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Does Communicating a Numerical Inflation Target Anchor Inflation Expectations? Evidence & Bond Market Implications*

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Abstract

High-frequency empirical evidence suggests that inflation expectations in the United States became better anchored after the Federal Reserve began communicating a numerical inflation target. Using an event-study approach, we find that forward measures of inflation compensation became unresponsive to news about current inflation after the adoption of an explicit inflation target. In contrast, we find that forward measures of nominal compensation in Japan continue to drift with news about current inflation, even after the Bank of Japan adopted a numerical inflation target. These empirical findings have implications for the term structure of interest rates in the United States. In a calibrated macro-finance model, we show that the apparent anchoring of inflation expectations implies a lower term premium in longer-term bond yields and decreases the slope of the yield curve.

JEL Classification: E31, E52, E58

Keywords: Monetary Policy, Inflation, Structural Breaks, Term Structure of Interest Rates

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1 Introduction

In January 2012, the Federal Open Market Committee (FOMC) adopted a numerical target of two percent inflation over the longer-run, which it believed was most consistent with its statutory mandates. Economic theory, such as Woodford (2003), predicts that such a policy change should lead to better economic outcomes. Anchoring long-run inflation expectations allows a central bank to respond aggressively to cyclical swings in the real economy without sacrificing its price stability mandate. Indeed, the FOMC alluded to such benefits when it adopted its long-run inflation objective:

"Communicating this inflation goal clearly to the public helps keep longer-term inflation expectations firmly anchored, thereby fostering price stability and moderate long-term interest rates and enhancing the Committee's ability to promote maximum employment in the face of significant economic disturbances." – Statement on Longer-Run Goals & Policy Strategy

One year after the Federal Reserve had adopted its inflation target, the Bank of Japan (BOJ) adopted a similar two percent inflation objective in January 2013. Merely publishing a numerical objective for inflation, however, does not necessarily cement inflation expectations at the central bank's target. For example, the announcement could lack credibility if the central bank failed to deliver on previous commitments. Instead, the degree to which inflation expectations are anchored remains an empirical question.

In this paper, we use data from bond markets to test whether the adoption of an explicit longer-run inflation objective better anchored inflation expectations in the United States (U.S.) and Japan. The anchored inflation expectations hypothesis implies a strong testable prediction, which we can evaluate statistically. In particular, if a central bank adopts a credible long-run inflation target, then expectations about inflation far in the future shouldn't respond to news about current inflation. In contrast, if inflation expectations are not well anchored, then recent inflation developments can sway longer-term inflation expectations. Using a high-frequency approach, we measure the responses of market-based measures of inflation compensation to data surprises contained in monthly Consumer Price Index (CPI) reports published by the Bureau of Labor Statistics in the U.S. and the Statistics Bureau in Japan.

Prior to the adoption of a numerical inflation target in the United States, we find that measures of far forward inflation compensation drift following inflation surprises. However, after the FOMC started communicating numerical objectives for longer-run inflation, a battery of econometric tests suggest that inflation compensation no longer comoves with inflation surprises. These results are consistent with an anchoring of inflation expectations as the FOMC moved towards adopting a numerical inflation target.

Structural break tests suggest a change in the relationship between inflation compensation and core inflation surprises shortly after FOMC participants began regularly publishing "longer-run" values for inflation in the quarterly Summary of Economic Projections (SEP). Moreover, narrative evidence also supports this timing of our estimated breakdate. For example, dialogue contained in the 2010 FOMC transcripts suggests that Committee members believed inflation expectations were anchored well before the formal adoption of the two percent inflation objective. Thus, by January 2012, we are able to statistically distinguish between the drifting inflation expectations regime of the late-1990's/2000's and the more recent anchored inflation expectations regime.

In contrast, we find no evidence of a similar anchoring in Japan. Our analysis suggests that inflation expectations continue to drift with inflation surprises in that country despite numerous changes in their monetary policy regime over the past two decades. However, the estimated response of nominal forward rates to Japanese inflation surprises appears to have diminished, albeit insignificantly, since 2013. Based on our findings from the U.S. experience, these recent results could signal the early stages of inflation expectations becoming better anchored in Japan. While we see some tentative signs of anchoring since the 2013 adoption of an inflation target, past inconsistencies between the Bank of Japan's communication and its policy actions has likely slowed the process of anchoring in Japan. This interpretation is consistent with the conclusions of De Michelis and Iacoviello (2016), which use structural models to argue that a lack of credibility has kept inflation expectations in Japan from moving closer to the Bank of Japan's target.

Finally, we show that these empirical findings have implications for the term structure of interest rates in the United States. Previous work by Rudebusch and Wu (2008) and Rudebusch and Swanson (2012) argues that drifting inflation expectations, a form of long-run nominal risk, is a key mechanism that helps macro-finance models match historical features of the U.S. yield curve. Using the Rudebusch and Swanson (2012) model, we examine the general-equilibrium implications of the anchoring of U.S. inflation expectations for Treasury yields. The model implies that anchoring inflation expectations reduces the average term premium on 10-year Treasury bonds by 5–16 basis points and flattens the yield curve.

2 Testing the Anchored Inflation Hypothesis

In this section, we present a simple model of longer-term inflation expectations that guides our intuition and our empirical specifications. Specifically, we model the evolution of long-term inflation expectations as follows:

$$\pi_t^{LT} = \pi_{t-1}^{LT} + \beta(\pi_t - \pi_{t-1}^{LT}), \tag{1}$$

where π_t^{LT} is the long-term inflation expectation in period t and π_t is the inflation rate in period t. We often associate π_t^{LT} with far forward measures of inflation expectations. This model builds upon the macro-finance literature of Gürkaynak, Sack and Swanson (2005), Rudebusch and Wu (2008), and Rudebusch and Swanson (2012), which finds that drifting long-term inflation expectations help explain characteristics of the U.S. Treasury yield curve.

