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Did the Federal Reserve Break the Phillips Curve? Theory & Evidence of Anchoring Inflation Expectations*

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Abstract

In a macroeconomic model with drifting long-run inflation expectations, the anchoring of inflation expectations manifests in two testable predictions. First, expectations about inflation far in the future should no longer respond to news about current inflation. Second, better anchored inflation expectations weaken the relationship between unemployment and inflation, flattening the reduced-form Phillips curve. We evaluate both predictions and find that communication of a numerical inflation objective better anchored inflation expectations in the US but failed to anchor expectations in Japan. Moreover, the improved anchoring of US inflation expectations can account for much of the observed flattening of the Phillips curve. Finally, we present evidence that initial Federal Reserve communication around its longer-run inflation objective may have led inflation expectations to anchor at a level below 2 percent.

JEL Classification: E31, E52, E58

Keywords: Monetary Policy, Inflation, Inflation Targeting, Central Bank Communication, Structural Breaks, Phillips Curve

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1 Introduction

In January 2012, the Federal Open Market Committee (FOMC) adopted a longer-run target for inflation of 2 percent, which it believed was most consistent with its price stability mandate. One year later, the Bank of Japan (BOJ) followed suit and also adopted a 2 percent inflation objective. Economic theory, such as Woodford (2003), predicts that such a policy change can lead to better economic outcomes. Specifically, if a central bank successfully anchors long-run inflation expectations, then it can respond more aggressively to cyclical swings in the real economy without sacrificing its price stability objective. Indeed, the FOMC alluded to such benefits when it adopted its numerical 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."

– FOMC, January 2012 Statement on Longer-Run Goals & Policy Strategy

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 is an empirical question.

In this paper, we test whether the adoption of a numerical longer-run inflation objective better anchored inflation expectations in the United States (US) and Japan. We begin our analysis using a theoretical model to identify testable predictions that emerge after a central bank adopts a credible inflation target. In particular, we embed potentially time-varying long-run inflation expectations in a macroeconomic model with nominal rigidities and labor search frictions. Then, guided by the predictions from this model, we use both high-frequency financial market data and monthly macroeconomic data to examine the degree to which inflation expectations became better anchored after the Federal Reserve and the BOJ each adopted a numerical inflation objective.

In our theoretical model, the adoption of a credible long-run inflation target manifests in two testable predictions. First, after adopting a credible inflation target, expectations about inflation far in the future no longer respond to unexpected changes in current inflation. In contrast, if inflation expectations are not well anchored, then recent inflation developments can sway longer-term inflation expectations. Second, the anchoring of inflation expectations weakens the typical negative relationship between unemployment and inflation, thereby flattening the reduced-form Phillips curve. Models with nominal rigidities predict that demand-driven changes in the unemployment rate lead to increases in inflation. If expectations are not well anchored then these transitory increase in inflation will increase expectations for inflation far in the future. Forward-looking firms internalize this increase in inflation expectations and set higher prices today as a result. Thus, the anchoring of expectations removes these second-round inflationary effects from reductions in unemployment which results in a flatter reduced-form Phillips curve.

Using high-frequency bond market data and monthly macroeconomic data, we examine these two predictions for both the US and Japan. High-frequency evidence as well as Phillips curve estimates suggest that the FOMC's policy of communicating a longer-run numerical inflation objective achieved its stated goal of better anchoring inflation expectations. We first document that inflation expectations in the US were not well anchored prior to the adoption of a numerical inflation target. In particular, we show that far forward measures of inflation compensation from financial markets respond significantly to inflation news contained in the monthly release of the Consumer Price Index (CPI). However, not long after the FOMC started publishing individual numerical objectives for longer-run inflation in the quarterly Summary of Economic Projections (SEP) in 2009, a battery of econometric tests suggest that far forward inflation compensation ceased to respond to inflation news. While our analysis does not provide estimates of the level around which inflation expectations are anchored, we show that from 2009 to 2012, FOMC SEP projections indicated a preference for longer-run inflation between 1.5% and 2%, perhaps leading inflation expectations to anchor at a level below 2%.

In further evidence of anchoring in the US, we also find a statistical break in the reduced-form Phillips curve. Inflation became less sensitive to fluctuations in the unemployment rate after the FOMC's adoption of a longer-run inflation target. According to small-sample simulations from our theoretical model, the anchoring of inflation expectations can explain nearly all of the recent flattening of the reduced-form Phillips curve in the United States. However, small-sample simulations from our theoretical model also predict that the slope of the near-term expectations-augmented Phillips curve is invariant to changes in the degree of anchoring. Therefore, we are able to reconcile our findings with the conclusion in Coibion and Gorodnichenko (2015), who find the household expectations-augmented Phillips curve in the US has remained stable in recent decades. These results more generally help to explain why some economists and policymakers claim that the Phillips curve in the United States is "dead" while others find that it is "alive and well."

In contrast to the US experience, we find little evidence of anchoring in Japan following the 2013 adoption of a numerical inflation target by the BOJ. High-frequency evidence suggests that Japanese inflation compensation continues to drift with inflation surprises. Moreover, we fail to find evidence of instability in the reduced-form Phillips curve in Japan after the adoption of a numerical inflation objective. However, by both benchmarks, there appears to have been some progress towards anchoring: the estimated response of nominal forward rates to Japanese inflation surprises appears to have diminished and the reduced-form Phillps curve in Japan appears to be flattening. Although, statistical evidence of both of these changes is insignificant. Similar to the conclusions of De Michelis and Iacoviello (2016), our results suggest that a lack of credibility, perhaps due to past inconsistencies between the Bank of Japan's communication and its policy actions, has slowed the process of anchoring in Japan. Overall, our work underscores the importance of central-bank credibility for anchoring inflation expectations.

2 Predictions of Anchoring Inflation Expectations

We now study the implications of anchoring expectations in a theoretical model which guides our later empirical tests of whether inflation expectations are anchored in the US and Japan. The central feature of our model is the potential for long-term inflation expectations of price setters to vary over time in response to realized inflation outcomes. We embed these dynamics for long-run inflation expectations in a macroeconomic model with nominal rigidities and unemployment to shed light on the general-equilibrium consequences of anchoring inflation expectations.

2.1 Long-Term Inflation Expectations

Our specification of long-term inflation expectations allows realized inflation outcomes to change longer-term inflation expectations, which captures the notion of drifting or unanchored inflation expectations. Specifically, long-term inflation expectations evolve according to:

$$\pi_t^{LT} = \pi_{t-1}^{LT} + \delta^\pi \left(\pi_t - \pi_{t-1}^{LT}\right), \tag{1}$$

where π_t^{LT} is the long-term inflation expectation in period t and π_t is the inflation rate in period t. Empirically, π_t^{LT} is often associated with far forward measures of inflation expectations. This specification 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 US Treasury yield curve. Ireland (2007) embeds a similar mechanism in an otherwise standard macroeconomic model to study the dynamics that drive the Federal Reserve's implicit inflation target. This specification of long-term inflation expectations also mirrors the model of "trend" inflation that emerges from the forecasting equation in Stock and Watson (2007) where, assuming no stochastic volatility, agents filter unexpected changes in inflation into its permanent and transitory components.

The coefficient δ^{π} determines the degree to which long-term inflation expectations are anchored. In the extreme case, if $\delta^{\pi} = 0$, then long-term inflation expectations are fully anchored in the sense that they are invariant to realized inflation. On the other extreme, if $\delta^{\pi} > 0$, then inflation expectations are unanchored and drift with realized inflation. In the aforementioned forecasting literature, δ^{π} has a similar interpretation as it governs the signal to noise ratio placed on unanticipated or unforecastable changes in inflation.

2.2 A Macroeconomic Model with Nominal Rigidities and Labor Search Frictions

We embed our specification of long-term inflation expectations into a relatively standard model of nominal rigidities and labor market search frictions. For expositional purposes, we present our model at a high-level in the main text. We provide additional details of our theoretical model and its calibration in the Appendix. Our model combines features of previous work by Leduc and Liu (2016) and Ireland (2007). The key agents in our model are a representative household, a retail goods sector which produces differentiated products subject to nominal rigidities, an aggregation sector which aggregates the differentiated products into the final output, intermediate goods producers which hire labor in a frictional labor market, and a monetary authority which sets the short-term nominal interest rate.

2.2.1 Households

The model features a representative household populated by a continuum of worker members which maximize utility from consumption and leisure:

$$\max \mathrm{E}_t \sum_{s=0}^{\infty} a_{t+s}\, \beta^s \left\{ \log\left(C_{t+s}\right) - \chi N_{t+s} \right\}$$

where C_t denotes consumption, N_t is the fraction of employed household members, χ denotes the disutility from working, β is the household's discount factor, and a_t is an exogenous preference shock which triggers unexpected fluctuations in household demand.[1]

2.2.2 Retail Goods Producers

A continuum of firms in the monopolistically-competitive retail goods sector each produce a differentiated product $Y_t(i)$ using a homogeneous intermediate good as input. An aggregation sector purchases each intermediate good $Y_t(i)$ with price $P_t(i)$ and aggregates the differentiated retail goods into output of the final consumption good Y_t. Firm i faces a quadratic cost to adjusting its nominal price $P_t(i)$:

$$\frac{\phi_P}{2}\left[\frac{P_t(i)}{\Pi_t^{LT} P_{t-1}(i)} - 1\right]^2 Y_t$$

where $\Pi_t^{LT} = \exp(\pi_t^{LT})$ is the gross rate of long-term inflation expectations from Equation (1), ϕ_P governs the magnitude of the adjustment costs, and Y_t is output of the final output good.

