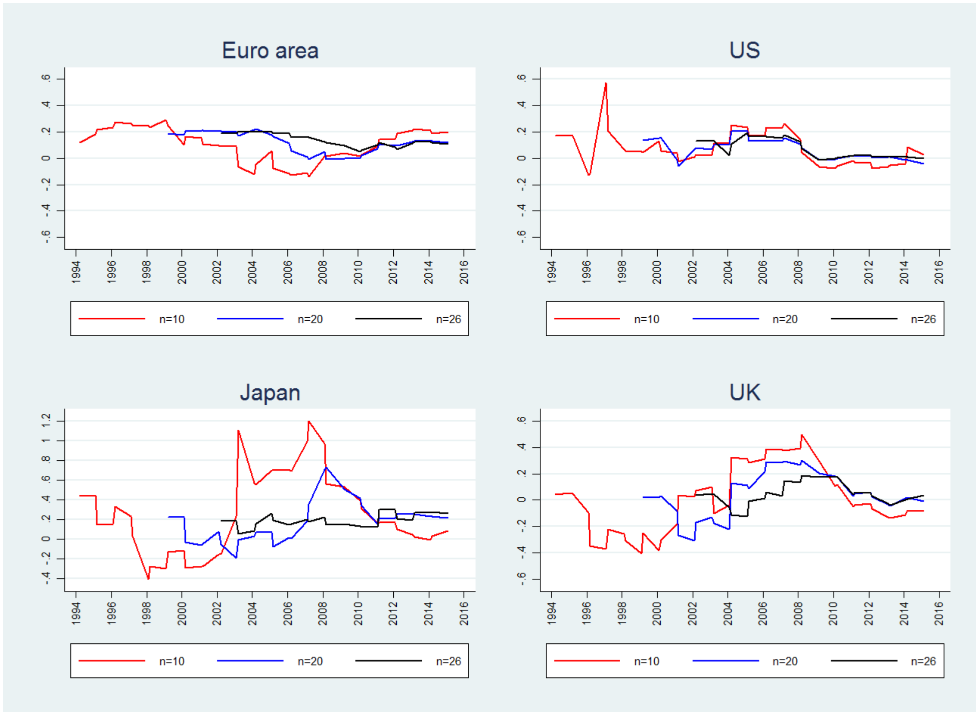


Appendix

Table A1: Summary Statistics

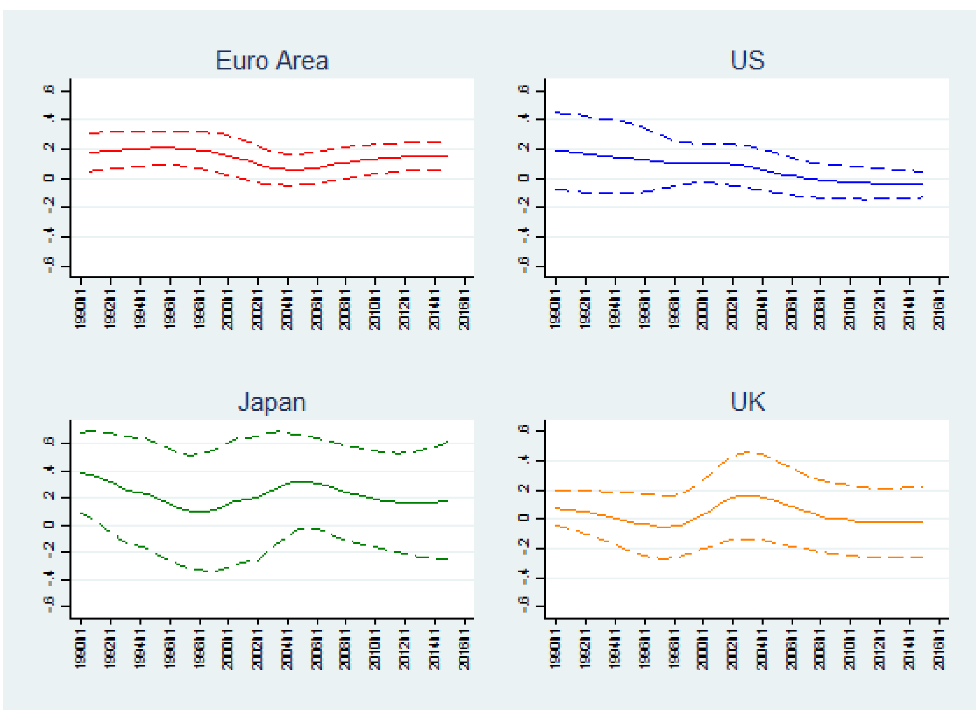
	MEAN		STD. DEVIATION		N	
	$\Delta\pi^e_{t,1y}$	$\Delta\pi^e_{t,5y}$	$\Delta\pi^e_{t,1y}$	$\Delta\pi^e_{t,5y}$	$\Delta\pi^e_{t,1y}$	$\Delta\pi^e_{t,5y}$
Euro Area	-0.06	-0.02	0.41	0.10	50	50
US	-0.07	-0.04	0.64	0.15	49	49
Japan	-0.01	-0.01	0.46	0.30	49	47
UK	-0.11	-0.05	0.55	0.27	51	51

Figure A1: Inflation pass-through from one-year-ahead to five-year-ahead expectations: rolling window estimation for varying n.



Note: Different y-scale for Japan.

Figure A2: Gaussian Kernel estimation for H=9.



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