The coefficient β determines the degree to which long-term inflation expectations are anchored. In one extreme, if $\beta = 1$ then long-term inflation expectation are completely unanchored and they move in lockstep with current inflation. On the other extreme, if $\beta = 0$, then long-term inflation expectations are anchored, in the sense that they are invariant realized inflation. For the intermediate cases that $0 < \beta < 1$, then inflation expectations drift with current inflation.

To test the degree to which inflation expectations are anchored, we must estimate the value of β. However, a simple regression of measures of long-term inflation expectations on current inflation is likely to yield biased estimates of β. Equation (1) is typically thought to be part of a larger macroeconomic model which contains an expectations-augmented Phillips curve, in which current inflation also depends on long-term inflation expectations. Therefore, the simultaneity between long-term inflation expectations when $\beta > 0$ and actual inflation makes the estimation of Equation (1) problematic.

However, a slight algebraic manipulation of Equation (1) allows us to easily estimate β directly using an event-study approach. If we take the expectations of Equation (1) at time $t-1$ and subtract it from Equation (1) above, we arrive at the following equation:

$$\pi_t^{LT} - \mathbb{E}_{t-1}\pi_t^{LT} = \beta(\pi_t - \mathbb{E}_{t-1}\pi_t), \tag{2}$$

where the right-hand side captures the news about current inflation that was revealed between time $t-1$ and t, and the coefficient β captures how that news about inflation affects long-term inflation expectations.

To measure β, we estimate Equation (2) using the one-day change in far forward yields around the release of CPI reports. Our preferred measure of π_t^{LT} is five-year, five-year forward inflation compensation implied by the spread between nominal Treasury yields and yields on Treasury Inflation-Protected Securities (TIPS). For the United States, we obtain daily data on this measure from the Federal Reserve Board. By focusing on forward measures of inflation compensation, we cleanse any direct effect that current inflation has on average expected inflation over the next decade. As an alternative, we also use the nominal one-year forward rate maturing in ten years (nominal one-year rate beginning in nine years) as a measure of forward nominal compensation. In U.S. data, for the samples that the two measures overlap, we find similar results. For Japan, where data on inflation-protected bond yields are unavailable, we use the nominal one year forward rate maturing in ten years as our measure of long-term inflation compensation.[1]

For our measure of $\pi_t - \mathbb{E}_{t-1}\pi_t$ in Equation (2), we use data surprises emanating from the release of monthly CPI reports.[2] For the U.S. and Japan, we measure $\mathbb{E}_{t-1}\pi_t$ using the median forecast from the surveys of professional forecasters compiled by Bloomberg prior to each data release. Furthermore, Bloomberg maintains data on the actual reported value of π_t in the report (i.e. not the revised value). For the U.S., we have forecasts and the actual release for the month-over-month percent change in CPI for both headline and core inflation. Using these two forecasts, along with the weight of core components in the CPI basket, we construct an implied food and energy surprise component. As we will show, however, our results are robust to using the percent change in an index of energy and agricultural prices as an alternative control for changes in non-core prices on the day of the CPI release.[3] Our sample periods are generally limited by the availability of data on inflation surprises, which starts in 1997 for the United States and in 2001 for Japan.

[1]For both the U.S. and Japan, we calculate nominal forward rates from the yield on constant maturity zero coupon bond yields as described in Gürkaynak, Levin and Swanson (2010).

[2]Our regression model is therefore very similar to the model in Gürkaynak, Levin and Swanson (2010) except we focus exclusively on CPI reports (i.e. news about inflation) as prescribed from Equation (2).

[3]For Japan, using this additional control variable is essential since we are unable to infer the weights on the core component (which is prices excluding fresh food). Also, for Japan, our inflation surprises are for the year-over-year percent change in the core CPI inflation as opposed to the month-over-month percent change. Although this may have implications for interpreting the magnitude of β, the scaling in no way affects hypothesis tests against the null hypothesis of $\beta = 0$.

3 Inflation News & Inflation Compensation in the U.S.

Did the relationship between market-implied inflation compensation and unexpected news about inflation change around the time that the FOMC adopted an explicit inflation objective? We apply several different statistical methods to detect such a possible change. First, we look for a structural break using split-sample regressions, which impose a break in the regression model after the policy change. Using this approach, we find evidence consistent with an anchoring of inflation expectations after 2012. To get a more precise sense of when inflation expectations became anchored, we then apply tests for a structural break at an unknown date. These break tests suggest that the relationship between inflation compensation and inflation news changed in the first half of 2010. Although this candidate break date is about two years prior to the FOMC's adoption of a numerical inflation objective, it follows shortly after the Committee began publishing numerical ranges for "longer-run" inflation in the quarterly Summary of Economic Projections. Finally, we show that both narrative evidence as well as rolling window regressions further support these conclusions.

3.1 Did a Numerical Target Better Anchor Expectations?

Using our core CPI surprise and the surprise associated with the food and energy components, we estimate the following event-study regression to measure β, the response of long-term inflation compensation to news about inflation,

$$\Delta\pi_t^{LT} = \alpha + \beta\pi_t^{core} + \gamma\pi_t^{fe} + \varepsilon_t, \tag{3}$$

where $\Delta\pi_t^{LT}$ is the one-day change in the 5-year, 5-year forward measure of inflation compensation on the day of a CPI release, π^{core} is the core CPI surprise, and π^{fe} is the surprise associated with the food and energy component.[4] We will refer to this specification as the "inflation compensation model." We estimate Equation 3 using ordinary least squares where each observation corresponds to a given CPI release. We split our data into two distinct sample periods to determine if the underlying relationship between inflation surprises and inflation compensation changed after the FOMC's adoption of an explicit inflation target. First, we examine the January 1999 – December 2011 sample period, which is prior to the inflation target adoption. Then, we examine the January 2012 – October 2017 period following the policy change.