2.2.3 The Labor Market

At the beginning of each period, there exist N_{t-1} employed workers, u_t unemployed workers searching for jobs, and v_t vacancies posted by firms. Matches between unemployed workers and vacancies are created using a Cobb-Douglas matching function:

$$m_t = \mu u_t^{\alpha} v_t^{1-\alpha}, \tag{2}$$

where m_t is the number of successful matches, the parameter $\alpha \in (0,1)$ denotes the elasticity of job matches with respect to the number of searching workers, and the parameter μ scales the matching efficiency. A fraction ρ of the employed workers lose their jobs each period. Thus, the number of workers who survive the job separation is $(1-\rho)N_{t-1}$. At the same time, m_t new matches are formed.

Following Blanchard and Galí (2010), we assume that new hires start working in the period they are hired. Thus, aggregate employment in period t evolves according to:

$$N_t = (1-\rho)N_{t-1} + m_t. \tag{3}$$

[1] These demand shocks are the only exogenous shocks in our model. The inclusion of other shocks is not necessary to generate our key findings. McLeay and Tenreyro (2020) highlight the difficulties that arise in identifying the Phillips curve when cost-push shocks are the driving force of economic variation.

We assume full participation and define the unemployment rate as the fraction of the population who are left without a job after hiring takes place. Thus, we can write the unemployment rate as follows:

$$U_t = u_t - m_t = 1 - N_t. \tag{4}$$

2.2.4 Intermediate Goods Producers

Each intermediate good firm produces a homogenous intermediate good and hires at most one worker subject to search and matching frictions in the labor market. If a firm finds a match, it obtains a flow profit in the current period after paying the worker. In the next period, the match may survive with probability $1-\rho$ or dissolve with probability ρ. If the match breaks down, the firm posts a new job vacancy at a fixed cost κ units of the final good with the value V_{t+1}. The following Bellman equation captures the value of the firm:

$$J_t^F = q_t - W_t + \mathbb{E}_t \left\{ \left(\beta \frac{\lambda_{t+1}}{\lambda_t} \right) \left((1-\rho) J_{t+1}^F + \rho V_{t+1} \right) \right\}. \tag{5}$$

where q_t denotes the relative price of the intermediate good, W_t denotes the real wage, and λ_t is the representative household's marginal utility from consumption.

Firms and workers Nash bargain over wages. However, following Hall (2005) and Blanchard and Galí (2010), we assume actual wages adjust slowly to changing economic conditions:

$$W_t = W_{t-1}^{\gamma} \left(W_t^N \right)^{1-\gamma} \tag{6}$$

where W_t^N is the wage under Nash bargaining and $\gamma \in (0,1)$ represents the degree of real wage rigidity.

2.2.5 Monetary Policy

The central bank in the model sets its short-term nominal policy rate R_t to minimize fluctuations in inflation in deviation from its long-term expectations and output growth, smoothing changes in interest rates over time:

$$\log \Big(R_t \Big) = \log \Big(R_{t-1} \Big) + \phi_\pi \log \Big(\Pi_t / \Pi_t^{LT} \Big) + \phi_y \log \Big(Y_t / Y_{t-1} \Big), \tag{7}$$

where ϕ_π and ϕ_y denote the central banks response to inflation deviations and changes in output growth.[2]

[2] Our random walk specification for long-term inflation expectations implies that inflation Π_t is nonstationary but detrended inflation Π_t/Π_t^{LT} is stationary. Therefore, we need a coefficient of one on lagged policy rates in our policy rule to write the policy rule in a stationary form. See the Appendix for more details.

2.3 Theoretical Predictions of Anchoring Inflation Expectations

We now use our theoretical model under both a drifting and anchored calibration of long-term inflation expectations to illustrate the implications of adopting a credible inflation target. After writing our model in stationary form, which is necessary due to the random-walk specification for long-term inflation expectations, we solve our model using a first-order approximation around the deterministic steady state. Table A.1 in the Appendix contains the calibrated values for the model parameters. Since our model shares many features with the models of Ireland (2007) and Leduc and Liu (2016), we calibrate many of the parameters to values used in those papers. However, there are two important exceptions. We calibrate the degree of anchoring of inflation expectations δ^{π} to align with our high-frequency estimates before the United States formally adopted its inflation objective. We then set $\delta^{\pi} = 0$ to simulate the dynamics under anchored expectations. We also calibrate the degree of nominal rigidities ϕ_P to reproduce the observed reduced-form Phillips curve prior to the adoption of the inflation target. We discuss the empirical support for these calibration choices in detail later in the paper.

Two testable predictions emerge from our general-equilibrium model once inflation expectations become anchored. Figure 1 illustrates these predictions by comparing the impulse responses to a one standard deviation aggregate demand (preference) shock under both drifting and anchored long-term inflation expectations. The dashed-red lines show the impulse responses under the drifting inflation expectations regime and the solid-blue lines show the impulse responses under the anchored inflation expectations regime.

The first model prediction is that long-run inflation expectations cease to comove with unanticipated changes in realized inflation under anchored inflation expecations. The top row of Figure 1 shows that the unanticipated increase in inflation induced by the demand shock spills over to long-run inflation expectations when inflation expectations drift. However, once anchored, the surprise increase in inflation following the demand shock no longer influences long-run inflation expectations. Intuitively, once inflation expectations are anchored, realized inflation no longer informs agent's views of where the central bank will steer inflation over the long run. Instead, when a credible inflation target is adopted, long-run inflation is pinned down solely by the central bank's communication.

The second model prediction is that the adoption of a credible inflation target mutes the response of inflation to changes in unemployment thus weakening the reduced-form

Phillips curve. The second row of Figure 1 shows that, for roughly the same reduction in the unemployment rate, the model generates a larger and more persistent increase in inflation when inflation expectations drift with realized inflation. In contrast, the response of inflation is more muted once inflation expectations are anchored. Intuitively, the theoretical model predicts that the successful anchoring of expectations flattens the slope of the reduced-form Phillips curve by removing the spillover effects from inflation into long-term inflation expectations for a given reduction in the unemployment rate.

2.4 Testing of the Anchored Inflation Hypothesis

We now translate these model predictions into empirically testable hypotheses using a mix of high-frequency financial market data and monthly macroeconomic data.

2.4.1 High-Frequency Evidence

The first model prediction rests on estimating δ^π, which measures the degree to which long-term inflation expectations respond to realized inflation in Equation (1). However, a simple regression of long-term inflation expectations on current inflation is likely to yield biased estimates of δ^π. In particular, Equation (1) is part of a larger macroeconomic model with an expectations-augmented Phillips curve which also links current inflation to long-term inflation expectations. Therefore, the simultaneity between long-term inflation expectations and actual inflation when $\delta^\pi > 0$ makes the direct estimation of Equation (1) problematic.

However, an algebraic manipulation of Equation (1) allows us to directly estimate δ^π using a high-frequency 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:

$$\pi_t^{LT} - \mathbb{E}_{t-1}\pi_t^{LT} = \delta^\pi(\pi_t - \mathbb{E}_{t-1}\pi_t), \tag{8}$$

where the right-hand side captures the news about current inflation revealed between time $t-1$ and t and the coefficient δ^π governs how that inflation news affects long-term inflation expectations. Equation (8) suggests that we can estimate δ^π using a high-frequency event-study approach of regressing the change in far forward inflation compensation on the news or unexpected component of the monthly CPI report. If we find that $\delta^\pi > 0$ such that forward measures of inflation compensation respond significantly to news about current inflation, then this would suggest that inflation expectations are unanchored. In the following sections, we formally test this prediction for both the US and Japan.

2.4.2 Implications for the Reduced-Form Phillips Curve

The second model prediction rests on estimating the correlation between inflation and unemployment. This aspect of our empirical analysis relates our work to a large literature which seeks to estimate Phillips curve relationships by regressing inflation on the unemployment rate. However, this literature is mired in debate around the appropriate measure of inflation expectations to include in the Phillips curve and issues of identification (Mavroeidis, Plagborg-Møller and Stock, 2014; Coibion, Gorodnichenko and Kamdar, 2018; McLeay and Tenreyro, 2020). We will return to some of these issues when we contrast the breakdown of reduced-form Phillips curves with potential stability of structural Phillips curve estimates amid the anchoring of long-run inflation expectations. For now, we simply seek to show that, despite these identification issues, our theoretical model suggests that simple ordinary least squares regressions of inflation on the unemployment rate can detect the weakening of the reduced-form Phillips curve relationship that is induced by anchoring inflation expectations, as illustrated in the impulse responses in Figure 1.

Figure 2 shows scatter plots and estimated regression lines for the relationship between the unemployment rate and inflation implied by our model under different degrees of anchoring. We begin by simulating data from the model for 2000 periods under the drifting inflation expectations calibration of $\delta^{\pi} > 0$, plotted by the red dots in Figure 2. We observe a tightly-estimated and steep reduced-form relationship between unemployment and inflation in the simulated data when long-term expectations drift with current inflation. We then simulate data from the model for 2000 periods when inflation expectations are well anchored ($\delta^{\pi} = 0$). In contrast to the unanchored Phillips curve relationship, the blue squares in Figure 2 show a weaker relationship between unemployment and inflation with more error around the estimated regression line. We show in the coming analysis that, even in a small sample of the size we consider in our empirical work, an econometrician would be able to detect a statistical break in this reduced-form Phillips curve regression following the anchoring of expectations.

Guided by these two model predictions, we now formally test the degree to which inflation expectations are anchored in the US and Japan. For each country, we examine: (1) changes in the high-frequency sensitivity of far forward inflation compensation to inflation surprises and (2) changes in the reduced form Phillips curve relationship — the slope coefficient when inflation is regressed on the unemployment rate — after the adoption of a numerical inflation target.

3 Are Inflation Expectations Anchored in the US?

Evidence from high-frequency event studies as well as Phillips curve regressions both suggest that inflation expectations became better anchored after the Federal Reserve began communicating a numerical inflation target. Beyond implementing the empirical tests for anchoring, as prescribed by our model predictions, we also shed light on the source of the anchoring as well as the quantitative implications of anchoring expectations for the US Phillips curve.

3.1 Inflation Compensation & Inflation News in the US

We begin our empirical analysis by examining the high-frequency sensitivity of US inflation compensation to unexpected news about inflation. Our key question is: Did the coefficient δ^{π} change after the FOMC adopted an explicit inflation objective? To measure δ^{π}, we estimate Equation (8) 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. Focusing on forward measures of inflation compensation cleanses any direct effect that current inflation has on average inflation over the next decade.

For our measure of $\pi_t - \mathbb{E}_{t-1}\pi_t$ in Equation (8), we use data surprises emanating from the release of monthly CPI reports.[3] 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.[4] For the US, Bloomberg provides 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. Our sample period is limited by the availability of data on inflation surprises, which starts in 1997 for the United States.

Using the core CPI surprise and the surprise associated with the food and energy components, we estimate the following event-study regression to measure δ^{π}, the sensitivity of long-term inflation compensation to news about inflation,

$$\Delta\pi_t^{LT} = \delta^0 + \delta^{\pi}\pi_t^{core} + \delta^{fe}\pi_t^{fe} + \varepsilon_t, \tag{9}$$

[3] 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 (8).

[4] Bloomberg also maintains data on the actual value of π_t in the CPI release (i.e. not the revised value)

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. We will refer to this specification as the *inflation compensation model.* We estimate Equation (9) using ordinary least squares where each observation corresponds to a given CPI release. We apply several different statistical methods to detect if the underlying relationship between inflation surprises and inflation compensation changed after the FOMC's adoption of an explicit inflation target. First, we estimate Equation (9) across two distinct sample periods. In the first sample, we examine the January 1999 – December 2011 period, which is prior to the inflation target adoption. Then, we examine the January 2012 – December 2019 period following the policy change.

We find that inflation compensation responds less to economic news about inflation after 2012, suggesting better anchored inflation expectations following the adoption of the inflation target. The first two columns of Table 1 show the estimated coefficients from Equation (9) across the two sample periods. Prior to January 2012, a positive core CPI surprise led 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. The finding that far-forward TIPS-based inflation compensation drifts with core CPI surprises in our early sample echos the findings in Beechey, Johannsen and Levin (2011) and Bauer (2015), both of which conclude their estimation in 2007. However, after the FOMC formally adopting its inflation target, the coefficient on the core CPI surprise falls and becomes 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 δ^π in 2012. The post-January 2012 dummy variable that interacts with the core inflation surprise is negative and statistically significant, which suggests a statistically significant reduction in the sensitivity of longer-term inflation compensation to news about core inflation after the FOMC adopted its formal inflation target.[5]

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 are statistically insignificant. Two intuitive reasons support this empirical finding. First, to the extent that food and energy price fluctuations are short-lived and often reverse in the coming months, we would expect them not to have an effect on longer-term inflation expectations. Second, thanks to vibrant

[5]In the Appendix, we provide additional evidence that shows that the break in the estimate of δ^π appears to reflect a change in the reaction of inflation expectations to CPI surprises, rather than a change in the nature of CPI surprises during the post-January 2012 period.

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.

3.1.1 Testing for a Structural Break at an Unknown Date

Rather than imposing a break in 2012 in the relationship between far forward inflation compensation and inflation surprises, we now test for a structural break at an unknown date. Probing the timing of the break allows us to provide some further interpretation of the source of the break in the δ^{π} coefficient. If a change in δ^{π} reflected a better anchoring of inflation expectations, we would expect the estimated break date to follow a change in US monetary policy. If instead the estimated break date is not supported by 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 δ^{π}, but not any of the other parameters in the regression model. Table 2 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 (9). These break tests suggest that the relationship between inflation compensation and inflation news changed in May 2010. The candidate break is significant at nearly the 5% level for both tests, indicating strong statistical evidence of a change in δ^{π}. There is no evidence of a break in any of the other regression parameters, including the variance of the regression residual.[6] The solid black line in Panel A of Figure 3 plots the time series of Chow test statistics for a break in δ^{π}. 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.[7]

3.1.2 The Summary of Economic Projections as Source of Anchoring

The estimated break date occurs before the Federal Reserve formally adopted a 2 percent inflation target in January of 2012, but shortly after FOMC participants began to pub-

[6]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.

[7]More formally, if we split the sample into two subsamples 1999-2010 and 2010-2019, 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 April of 2014. 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 1999-2014.

lish quarterly projections for “longer-run” inflation. In January of 2009, the Summary of 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 target for the FOMC’s longer-term inflation objective.

Reviewing historical FOMC transcripts corroborates this view that the numerical SEP projections for longer-run inflation were intended to better anchor inflation expectations. In a January 2009 conference call, the FOMC discussed adding a projection for longer-run inflation to the SEP. In that call, then San Francisco Fed President Janet Yellen noted that:

“Greater transparency about how we think the future will likely unfold could help anchor inflationary expectations . . . But our existing FOMC projections, which have the three-year forecast horizon, obviously aren’t up to the task . . . The obvious solution to this problem is to provide economic projections with a longer horizon . . . ”

– Janet Yellen, January 16, 2009 FOMC Conference Call

The evident timing of the break we identify in the relationship between inflation surprises and inflation compensation largely supports her claim. In particular, our break date suggests that the modifications to the SEP — and not necessarily the adoption of a formal 2-percent target in 2012 — played an instrumental role in helping the FOMC anchor inflation expectations. On the same 2009 conference call, Yellen hypothesized this may well be the case. In particular, she saw the addition of longer-run projections as largely accomplishing the desired degree of anchoring without needing to adopt a Committee-wide longer-run numerical target. Specifically, she stated, “I don’t expect the associated gains from transparency and better anchoring of inflation expectations from the enunciation of an explicit numerical inflation objective to be a lot larger than those that we would achieve just from extending the forecast horizon.”

While the longer-run SEP projections for inflation appeared to have been the catalyst for anchoring inflation expectations, these projections were dispersed and centered below 2 percent prior to 2012. Figure 4 shows the range of FOMC participant’s projections for longer-run inflation spanned 1.5 percent to 2.1 percent during the 2009-2012 period. Moreover, the central tendency and midpoint of FOMC projections were below 2 percent for three years prior to the FOMC's adoption of a 2 longer-run inflation target. This distribution of longer-run inflation projections from 2009-2012 highlights an important implication of our results. The downward skew of the SEP projections for longer-run inflation over this period may have caused inflation expectations to become initially anchored below 2 percent. This revelation could help to explain why US PCE inflation has been stable but below 2 percent for almost the entire time since the FOMC's 2012 adoption of a numerical inflation target.[8]

3.1.3 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 examine the robustness of this finding to 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 of these alternative specifications, we continue to find evidence that nominal compensation became unresponsive to inflation news after the FOMC communicated an 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.[9] Our baseline model also uses the weight of core goods 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 estimate the following alternative regression model around CPI releases:

$$\Delta y_t^{LT} = \delta^0 + \delta^\pi \pi_t^{core} + \delta^f \pi_t^{food} + \delta^e \pi_t^{energy} + \varepsilon_t, \tag{10}$$

where Δy_t^{LT} is the one-day change in the one-year, nine-year forward rate and π_t^{food} and π_t^{energy} are the one-day percent changes in the Goldman Sachs agricultural and energy price indexes, respectively.[10] We refer to this as the *forward rate model.*

[8] Shapiro and Wilson (2019) perform text analysis of FOMC communication to estimate the FOMC's objective function and similarly find that the FOMC's implicit objective for inflation is below 2 percent.

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

[10] We calculate nominal forward rates from the yield on constant maturity zero coupon bond yields as described in Gürkaynak, Levin and Swanson (2010).

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 is 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.

Using this alternative forward rate model, the regression results in Table 3 show a decline in the response of inflation compensation to inflation news following the adoption of the inflation target.[11] Before 2012, nominal compensation significantly comoved with inflation surprises. However, after 2012, nominal compensation became unresponsive.[12] The general robustness of our findings using the forward rate model is important as we move to our analysis of the BOJ's adoption of a numerical inflation target. For Japan, 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 (9). 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 in Panel B of Figure 3 plots the Chow test sequence for δ^{π} 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. Both the 2008 and 2010 breaks fail to exceed the 10% critical value of the Andrews-Quandt test in the second panel of Figure 3. 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.

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

[12] While the Chow test statistic for a break after 2012 is slightly below the critical value, the p-value of the Chow test is 0.1003, indicating that our findings are generally robust. This specification reveals a serial correlation in the errors which, once corrected, drives the p-value on the Chow test below 0.01.

If we drop the precipice of the global financial crisis, we find evidence indicating the presence of a single structural break in early 2010. For our baseline inflation compensation model, Table B.2 in the Appendix shows that we estimate the exact same break date of May 2010 for the core inflation coefficient if we drop the fourth quarter of 2008 and first quarter of 2009 from the estimation. The blue dotted lines in Panels A and B of Figure 3 plot the time series of the Chow statistics for samples that exclude the financial crisis. For both the inflation compensation and forward rate 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 B.3 in the Appendix shows that the estimated break date for the forward rate model is February of 2010 and that break is estimated to be statistically significant using the Andrews-Quandt test and the Andrews-Ploberger test. 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 indicate a similar decline in δ^{π} 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 change in the coefficients over time. Panel A of Figure 5 illustrates the time variation in δ^{π} from the inflation compensation regression model specified in Equation (9) using 10-year rolling samples. We observe the same pattern of structural change as our previous findings. Early in the sample, prior to 2012, δ^{π} is estimated to be statistically significant and positive. However, the point estimate of δ^{π} begins to decline in 2010 and falls to values not different from zero by 2012. Panel B of Figure 5 shows similar time variation in δ^{π} as estimated from the forward rate regression model in Equation (10). The point estimate of δ^{π} from the forward rate model is positive and significant before 2009, and thereafter begins a steady descent towards zero. The results of these alternative specifications provide further evidence that the FOMC's decision to communicate a numerical inflation objective helped better anchor US inflation expectations.