[4] We scale the core CPI surprises by the weight of core items in the CPI basket. Our results are insensitive to this scaling.

We find statistically significant evidence that inflation compensation responds less to economic news about inflation following the adoption of the inflation target. The first two columns of Table 1 show the estimated coefficients from Equation 3 across sample periods. Prior to January 2012, a positive core CPI surprise leads to a statistically significant increase in inflation compensation. A ten-basis point core CPI surprise typically raises five-year, five-year forward inflation compensation by one and half basis points. After the FOMC formally adopting its inflation target, however, the coefficient on core CPI falls and is statistically indistinguishable from zero. In the third column of Table 1, we formally conduct a Chow (1960) test, which suggests the presence of a structural break in β in 2012. The post-January 2012 dummy variable that interacts with the core inflation surprise is negative and statistically significant, which suggests a change in the relationship between news about core inflation and longer-term inflation compensation after the FOMC adopted its formal inflation target.

The break in the estimate of β appears to reflect a change in the reaction of inflation expectations to CPI surprises after 2012, rather than a change in the nature of CPI surprises. Table 2 shows the summary statistics for CPI surprises both before and after 2012. The standard deviation of the Bloomberg core inflation surprises is equal to 0.072 prior to 2012 and 0.066 thereafter. Furthermore, the surprises in both periods are not significantly skewed nor is there evidence that they are non-normal as the Jarque-Bera statistic falls below its critical value in both samples. The most notable difference between the two samples is the presence of an average downside inflation surprise after 2012. The negative mean surprise largely reflects the very most recent string of downside core CPI surprises beginning in March of 2017. From the viewpoint of our regression model, this change in the distribution of inflation surprises, all else equal, has the potential to impact the intercept α, but not the slope coefficient β. However, Table 1 shows that we find no statistically significant evidence of a change in the regression intercept across the two sample periods.

We consistently find that CPI surprises associated with the food and energy components of the consumption basket don't significantly affect longer-term inflation compensation. Two intuitive reasons support this empirical finding. First, to the extent that food and energy price fluctuations are short-lived and often reverse, we would expect them not to have a big effect on longer-term inflation expectations. Second, thanks to vibrant spot and derivatives markets based on food and energy commodities, bond investors already have some information about the food and energy components ahead of the CPI release. Both prior to and after the adoption of the inflation target, we find that the coefficients on the food and energy surprises remain near zero and this relationship appears stable over time.

3.2 Testing for a Structural Break at an Unknown Date

Rather than imposing a break after 2012 in the relationship between far forward inflation compensation and inflation surprises, we now estimate the most likely timing of the break. In particular, we test for a structural break at an unknown date. The estimated breakdate allows us to provide some further interpretation of the source of the break in the β coefficient. If a change in β reflected a better anchoring of inflation expectations, we would expect the estimated break date to lag or coincide with a change in U.S. monetary policy. If instead the estimated break date is not supported with corroborating narrative evidence, it could reflect general instability in the regression model rather than deep structural change.

Tests for a structural break at an unknown date reveal evidence of a break in β, but not any of the other parameters in the regression model. Table 3 shows the results of Andrews (1993) and Quandt (1960) test and the Andrews and Ploberger (1994) test for a structural break in the inflation compensation regression model in Equation 3. Both tests suggests that the relationship between inflation compensation and inflation news changed in May 2010. The candidate break date is significant at the 5% level for both tests, indicating strong statistical evidence of a change in β. There is no evidence of a break in any of the other regression parameters, including the variance of the regression residual. Since our regression model has a relatively low R^2, instability due to changes in the liquidity of the TIPS market over time would likely appear as a break in the residual variance. Importantly, we find no evidence of such a break.

The solid black line in the top panel of Figure 1 plots the time series of Chow test statistics for a break in β. The breaktest sequence has a fairly well-defined maximum at the estimated break date with no other clear peaks. This pattern suggests a one-time structural break in the sensitivity of long-term inflation expectations to core CPI surprises occurring around 2010.[5]

The estimated break date occurs before the Federal Reserve formally adopted a two percent inflation target in January of 2012, but shortly after FOMC participants began to publish quarterly projections for "longer-run" inflation. In April of 2009, the Summary of

[5] More formally, if we split the sample into two subsamples 1999-2010 and 2010-2017, the Andrews-Quandt test indicates no other breaks while the Andrews-Ploberger test indicates weak evidence (significant at the 10% level) of a second break in November of 2012. The 2010 break is again found with a high level of statistical significance from both the Andrews-Quandt and Andrews-Ploberger test when we perform the refinement proposed in Bai (1997), which tests for a break at an unknown date from 1999-2012.

Economic Projections (SEP) added longer-run inflation which, according to the language summarizing these economic projections, "[...] represent each participant's assessment of the rate to which each variable would be expected to converge under appropriate monetary policy and in the absence of further shocks to the economy." According to most economic theories, monetary policy solely determines inflation in the longer run. Therefore, one interpretation of this estimated break date is that public expectations began to fixate on these projections as an initial numerical range for the FOMC's longer-term inflation objective.

This interpretation of the timing and the source of the break in β aligns with the thinking of the FOMC at the time. In an October 2010 conference call, the Committee discussed making changes to the FOMC's policy and communication framework. As the Committee debated specific language to describe its inflation objective, then recently appointed Vice-Chair Janet Yellen asserted that the FOMC's SEP was serving to provide numerical guidance around the FOMC's inflation objective: "I don't think we need to seek Committee agreement on a single specific inflation target at this point. To say it is 'about two percent,' or 'two percent or a bit less,' strikes me as an accurate characterization of our SEP responses." Vice-Chair Yellen went on to say that while she supported the Committee's adoption of a numerical inflation target, "The Committee's objectives are already pretty well understood by markets, so they'll probably get the message without the numbers."