3.2 Phillips Curve Estimates in the United States

In our first testable prediction, high-frequency empirical evidence suggests that communicating an inflation objective coincided with a better anchoring of US inflation expectations. We now examine our second testable implication of anchoring inflation expectations: An anchoring of expectations weakens the relationship between inflation and unemployment and thus flattens the reduced-form Phillips curve. Empirically examining this prediction underscores the potential role monetary policy plays in shaping the relationship between inflation and unemployment and sheds light on the source of conflicting evidence of Phillips curve instability in the recent macro literature.

The reduced-form Phillips curve in the US appears to have significantly flattened over the last two decades. Using monthly data on inflation and unemployment, the Panel A of Figure 6 shows a scatter plot of the unemployment rate versus year-over-year core inflation as measured by the consumer price index excluding food and energy. Over the 1999-2019 period, the same sample we use in our high-frequency analysis, we observe a visible reduction in the slope of the reduced-form Phillips curve in the 2012-2019 sample.

Regression analysis confirms a meaningful breakdown in the relationship between inflation and the unemployment rate after January 2012. Table 5 contains the results from a regression of inflation on a constant and the unemployment rate before and after January 2012. We highlight three key findings. First, the slope of the Phillips curve has flattened since 2012. Prior to 2012, Table 5 illustrates a statistically significant and downward-sloping relationship between inflation and unemployment, with a Phillips curve slope of roughly -0.19. After 2011, however, the slope of the Phillips curve becomes much flatter, declining in magnitude to less than -0.05. Second, a Chow (1960) test for a break in January 2012 confirms that this flattening is statistically significant. And, third, Table 5 reveals an increase in unexplained fluctuations in inflation after 2012, as evident through a lower R^2. In light of our theoretical model, the breakdown in the reduced-form Phillips curve offers further evidence that the Federal Reserve's 2012 adoption of a numerical inflation target better anchored inflation expectations in the United States. This evidence builds on work by Jorgensen and Lansing (2019) and McLeay and Tenreyro (2020) that also studies the role that shifts in monetary policy may play in altering the reduced-form Phillips curve relationship.[13]

[13]Our focus on the role of recent changes in FOMC communication together with the link we draw between the timing of those communications and changes in the reduced-form Phillips curve distinguishes our contribution. For instance, Jorgensen and Lansing (2019) emphasize the effect that the Volcker regime change induced on the Phillips curve relationship. Meanwhile, McLeay and Tenreyro (2020) focus on how an optimal targeting rule can mask the relationship between unemployment and inflation with less focus on the source or timing of any specific break in Federal Reserve policy.

3.2.1 Quantitatively Accounting for the Flattening of the US Phillip Curve

The left three columns in Table 5 illustrate a large reduction of the sensitivity of inflation to the unemployment rate since 2012. How much of this flattening of the reduced-form Phillips curve can be explained by the anchoring of inflation expectations as opposed to other changes in the economy? To answer this question, we use our theoretical model to quantitatively assess how much the slope of the reduced-form Phillips curve changes given the observed changes in the degree of anchoring of inflation expectations.

To examine the quantitative predictions of the model, we replicate our empirical Phillips curve exercise using simulated data from our theoretical model both before and after the adoption of the inflation objective. To conduct this exercise, we use our high-frequency estimates from Table 1 to generate two different calibrations for our theoretical model. In the first calibration, we generate a drifting inflation target economy by setting δ^{π} in Equation (1) equal to 0.15, our high-frequency coefficient on core CPI from Table 1 over the 1999-2011 period. In the second calibration, we calibrate δ^{π} using the estimated value over the 2012-2019 sample period and hold all other model parameters fixed.

To generate the model-implied Phillips curves, we first simulate the model with the drifting inflation target specification for 156 periods, the same length as our empirical Phillips curve specification in Column 1 of Table 5. Then, in period 157, we assume that inflation expectations become anchored and continue to simulate the model for another 96 periods under the anchored inflation expectations calibration. We then estimate reduced-form Phillips curve regressions on this simulated model data both before and after anchoring and test for a structural break in the slope coefficient after anchoring.[14] We repeat this exercise 1000 times and generate small-sample bootstrapped confidence intervals. To facilitate comparison with our empirical evidence, we calibrate the degree of nominal rigidity ϕ_P such that our model with the drifting inflation target specification generates the same average reduced-form Phillips curve slope (-0.19) that we observe in the data during the 1999-2011 period.[15]

[14] We also simulate a burn-in sample prior to conducting this exercise in order to ensure that our conclusions are not driven by initial conditions. As in our empirical evidence, we regress year-over-year inflation onto a constant and the unemployment rate in our model-implied Phillips curves. However, we find similar results if we instead use annualized quarterly inflation rather than year-over-year inflation in our simulated Phillips curve exercise.

[15] Our calibrated value of $\phi_P = 170$ is very close to the value estimated by Ireland (2003) and is well within the accepted range of calibrated values from the literature.

The far right columns of Table 5 shows the resulting regression coefficients and their associated 95% bootstrapped standard errors. These model simulations reveal that the anchoring of expectations in the model qualitatively reproduces the three empirical findings that we observed in the reduced-form Phillips curve in the data following the adoption of a formal inflation target: (1) a flattening in the slope coefficient, (2) a statistical break in the slope coefficient, and (3) a decline in the regression R^2 during the post-anchoring period. Importantly, these findings confirm that, even in a small sample, an econometrician would be able to detect a breakdown in the reduced-form Phillips curve after the adoption of an inflation target.

Quantitatively, our theoretical model simulations suggest that the anchoring of inflation expectations in the United States can explain nearly all of the observed flattening in the reduced-form Phillips curve. In the model simulations, the slope of the reduced-form Phillips curve shrinks from -0.19 to -0.08. While the flattening of the slope coefficient in the data (0.14) is a bit larger than the point estimate of what the model predicts would occur solely due to anchoring (0.11), the 95% model-implied confidence interval of the post-anchoring flattening is (0.14, 0.02), which contains the empirical estimate of 0.14. Therefore, at the 5% level, we cannot reject the null hypothesis that the better of anchoring of inflation expectations explains the observed changes in the US reduced-form Phillips curve.

3.2.2 Slope of the Structural Versus Reduced-Form Phillips Curve

In sharp contrast to our results, Coibion and Gorodnichenko (2015) argue that the Phillips curve has remained stable in recent decades once it is augmented with household's near-term inflation expectations. In this section, we aim to reconcile our results with their findings. In particular, we show that their specification exhibits stability amid the anchoring of inflation expectations because it proxies well the underlying structural Phillips curve in our model, which is also invariant to the degree of anchoring. The analysis in this section also helps to reconcile our findings with the conclusion in Coibion and Gorodnichenko (2015) that the stability of their expectations-augmented Phillips curve results from unanchored rather than anchored inflation expectations.

The slope of the underlying structural Phillips curve in our model is invariant to the degree of anchoring. For example, in our theoretical model, taking a first-order approximation to the optimal pricing decision of retail firms implies the following relationship:

$$\pi_t - \pi_t^{LT} = \beta \, \mathrm{E}_t \left\{ \pi_{t+1} - \pi_{t+1}^{LT} \right\} + \Psi \, \Xi_t \tag{11}$$

where Ξ_t denotes firm marginal costs and Ψ is the slope of the structural Phillips curve. Importantly, Ψ is invariant to changes in the degree of anchoring. Ravenna and Walsh (2011) further show that marginal costs in a model like ours can be written as a function of the unemployment rate. Therefore, the instability we document in the reduced-form Phillips curve results partly from the failure to account for firm inflation expectations.

We now demonstrate that a Phillips curve augmented with near-term firm inflation expectations is stable amid changes in the degree of anchoring. We illustrate this stability using small-sample simulations from our theoretical model. Motivated by the Coibion and Gorodnichenko (2015) specification, we regress year-over-year inflation less 1-year ahead inflation expectations on the unemployment rate using model-generated data. The far right columns of Table 6 show the resulting regression coefficients when we estimate this expectations-augmented Phillips curve using a sample size of 250 periods with a break in the degree of anchoring in period 157 and repeat this exercise 1000 times. Despite estimating a miss-specified Phillips curve relative to the actual data generating process in Equation (11), we find that the slope of the near-term firm inflation expectations augmented Phillips curve specification remains stable amid the anchoring of expectations. In particular, even in a controlled experiment where there is a known break in the degree of anchoring in period 157, the Chow (1960) test for a break shows no evidence of instability.

These model simulation results suggest that the Coibion and Gorodnichenko (2015) Phillips curve specification, which replaces firm expectations with household expectations, should remain stable amid changes in the degree of anchoring.[16] Coibion and Gorodnichenko (2015) use quarterly data and end their analysis in 2013, close to the FOMC's adoption of a numerical inflation target in 2012. Therefore, we now verify that the monthly variant of the household-inflation-expectations augmented Phillips curve has indeed remained stable since 2012. Using the same sample periods used in our reduced-form Phillips curve regressions, we estimate expectations-augmented Phillips curves with year-over-year inflation inflation less 1-year ahead household inflation expectations (measured by the University of Michigan Survey of Consumers) as the dependent variable. The far left columns of Table 6 reveals no evidence of a change in the underlying slope of the Coibion and Gorodnichenko (2015) Phillips curve in January 2012.[17]

[16] Coibion and Gorodnichenko (2015) argue that consumer inflation expectations are linked most closely with the inflation expectations of firms, the actual price and wage setters in the economy.

[17] This stability in the semi-structural version of the Phillips curve also suggests that the instability we document in the reduced-form Phillips curve is not merely coincidental with a change in the slope of the

Taken together, our model simulation and empirical results in this section explain why some researchers and policymakers claim that the Phillips curve in the United States is "dead" while others find that it is "alive and well." For example, Erceg et al. (2018) illustrate that the slope of the Phillips curve has flattened significantly over the past two decades in a backward-looking reduced-form model. In contrast, Coibion and Gorodnichenko (2015) conclude that, if we account for changes in inflation expectations, then the Phillips curve slope has remained stable. Our results help to resolve this tension in the previous literature.

The dichotomy between the reduced-form and expectations-augmented Phillips curves also helps to reconcile our finding that inflation expectations in the US are better anchored since 2012 with the conclusion in Coibion and Gorodnichenko (2015) that inflation expectations are largely unanchored. In particular, simulations from our model show that the anchoring of long-term inflation expectations does not eliminate fluctuations in 1-year ahead inflation expectations. Therefore, while our model results offer support for the Coibion and Gorodnichenko (2015) specification as a stable description of inflation dynamics, we can not conclude from this stability that fluctuations in 1-year ahead inflation expectations are inconsistent with the anchoring of longer-run inflation expectations that central banks often pursue.

4 Are Inflation Expectations Anchored in Japan?

One year after the FOMC formally adopted its longer-run inflation target of 2 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. In January 2013, the BOJ set the "price stability target" at 2 percent in terms of the year-on-year rate of change in the consumer price index (CPI) with the goal of achieving this target at the earliest possible time. 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 and introduced an inflation-overshooting commitment, whereby

underlying structural Phillips curve. Therefore, this stability offers further evidence in favor of our interpretation that the instability in the reduced-form Phillips curve reflects the better anchoring of inflation expectations in the US.

the BOJ expands the monetary base until inflation exceeds 2 percent. 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.

4.1 Inflation Compensation & Inflation News in Japan

Guided by the predictions of our theoretical model in Section 2, we first examine high-frequency changes in far forward measures of nominal compensation to news about current inflation in Japan. Specifically, we look for evidence of parameter instability using our forward rate model:

$$\Delta y_t^{LT} = \delta^0 + \delta^\pi \pi_t^{core} + \delta^f \pi_t^{food} + \varepsilon_t. \tag{12}$$

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.[18] Data on inflation surprises in Japan begin in 2001, so we estimate our forward rate model over the 2001-2019 sample period.

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 US, we initially impose a break in the regression relationship in January of 2013 after the election of Shinzo Abe. Table 7 shows the split-sample regression estimates over the 2001-2012 sample period and 2013-2019 sample periods. In both subsamples, we observe a positive and statistically significant coefficient δ^π on the core inflation surprise, suggesting far forward nominal compensation continues to drift with news about current inflation. However, the point estimate of δ^π 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 δ^π.

Unlike our findings for the United States, tests for a structural break at an unknown date indicate *stability* in the regression model in Equation (12) over time. The Andrews-Quandt

[18] As we discussed previously, we lack data on real (inflation-indexed) bonds and the knowledge about the weight of core components in the CPI basket for Japan. Thus, we cannot estimate our preferred inflation compensation model specified in Equation (9). However, we can estimate the forward rate model we used in Section 3.1.3. For Japan, Bloomberg surveys and reports the year-over-year percent change in the core CPI inflation as opposed to the month-over-month percent change we used for the United States. Although this may have implications for interpreting the magnitude of δ^π, the scaling does not affect hypothesis tests against the null of $\delta^\pi = 0$.

and Andrews-Ploberger tests indicate no evidence of significant time variation in δ^π. Panel C of Figure 3 shows the Chow test sequence over candidate breakdates. The time series of Chow tests has no well defined peaks any where near the 10% critical value for a structural break. Rolling-window regressions also support these findings of a lack of structural change. Panel C of Figure 5 illustrates estimates of the sensitivity of nominal forward rates to core inflation surprises over 10-year rolling windows advanced one month at a time. As with the split-sample estimates and the break tests, the time series of estimated δ^π 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 markets – remain unanchored in Japan.

4.2 Phillips Curve Estimates in Japan

Corroborating our high-frequency evidence that inflation expectations remain unchanged despite adoption of an inflation target, we find evidence of a *stable* reduced-form Phillips curve in Japan. Panel B of Figure 6 and Table 8 illustrate the corresponding scatter plot and regression results for inflation and unemployment for Japan over the 2001-2019 period (the same period we use in our high-frequency empirical work in Table 7). Unlike the United States, we find no statistically significant evidence of a break in the reduced-form relationship between inflation and unemployment. However, we observe some reduction in the magnitude of the reduced-form Phillips curve and we see more unexplained variation in inflation during the 2013-2019. Consistent with the reduced estimate of δ^π in the post-2013 period, the insignificant flattening of the reduced-form Phillips curve in the post-2013 period might also indicate initial signs of anchoring in Japan. However, at this time, the combined evidence from high-frequency nominal compensation and Phillips curve estimates suggest that adopting an inflation objective has yet to fully anchor inflation expectations in Japan.

5 Conclusions & Implications for Central Bank Credibility

Over the past ten years, Gürkaynak et al. (2007), Gürkaynak, Levin and Swanson (2010) and Beechey, Johannsen and Levin (2011) have conducted 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 Euro Area, Canada, the UK and Sweden than in the United States because, at that time, the United States had not yet adopted a numerical long-term inflation objective. Our paper builds on this influential work in several dimensions. To begin with, our results provide external validity to their analysis and conclusions. Indeed, we find that inflation expectations became better anchored as the FOMC began communicating a numerical objective for longer-run inflation. Specifically, along the lines of Gürkaynak, Levin and Swanson (2010), we find a reduction in the sensitivity of far forward measures of nominal compensation to current inflation surprises in the United States after 2012.

Moreover, we develop a macro model with drifting long-run inflation expectations, nominal rigidities, and labor search frictions to identify a second testable prediction: the anchoring of inflation expectations flattens the slope of the reduced-form Phillips curve. Using US data on inflation and the unemployment rate, we then provide empirical evidence consistent with this prediction which further supports our claim that inflation expectations became better anchored after the FOMC adopted a numerical long-term inflation objective. In contrast, we show that the slope of the expectations-augmented Phillips curve, similar to the type proposed in Coibion and Gorodnichenko (2015), is invariant to the degree of anchoring. Therefore, our results help to resolve the tension between some research which claims the Phillips curve relationship has disappeared in recent US data with other research which argues that a stable relationship exists between inflation, inflation expectations, and the unemployment.

However, our analysis of Japan underscores the importance of central bank credibility in better anchoring 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. Moreover, there is no significant evidence of flattening of the reduced-form Phillips curve in Japan. To our knowledge, our paper presents the first high-frequency financial market evidence demonstrating that inflation expectations may remain unanchored in an advanced economy even after adopting a numerical inflation target. These results for Japan underscore 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 inconsistencies between policymakers' words and actions may slow or prevent inflation expectations from becoming anchored. The varied experiences of the US and Japan illustrate the practical value of maintaining central bank credibility. This credibility may be particularly valuable should central banks rely more heavily on guiding expectations through communication during more frequent and longer-lasting spells at the zero lower bound.

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Table 1: US Inflation Compensation Model

	5-Year, 5-Year Forward Inflation		
	1999-2011	2012-2019	1999-2019
Constant	0.00	0.00	0.00
	(0.00)	(0.00)	(0.00)
Core CPI surprise	0.15^{**}	-0.04	0.15^{**}
	(0.07)	(0.06)	(0.07)
Food & Energy CPI surprise	-0.02	-0.02	-0.02
	(0.04)	(0.05)	(0.04)
Constant $\times \mathcal{I}_{t \geq 2012}$			0.00
			(0.01)
Core CPI surprise $\times \mathcal{I}_{t \geq 2012}$			-0.20^{**}
			(0.09)
Food & Energy CPI surprise $\times \mathcal{I}_{t \geq 2012}$			0.00
			(0.07)
Observations	155	94	249
R^2	0.04	0.01	0.03

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

Table 2: Structural Break Tests at an Unknown Date in US Inflation Compensation Model

		5-Year, 5-Year Forward Inflation	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2002:08	1.88	0.25
		[0.82]	[0.71]
Core CPI surprise	2010:05	9.79^{**}	2.02^{*}
		[0.03]	[0.05]
Food & Energy CPI surprise	2004:05	1.10	0.14
		[0.99]	[0.92]
All Coefficients	2010:05	11.47	2.78
		[0.13]	[0.19]
Residual Variance	2013:03	3.01	0.80
		[0.66]	[0.27]

Note: Approximate asymptotic p-values from Hansen (1997) in brackets.
Observations: 249. $^{*}p < 0.10$,$^{**}p < 0.05$

Table 3: US Forward Rate Model

	1-Year, 9-Year Forward Rate		
	1997-2011	2012-2019	1997-2019
Constant	0.00	0.00	0.00
	(0.01)	(0.01)	(0.01)
Core CPI surprise	0.11*	−0.04	0.11*
	(0.06)	(0.06)	(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 $\times \mathcal{I}_{t \geq 2012}$			−0.14
			(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	95	274
R^2	0.04	0.13	0.06