3.3 Robustness to Alternative Data, Samples, & Specifications

Our baseline model shows that market-based measures of inflation expectations became less sensitive to news about inflation after the FOMC began to communicate a numerical inflation objective. We now show that this finding is robust to a number of alternative specifications. Specifically, we illustrate this robustness using: (i) alternative measures of nominal compensation and food and energy price controls, (ii) data samples that exclude the global financial crisis, and (iii) specifications that allow for more gradual parametric change. Under all these alternatives, we find that nominal compensation responds less to inflation news after the adoption of the explicit inflation objective.

In our baseline inflation compensation model, we proxy forward inflation expectations by using inflation compensation measured from inflation-indexed bonds. However, TIPS yields may contain a non-trivial, time-varying liquidity premium which could distort our measure of inflation expectations.[6] Our baseline model also uses the weight of core goods

[6] As long as this premium is uncorrelated with core inflation surprises, our baseline results remain unbiased.

and services in the overall CPI basket, along with the headline and core CPI surprises, to infer the information content emanating from food and energy components. While this weight varies little month to month, its value is not exactly known in real time. To address both of these concerns, we instead estimate the following alternative regression model:

$$\Delta y_t^{LT} = \alpha + \beta \pi_t^{core} + \gamma^f \pi_t^{food} + \gamma^e \pi_t^{energy} + \varepsilon_t, \tag{4}$$

where Δy_t^{LT} is the one-day change in the one-year, nine-year forward rate around CPI announcements and π_t^{food} and π_t^{energy} are the one-day percent changes in the Goldman Sachs agricultural and energy price indexes, respectively. We will refer to this specification as the "forward rate model."

Rather than using inflation compensation measured from inflation-indexed bonds, this alternative model uses far forward measures of nominal compensation as a proxy for long-term inflation expectations. Although real factors could influence this measure of forward compensation, Gürkaynak, Sack and Swanson (2005) argue that most macroeconomic models would predict that real variables return to their steady state values following a disturbance before nine years. In addition, this specification uses the change in spot prices for food and energy inputs instead of the implied surprise from the CPI measure of food and energy prices. Given that timely information on the previous month's food and energy prices already available to bond investors at the time of the CPI release, the change in spot prices for food and energy inputs might be a more appropriate control for these non-core items on the day of the CPI release.

Table 4 reports our regression estimates for the United States using the forward rate model, which continues to show a statistically significant decline in the response of inflation compensation to inflation news following the adoption of the inflation target.[7] The robustness of our findings using the forward rate model is important as we move to our cross-country analysis. In Section 4, we repeat a similar exercise for Japan. For that country, however, we lack data on real (inflation-indexed) bonds and the knowledge about the weight of core components in the CPI basket. Thus, we cannot estimate our preferred inflation compensation model specified in Equation 3. However, we can estimate the forward rate model.

Using this alternative model, tests for a structural break at an unknown date also suggest a break in the coefficient on the core inflation surprise around 2010. The solid black line

[7] We no longer scale the core CPI surprises by the weight of core components in the CPI basket.

in the second panel of Figure 1 plots the Chow test sequence for β using the forward rate regression model over time. Once again, we see a clear peak in the time series of the test statistic in the first half of 2010. However, there is also a sharp spike in the sequence of Chow tests in late 2008. Table 5 shows that the test statistic at the December 2008 break date exceeds the 10% critical value of the Andrews-Quandt test. However, the local maximum in 2010 also exceeds this critical value. The presence of two local maxima could signal either two breaks or, based on the timing, instability during the financial crisis. This latter possibility of instability in the regression model due to the global financial crisis leads us to further examine the robustness of our candidate break dates.

If we drop the precipice of the global financial crisis, we find evidence indicating the presence of a single structural break in 2010. For the inflation compensation model, Table 6 shows that if we drop the fourth quarter of 2008 and first quarter of 2009 from the estimation, we estimate the exact same break date of May of 2010 for the core inflation coefficient. The blue dotted lines in Figure 1 plot the time series of the Chow statistics for samples that exclude the financial crisis. For both regression models, the presence of a peak in the time series of the break statistics in 2010 is insensitive to the inclusion or exclusion of the financial crisis. After excluding the precipice of the financial crisis, Table 7 shows that the estimated break-date for the forward rate model is February of 2010, which also matches the maximum of the sequence of Chow test statistics. This finding suggests that the source of instability in the response of forward bond yields to inflation surprises occurring around 2010 is not simply a reflection of financial market volatility but, instead, is likely due to deeper structural change.

Rolling-window regressions also suggest a similar decline in β over time. This alternative approach to measuring the time variation in the sensitivity of long-term inflation expectations to inflation surprises is well suited to capture a more gradual changes in the coefficients over time. The top panel of Figure 2 illustrates the time variation in β from the inflation compensation regression model specified in Equation (3) using 10-year rolling samples. We observe the same pattern of structural change as our previous findings. Early in the sample, prior to 2012, β is estimated to be statistically significant and positive. However, by 2010, the point estimate of β begins to decline and falls to values not different from zero by 2012. The second panel of Figure 2 shows similar time variation in β as estimated from the forward rate regression model in Equation 4. The point estimate of β from the forward rate model is positive and close to being significant in 2010, before beginning a steady descent.

4 Inflation News & Inflation Compensation in Japan

About one year after the FOMC formally adopted its longer-run inflation target of two percent, the Bank of Japan (BOJ) followed suit. After years of deflation and slow growth, Shinzo Abe campaigned on a platform of reflation through an official inflation target and aggressive quantitative easing. After taking office in December of 2012, Prime Minister Abe appointed Haruhiko Kuroda as the Governor of the Bank of Japan. Shortly after taking office, Governor Kuroda implemented a more aggressive quantitative easing campaign, which was further expanded in October 2014. In January 2016, the BOJ implemented a negative interest rate on reserves policy. Later that year, the Bank pursued a policy of yield curve control, which buys and sells bonds as necessary to achieve a 0% yield on 10-year Japanese government bonds.