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

Table 4: US Forward Rate Model: Structural Break Tests

		1-Year, 9-Year Forward Rate	
		Andrews-Quandt	Andrews-Ploberger
	Date	Test Statistic	Test Statistic
Constant	2003:09	3.46	0.55
		[0.46]	[0.40]
Core CPI surprise	2008:12	6.39	1.49*
		[0.14]	[0.10]
GS Agriculture Price Index	2003:09	5.26	1.23
		[0.22]	[0.14]
GS Energy Price Index	2008:10	3.60	0.57
		[0.44]	[0.39]
All Coefficients	2008:12	9.71	2.88
		[0.41]	[0.34]
Residual Variance	2013:11	2.38	0.49
		[0.69]	[0.45]

Note: Approximate asymptotic p-values from Hansen (1997) in brackets.
Observations: 274
$^{*}p < 0.10, ^{**}p < 0.05$

Table 5: US Reduced-Form Phillips Curve Regressions

	US Data			**Model Simulations**		
	Core Inflation			Inflation		
	1999-2011	2012-2019	1999-2019	1999-2011	2012-2019	1999-2019
Constant	3.18^{**}	2.22^{**}	3.18^{**}	0.00	0.00	0.00
	(0.22)	(0.22)	(0.22)	(0.15)	(0.09)	(0.15)
Unemployment Rate	-0.19^{**}	-0.04	-0.19^{**}	-0.19^{**}	-0.08^{**}	-0.19^{**}
	(0.03)	(0.04)	(0.03)	(0.02)	(0.03)	(0.02)
Constant $\times \mathcal{I}_{t\geq 2012}$			-0.96^{**}			-0.01
			(0.28)			(0.16)
Unemployment Rate $\times \mathcal{I}_{t\geq 2012}$			0.14^{**}			0.11^{**}
			(0.04)			(0.04)
Observations	156	96	252	156	96	252
R^2	0.49	0.09	0.45	0.62	0.22	0.55

Note: Core Inflation is measured as the year/year percent change in the CPI excluding food and energy. For the regressions on US Data, Newey-West standard errors with 12 lags are shown in parenthesis. For the regressions on Model-Simulated Data, bootstrapped standard errors are shown. See Section 3.2 for more details. $^{*}p < 0.10, ^{**}p < 0.05$

Table 6: US Expectations-Augmented Phillips Curve Regressions

	US Data			Model Simulations		
	Core Inflation less Expected Inflation			Inflation less Expected Inflation		
	1999-2011	2012-2019	1999-2019	1999-2011	2012-2019	1999-2019
Constant	0.09	0.21	0.09	-0.02	0.00	-0.02
	(0.41)	(0.21)	(0.41)	(0.63)	(0.28)	(0.63)
Unemployment Rate	-0.18^{**}	-0.20^{**}	-0.18^{**}	-0.34^{**}	-0.25^{**}	-0.34^{**}
	(0.07)	(0.04)	(0.07)	(0.13)	(0.07)	(0.13)
Constant $\times$ $\mathcal{I}_{t\geq 2012}$			0.12			0.01
			(0.46)			(0.67)
Unemployment Rate $\times$ $\mathcal{I}_{t\geq 2012}$			0.02			0.09
			(0.08)			(0.15)
Observations	156	96	252	156	96	252
R^2	0.17	0.51	0.22	0.26	0.49	0.50

Note: Core Inflation is measured as the year/year percent change in the CPI excluding food and energy. Following Coibion and Gorodnichenko (2015), we use 1-year ahead household inflation expectations as measured by the University of Michigan Survey of Consumers to measure expected inflation. For the regressions on US data, Newey-West standard errors with 12 lags are shown in parenthesis. For the regressions on Model-Simulated Data, bootstrapped standard errors are shown. See Section 3.2 for more details. $^{*}p < 0.10$,$^{**}p < 0.05$

Table 7: Japan Forward Rate Model

	1-Year, 9-Year Forward Rate		
	2001-2012	2013-2019	2001-2019
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 $\times \mathcal{I}_{t \geq 2013}$			-0.05
			(0.07)
GS Agricultural Price Index $\times \mathcal{I}_{t \geq 2013}$			0.00
			(0.00)
Observations	136	84	220
R^2	0.02	0.04	0.02

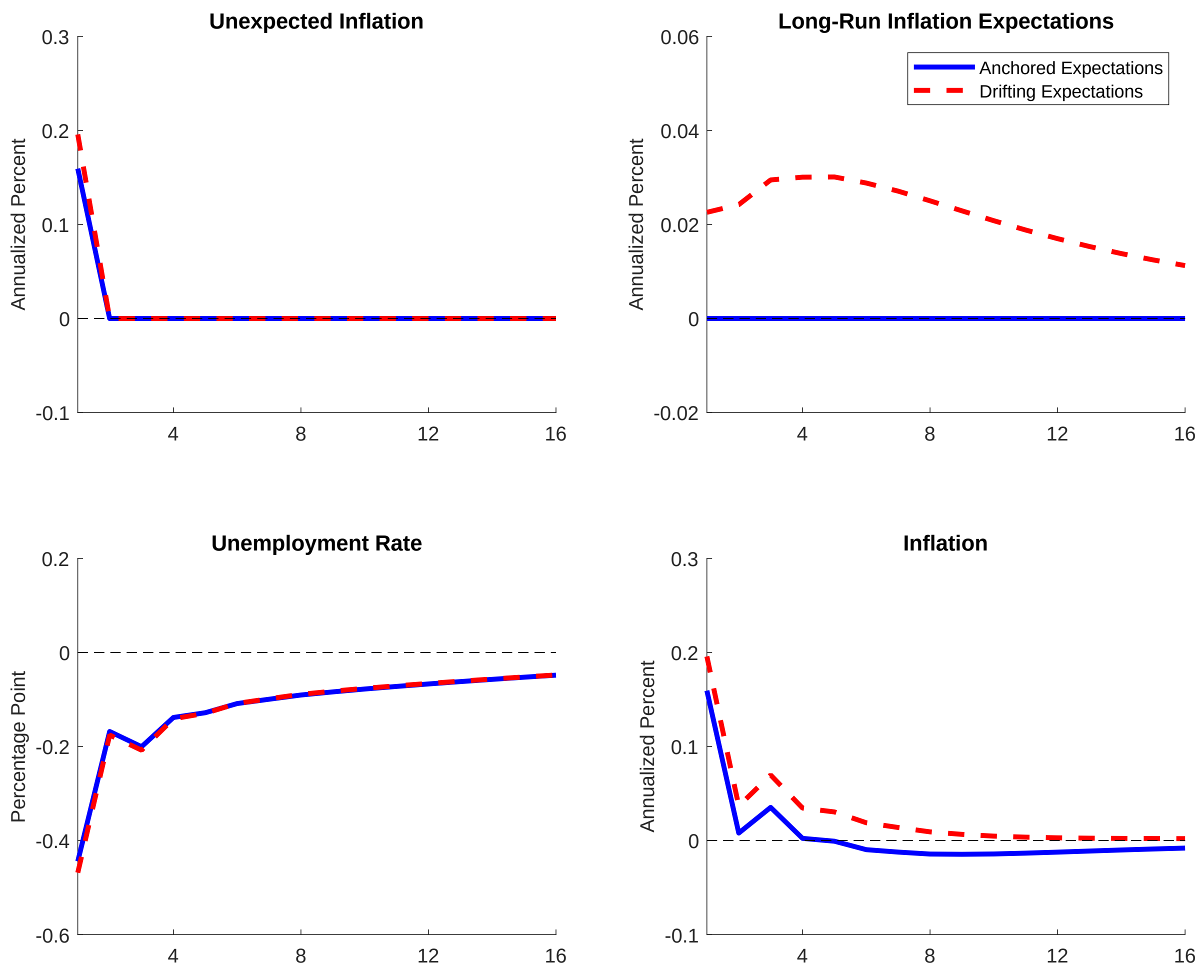
Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$

Table 8: Japan Reduced-Form Phillips Curve Regressions

	Core Inflation		
	2001-2012	2013-2019	2001-2019
Constant	1.33^{**}	0.38	1.33^{**}
	(0.56)	(0.57)	(0.56)
Unemployment Rate	-0.41^{**}	-0.04	-0.41^{**}
	(0.13)	(0.20)	(0.13)
Constant $\times \mathcal{I}_{t \geq 2013}$			-0.95
			(0.82)
Unemployment Rate $\times \mathcal{I}_{t \geq 2013}$			0.37
			(0.25)
Observations	136	84	220
R^2	0.31	0.00	0.62