Given this narrative evidence of several regimes changes in Japanese monetary policy, we empirically evaluate whether the adoption of these policies has better anchored inflation expectations in Japan. To this end, we look for evidence of parameter instability in the following statistical model for Japan:

$$\Delta y_t^{LT} = \alpha + \beta \pi_t^{core} + \gamma^f \pi_t^{food} + \varepsilon_t. \tag{5}$$

where Δy_t^{LT} is the one-day change in a 1-year, 9-year forward rate around Japanese CPI announcements, π^{core} is the core Japanese CPI surprise (which excludes the price of fresh food), and π_t^{food} is the one-day percent change in the Goldman Sachs agricultural price index.

We find no evidence of a change in the response of nominal forward rates to core inflation surprises in Japan. Following the same strategy as we did for the U.S., we initially impose a break in the regression relationship in January of 2013, after the election of Shinzo Abe. Table 8 shows the split-sample regression estimates over the 2001-2012 sample period and 2013-2017 sample periods. In both subsamples, the estimate of β, the coefficient on the core inflation surprise, is positive and statistically significant. However, the point estimate of β in the more recent sample is about half the size compared to its pre-Abe/Kuroda estimate. This finding may suggest some initial signs of anchoring in Japan. However, a more formal Chow (1960) test yields no statistically significant evidence of a break in β.

Using structural break tests for an unknown date, we can more generally test for a break in the relationship between nominal forward yields and core inflation surprises. However, unlike our findings for the United States, these tests indicate *stability* in the regression model in Equation (5) over time. The Andrews-Quandt and Andrews-Ploberger tests indicate no evidence of significant time variation in β. The bottom panel of Figure 1 shows the Chow test sequence over candidate breakdates. The time series of Chow tests has no well defined peaks that exceed the 10% critical value for a structural break. Rolling-window regressions also support these findings of a lack of structural change. The bottom panel of Figure 2 illustrates estimates of the sensitivity of nominal forward rates to core inflation surprises over 10-year rolling windows advanced by one month at a time. As with the split-sample estimates and the break tests, the time series of estimated β coefficients suggests a positive and stable relationship between inflation news and inflation compensation in Japan. This evidence indicates that despite the host of policy changes implemented by the BOJ, inflation expectations – as perceived by bond market investors – remain unanchored.

5 Implications for U.S. Bond Markets

To this point, we have used data on inflation compensation to analyze how bond markets perceive the inflation objectives of central banks. In this section, we now explore the general-equilibrium effects of the apparent anchoring of U.S. inflation expectations for the term structure of U.S. interest rates. Drifting inflation expectations embody a long-run risk for holders of nominal debt. As a result, Rudebusch and Swanson (2012) and others show that this mechanism helps dynamic, general-equilibrium models simultaneously match macroeconomic and financial market moments. Specifically, they find that that drifting inflation expectations help these models generate a significantly positive average term premium on long-term nominal bonds and an upward sloping yield curve. By reducing long-run nominal risk, these models predict that anchoring inflation expectations should reduce the term premium and flatten the yield curve.[8]

Using the calibrated model of Rudebusch and Swanson (2012), we now quantitatively assess the implications of anchored inflation expectations for the average term premium and slope of the yield curve. Similar to our Equation (1), Rudebusch and Swanson (2012) assume that long-term inflation expectations drift according to the following process:

$$\pi_t^{LT} = \rho_\pi \, \pi_{t-1}^{LT} + \vartheta_\pi \left(\bar{\pi}_t - \pi_t^{LT} \right), \tag{6}$$

where $\bar{\pi}_t$ is an infinite sum of past inflation with geometrically declining weights on more

[8]Rudebusch and Swanson (2012) show that price-level targeting, which completely eliminates inflation risk to nominal bond holders, implies that the term premium and the slope of the yield curve are essentially zero in their model.

distant inflation.[9] Using a moment matching exercise, Rudebusch and Swanson (2012) find that a value of $\vartheta_\pi = 0.003$ helps the model jointly match macroeconomic and yield curve moments. They also consider a calibrated model, in which they set $\vartheta_\pi = 0.01$, which is much more inline with our high-frequency empirical evidence.[10]

On average, the anchoring of inflation expectations causes a modest decline in the term premium and flattens the yield curve. The second two columns of Table 9 illustrate the average term premium and slope of the yield curve under the two different calibrations in Rudebusch and Swanson (2012). In the final column, we solve the model setting $\vartheta_\pi = 0$, which is consistent with our high-frequency evidence that inflation expectations became better anchored over the past few years. Depending on the calibration, the average term premium on a 10-year Treasury bond declines by 5–16 basis points. Moreover, we see that the yield curve is a bit flatter when inflation expectations do not respond to current inflation developments. These results highlight that small changes in agents' views about the central bank's inflation objective can have significant general-equilibrium implications for the term structure of interest rates.

6 Conclusions

Almost ten years ago, Gürkaynak, Levin and Swanson (2010) conducted a detailed, cross-country analysis on the effects that numerical inflation targeting has on long-run inflation expectations. They concluded that inflation expectations were generally better anchored in the UK and Sweden than in the United States because, at that time, the United States had not yet adopted an explicit long-term inflation objective. Our paper provides two key results which further their influential work. First, our results highlight an out-of-sample test of their findings. Indeed, we find that inflation expectations became better anchored as the FOMC began communicating a numerical value for its longer-term inflation objective. Specifically,

[9] In their model, Rudebusch and Swanson (2012) also incorporate exogenous shocks to long-term inflation expectations, which are unrelated to current inflation outcomes. Numerically, these shocks have only very small implications for the average term premium and slope of the yield curve. Therefore, we set the volatility of these shocks to zero in our analysis and focus on the endogenous drifting of inflation expectations due to realized inflation outcomes.