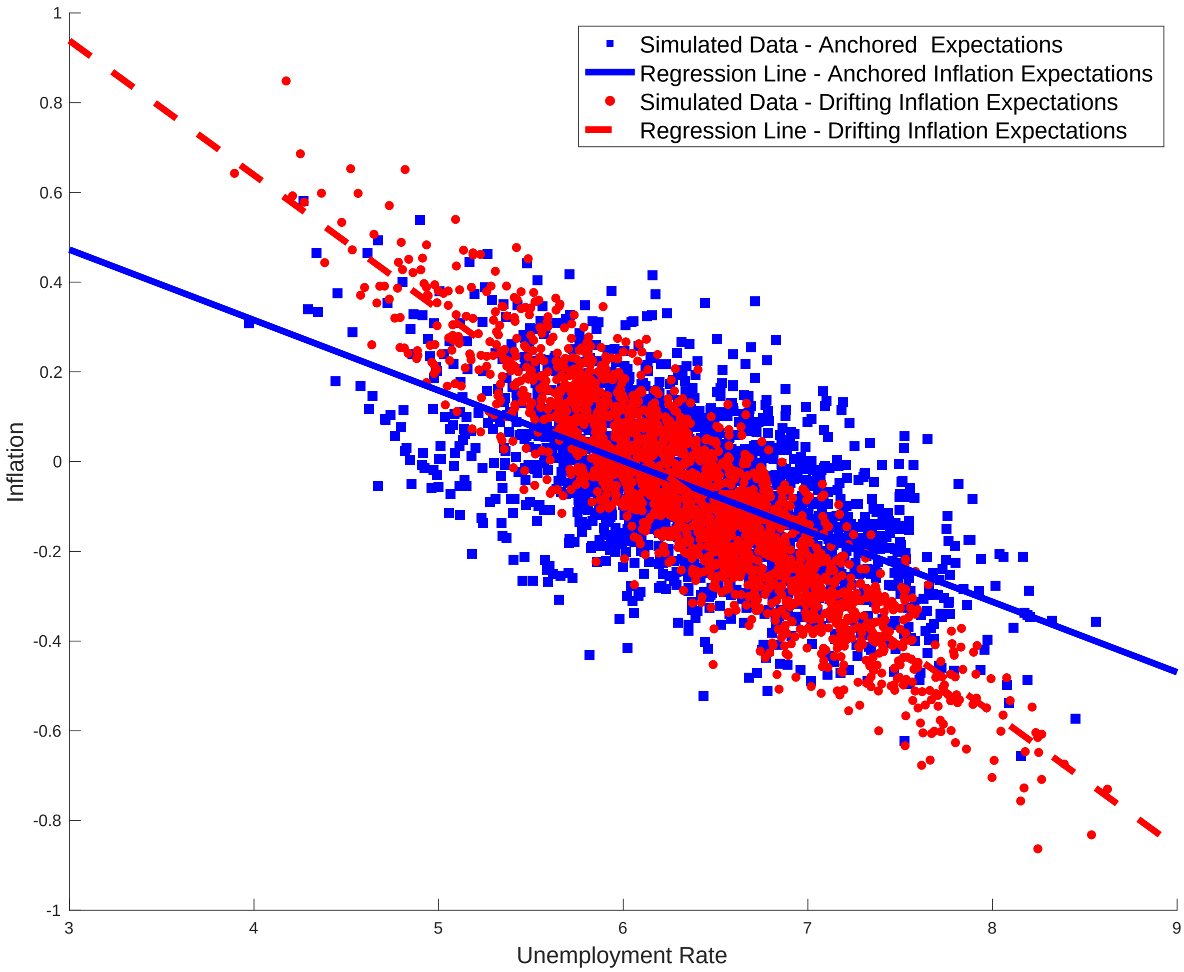
Note: Core Inflation is measured as the year/year percent change in the CPI excluding food, energy, and sales tax changes. Newey-West standard errors with 12 lags in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$

Figure 1: Impulse Responses to Demand Shock Under Drifting & Anchored Expectations



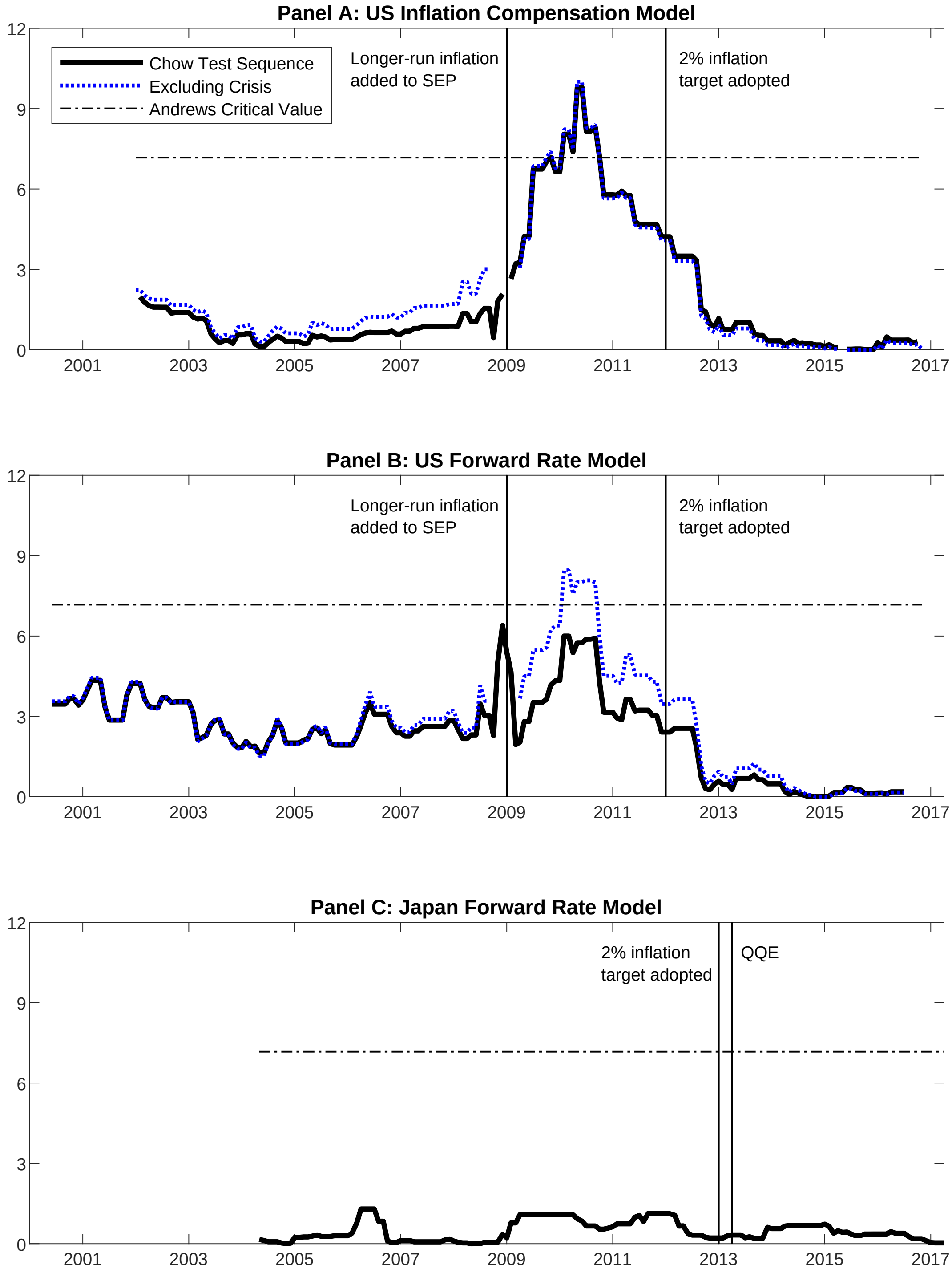
Note: The figure shows the impulse responses in the theoretical model to a one standard deviation aggregate demand (preference) shock under both drifting and anchored coefficient estimates from Table 1. See Section 2.3 for additional details.

Figure 2: Simulated Reduced-Form Phillips Curve in Theoretical Model



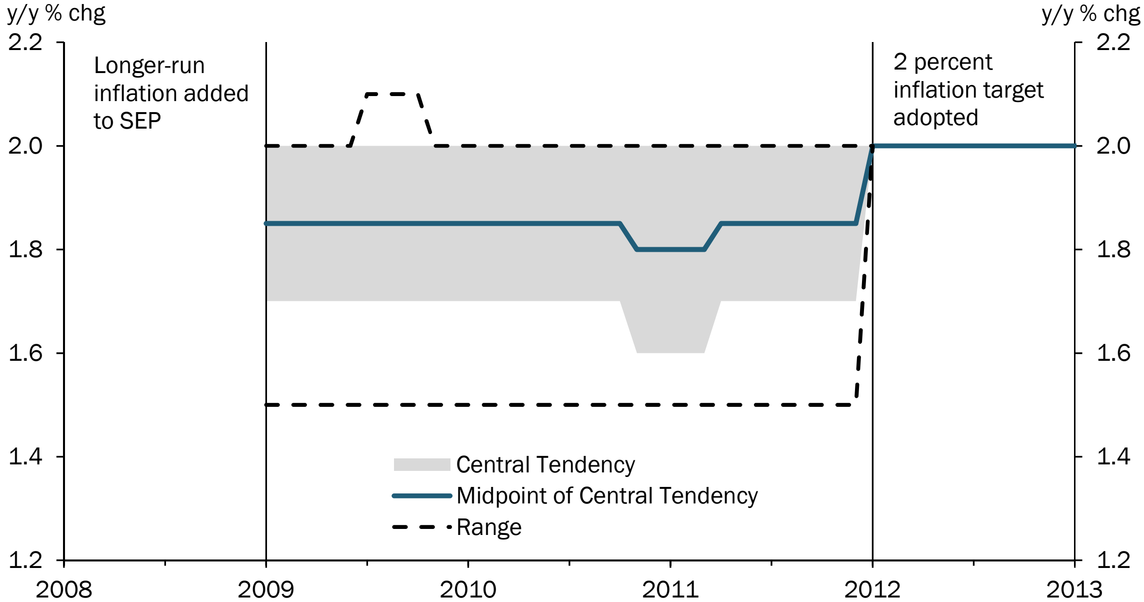
Note: The figure shows the simulated data and model-implied regression lines in the theoretical model under both drifting and anchored coefficient estimates from Table 1. Inflation is measured in annualized percent. See Section 2.4.2 for additional details.