[10] Rudebusch and Swanson (2012) calibrate their model at a quarterly frequency, while our high-frequency estimation of Equation 2 relates monthly inflation to changes in annualized inflation compensation. After accounting for the difference in data frequencies, as well as the calibration of the lag polynominal used to calculate $\bar{\pi}_t$ in their model, our high-frequency baseline model would imply a value of $\vartheta_\pi = 0.04$ over the 1999–2011 sample period.

we find a reduction in the sensitivity of far forward measures of nominal compensation to current inflation surprises in the United States after 2012. Thus, we find evidence that further validates their analysis and conclusion; a credible, numerical inflation objective can better anchor inflation expectations.

However, our second key result provides a caveat to the applicability of this result. Using data from Japan, we find that simply communicating an inflation target may be insufficient to anchor inflation expectations. Although the BOJ adopted an explicit numerical inflation objective about one year after the Federal Reserve, we find no statistically significant evidence that, as of yet, far forward nominal bond prices are less responsive to CPI surprises. However, the point estimate of the degree of sensitivity has fallen since the adoption of the BOJ's inflation target and aggressive monetary easing. These results for Japan suggest that words must be accompanied by either expected or actual actions to achieve credibility around an inflation target. Even if the central bank announces a numerical inflation objective, poor credibility and past policy actions may slow or prevent inflation expectations from becoming anchored.

Our results also suggest that monitoring the sensitivity of far forward nominal compensation to the flow of inflation data is a valuable tool for understanding how bond markets perceive monetary policy. Given the changes in the response of inflation compensation to inflation news that we document in this paper, further modeling the time variation in this relationship could be fruitful. In particular, maintaining more general time-varying parameter models could enhance real-time surveillance of inflation expectations for central banks.

Finally, we show that anchoring inflation expectations has implications for the shape of the yield curve. All else equal, a more credible commitment to a numerical inflation target reduces long-run nominal risk to bond holders. Therefore, anchoring inflation expectations reduces the extra compensation investors require for bearing inflation risk over the life of the bond (the term premium) and flattens the yield curve. We show, through the lens of a standard macro-finance model, that this effect may be quantitatively meaningful.

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Table 1: U.S. Inflation Compensation Model

	5-Year, 5-Year Forward Inflation		
	1999-2011	2012-2017	1999-2017
Constant	0.00	0.00	0.00
	(0.00)	(0.00)	(0.00)
Core CPI surprise	0.15^{**}	-0.07	0.15^{**}
	(0.07)	(0.08)	(0.07)
Food & Energy CPI surprise	-0.02	-0.02	-0.02
	(0.04)	(0.06)	(0.04)
Constant $\times\mathcal{I}_{t\geq 2012}$			0.00
			(0.01)
Core CPI surprise $\times\mathcal{I}_{t\geq 2012}$			-0.22^{**}
			(0.10)
Food & Energy CPI surprise $\times\mathcal{I}_{t\geq 2012}$			0.00
			(0.07)
Observations	155	68	223
R^2	0.04	0.01	0.03

Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$, $^{**}p < 0.05$

Table 2: Summary Statistics of U.S. Core CPI Inflation Surprises

	1997-2011	2012-2017
Mean	0.00	−0.02
	(0.63)	(0.01)
Standard Deviation	0.07	0.07
Skewness	0.16	−0.33
	(0.39)	(0.28)
Kutroisis	−0.33	0.41
	(0.37)	(0.51)
Jarque-Bera	1.58	1.72
	(0.45)	(0.42)
Observations	179	70

Note: p-values in parenthesis.

Table 3: U.S. Inflation Compensation Model: Structural Break Tests at an Unknown Date

		5-Year, 5-Year Forward Inflation	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2002:08	1.78	0.24
		(0.84)	(0.73)
Core CPI surprise	2010:05	9.76^{**}	2.07^{**}
		(0.03)	(0.05)
Food & Energy CPI surprise	2010:07	1.14	0.13
		(0.98)	(0.95)
All Coefficients	2010:05	11.37	2.78
		(0.13)	(0.20)
Residual Variance	2013:03	2.38	0.50
		(0.69)	(0.44)

Note: Approximate asymptotic p-values from Hansen (1997) in parenthesis.
Observations: 223
$^{*}p < 0.10$, $^{**}p < 0.05$

Table 4: U.S. Forward Rate Model

	1-Year, 9-Year Forward Rate		
	1997-2011	2012-2017	1997-2017
Constant	0.00	0.00	0.00
	(0.01)	(0.01)	(0.01)
Core CPI surprise	0.11^{*}	-0.08	0.11^{*}
	(0.06)	(0.07)	(0.06)
GS Agriculture Price Index	0.00	0.02	0.00
	(0.01)	(0.01)	(0.01)
GS Energy Price Index	0.00	0.01^{**}	0.00
	(0.00)	(0.00)	(0.00)
Constant $\times \mathcal{I}_{t \geq 2012}$			0.00
			(0.01)
Core CPI surprise $\times \mathcal{I}_{t \geq 2012}$			-0.18^{*}
			(0.09)
GS Agriculture Price Index $\times \mathcal{I}_{t \geq 2012}$			0.01
			(0.01)
GS Energy Price Index $\times \mathcal{I}_{t \geq 2012}$			0.00
			(0.00)
Observations	179	69	248
R^2	0.04	0.13	0.06

Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$

Table 5: U.S. Forward Rate Model: Structural Break Tests

		1-Year, 9-Year Forward Rate	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2003:09	2.94	0.40
		(0.57)	(0.53)
Core CPI surprise	2008:12	7.72*	2.06**
		(0.08)	(0.05)
GS Agriculture Price Index	2003:09	5.10	1.24
		(0.24)	(0.14)
GS Energy Price Index	2008:10	4.25	0.75
		(0.34)	(0.29)
All Coefficients	2008:12	10.22	3.23
		(0.36)	(0.26)
Residual Variance	2013:11	1.85	0.19
		(0.82)	(0.81)

Note: Approximate asymptotic p-values from Hansen (1997) in parenthesis.
Observations: 248
$^{*}p < 0.10$, $^{**}p < 0.05$

Table 6: U.S. Inflation Compensation Model: Structural Break Tests Excluding Financial Crisis

		5-Year, 5-Year Forward Inflation	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2002:08	2.13	0.27
		(0.75)	(0.70)
Core CPI surprise	2010:05	9.91^{**}	2.13^{**}
		(0.03)	(0.05)
Food & Energy CPI surprise	2010:07	1.20	0.13
		(0.97)	(0.93)
All Coefficients	2010:05	10.90	2.57
		(0.15)	(0.24)
Residual Variance	2001:09	2.55	0.60
		(0.65)	(0.37)

Note: Approximate asymptotic p-values from Hansen (1997) in parenthesis.
Observations: 218
$^{*}p < 0.10,^{**}p < 0.05$

Table 7: U.S. Forward Rate Model: Structural Break Tests Excluding Financial Crisis

		1-Year, 9-Year Forward Rate	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2003:09	1.99	0.24
		(0.79)	(0.71)
Core CPI surprise	2010:02	10.29**	2.72**
		(0.02)	(0.02)
GS Agriculture Price Index	2003:09	5.04	1.23
		(0.24)	(0.14)
GS Energy Price Index	2011:05	5.20	0.88
		(0.23)	(0.24)
All Coefficients	2003:09	14.08	4.42*
		(0.11)	(0.10)
Residual Variance	2013:11	2.06	0.37
		(0.77)	(0.56)

Note: Approximate asymptotic p-values from Hansen (1997) in parenthesis.
Observations: 242
$^{*}p < 0.10, ^{**}p < 0.05$

Table 8: Japan Forward Rate Model

	1-Year, 9-Year Forward Rate		
	2001-2012	2013-2017	2001-2017
Constant	0.00	0.00	0.00
	(0.01)	(0.00)	(0.01)
Core CPI surprise	0.13^{**}	0.07^{*}	0.13^{**}
	(0.06)	(0.04)	(0.06)
GS Agricultural Price Index	0.00	0.00	0.00
	(0.00)	(0.00)	(0.00)
Constant $\times \mathcal{I}_{t \geq 2013}$			0.00
			(0.01)
Core CPI surprise $\times \mathcal{I}_{t \geq 2013}$			-0.06
			(0.08)
GS Agricultural Price Index $\times \mathcal{I}_{t \geq 2013}$			0.00
			(0.00)
Observations	136	57	193
R^2	0.02	0.04	0.02

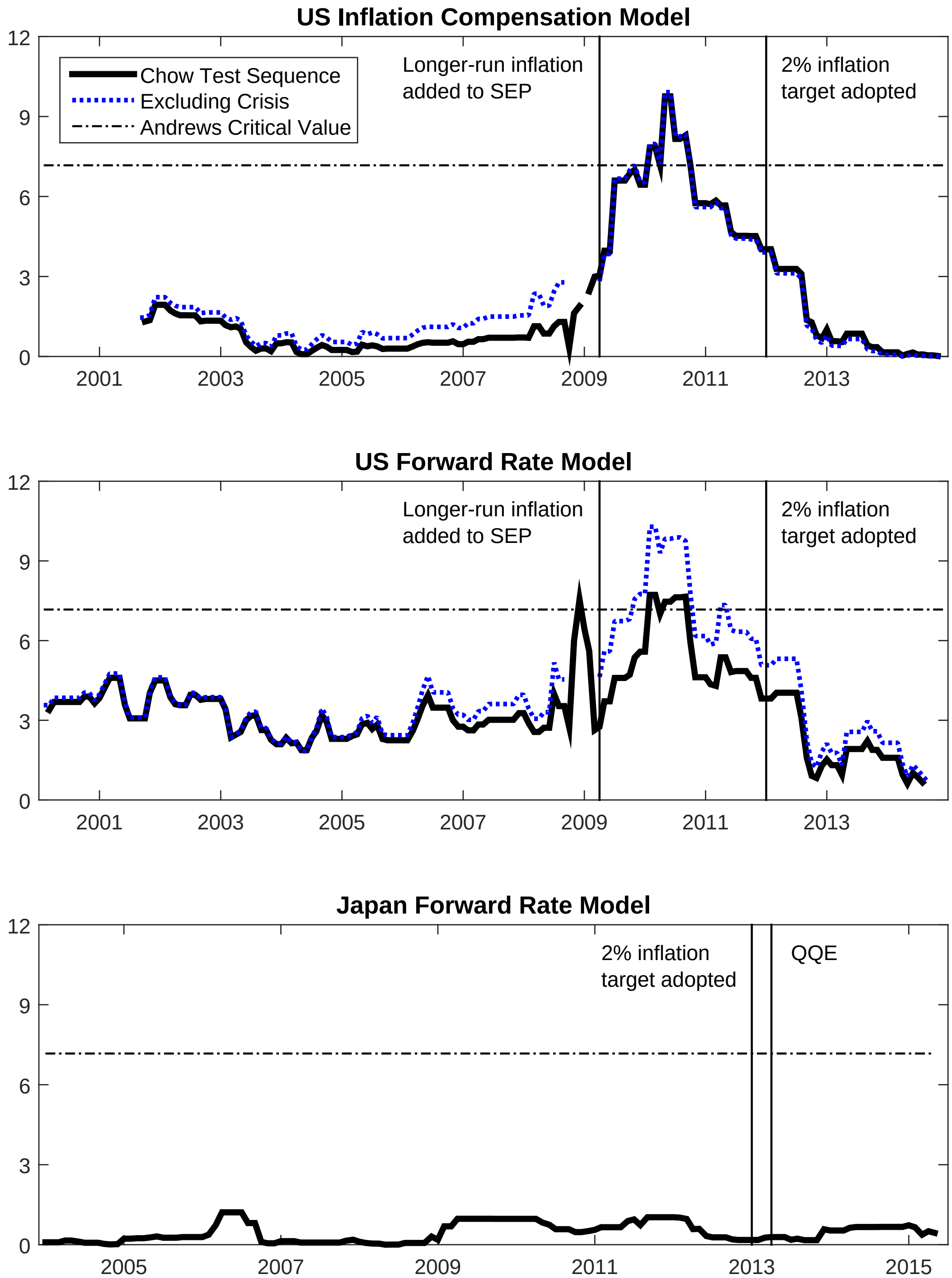
Note: Eicker-White standard errors in paranthesis. $^{*}p < 0.10$,$^{**}p < 0.05$