Figure 3: Chow Test Sequence for Core Inflation Coefficient as a Function of Breakdate



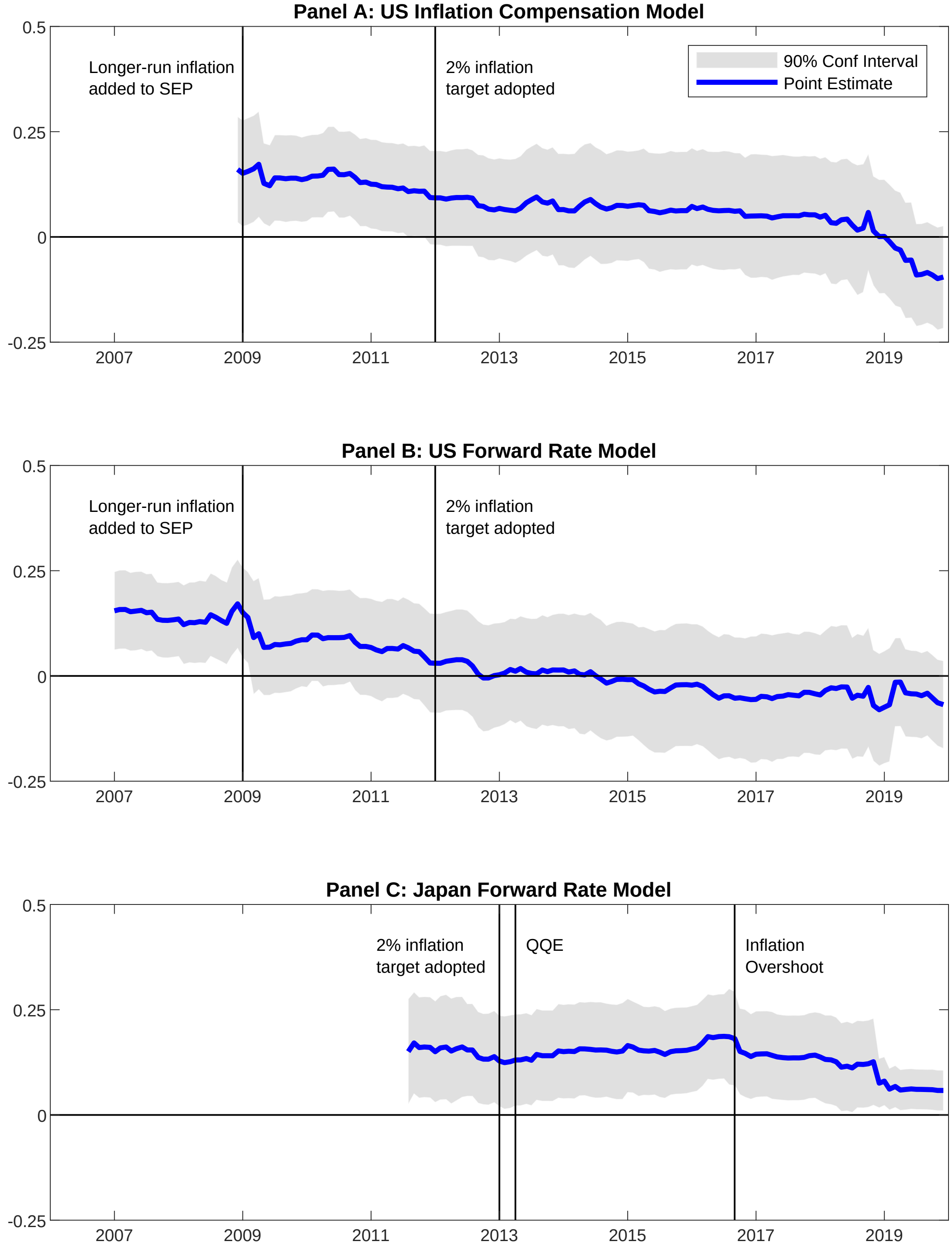
Note: Each panel shows the sequence of Chow test statistics as a function of candidate break dates. For each model, 15% of the observations on the ends of the sample are not examined as break points. 10% critical values are obtained from Andrews (1993) for $\pi_0 = 0.15$ and $p = 1$.

Figure 4: FOMC Summary of Economic Projections: Longer-run PCE Inflation



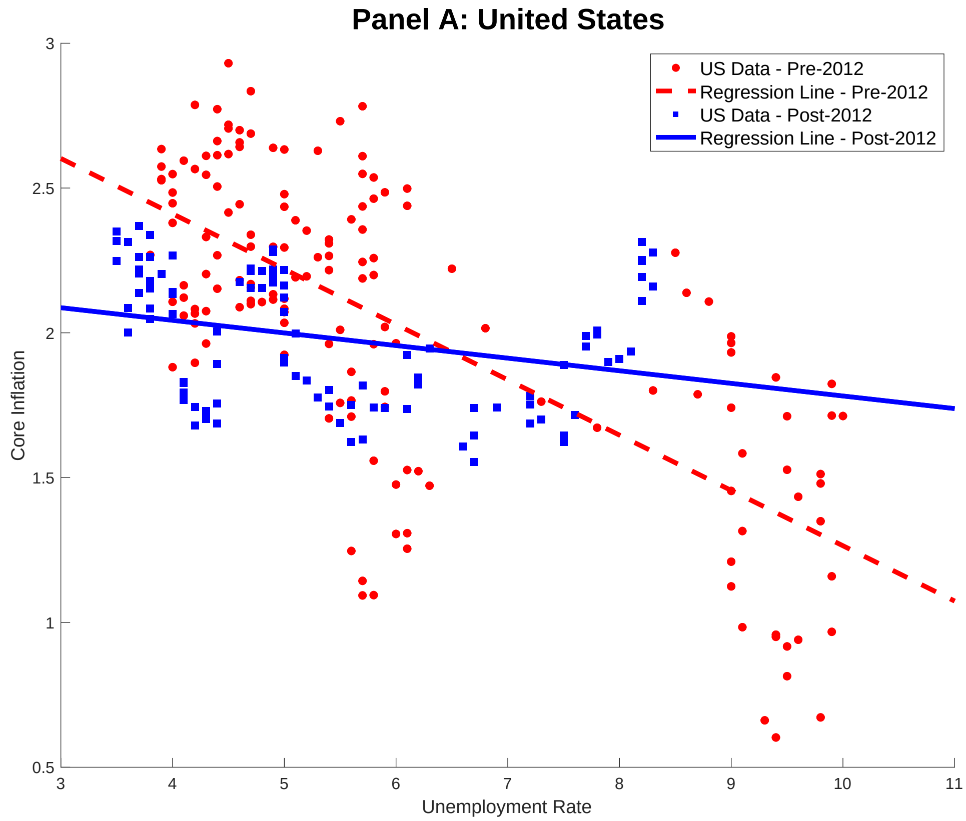
Note: Longer-run projections represent each FOMC participant's assessment of the rate to which inflation would be expected to converge under appropriate monetary policy and in the absence of further shocks to the economy. The central tendency discards the three highest and three lowest projections.

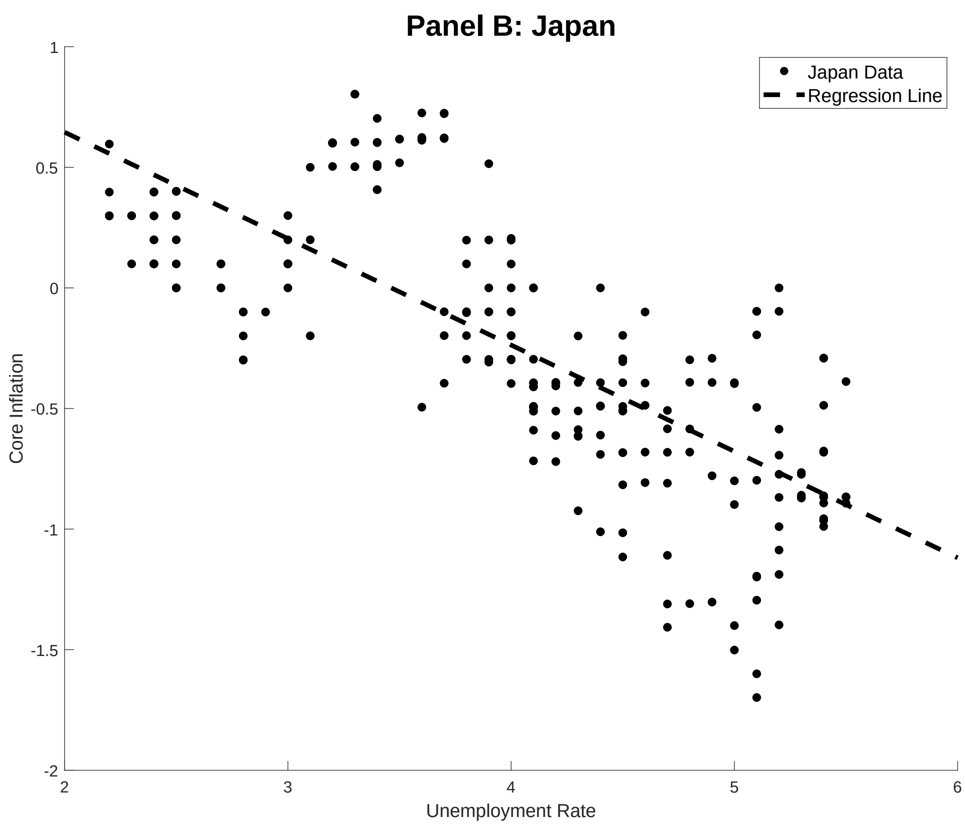
Figure 5: Rolling Window Estimates of Core Inflation Coefficient



Note: Each panel shows the sequence of estimates of δ^{π} 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 Eicker-White standard error.

Figure 6: Phillips Curves Before & After Adoption of an Inflation Target in the US & Japan





Note: Core Inflation is measured as the year/year percent change in the CPI excluding food and energy for the US and the year/year percent change in the CPI excluding food, energy, and sales tax changes for Japan. The US sample period is January 1999 - December 2019 and the sample period for Japan is September 2001 - December 2019.