Table 9: Model-Implied Implications of Anchored Inflation Expectations

		Model		Model
		Drifting Expectations		Anchored Expectations
	Data	$\vartheta_\pi = 0.003$	$\vartheta_\pi = 0.01$	$\vartheta_\pi = 0.0$
Average 10-Year Term Premium	1.06	1.00	1.11	0.95
Average Slope of Yield Curve	1.43	0.88	0.96	0.85

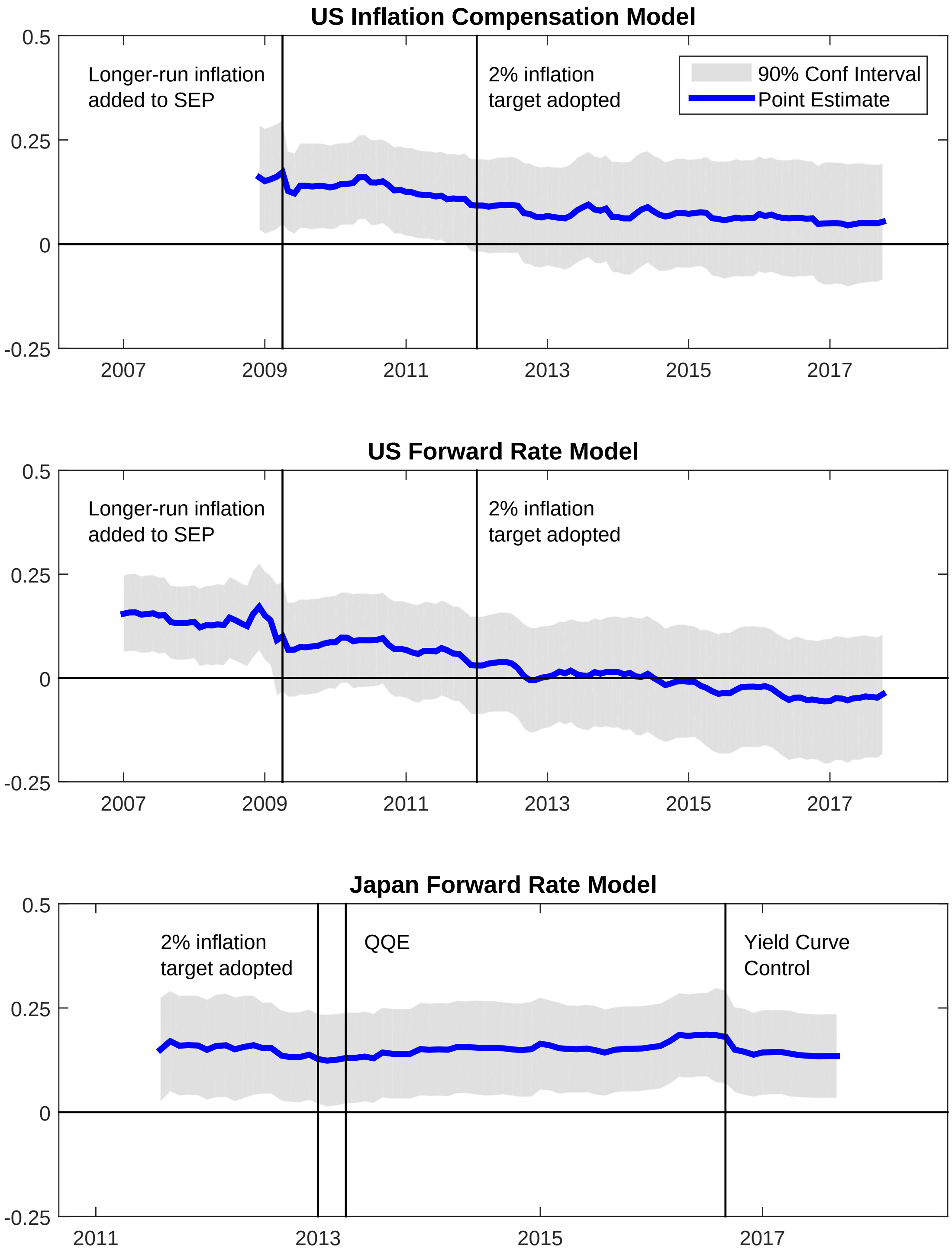
Note: We compute the model-implied moments using the model in Section III.B of Rudebusch and Swanson (2012). With the exception of ϑ_π, all other model parameters are calibrated to match their best-fit values reported in the paper. The empirical moments are from Table 3 of their paper.

Figure 1: Chow Test Sequence for Core Inflation Coefficient as a Function of Breakdate



Note: Each panel shows the sequence of Chow test statistics as a function of candidate break dates. For each model, 15% of the observations on the ends of the sample are not examined as break points. 10% critical values are obtained from Andrews (1993) for $\pi_0 = 0.15$ and $p = 1$.

Figure 2: Rolling Window Estimates of Core Inflation Coefficient



Note: Each panel shows the sequence of estimates of β as a function of time. The date on the x-axis denotes the end point of the 10-year rolling sample. The 90% confidence intervals are computed as the point estimate plus or minus 1.645 times the robust standard error.