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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 the Federal Reserve's communication of a numerical inflation objective in 2012 better anchored inflation expectations. Moreover, this improved anchoring can account for much of the observed flattening of the Phillips curve in the US over the past decade. Similar analysis reveals no evidence of anchoring in Japan despite the Bank of Japan's announcement of a numerical target, suggesting that merely announcing an inflation objective may not be sufficient to better anchor inflation expectations.

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 inflation target of 2 percent. 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 empirically 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 observed 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 increases 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 following 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 communication of a numerical inflation target. In particular, we show that prior to 2012 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, once the Federal Reserve began communicating a numerical inflation objective, a battery of econometric tests suggest that far forward inflation compensation ceased to respond to unanticipated inflation. In further evidence of anchoring in the US, we also find a statistical break in the reduced-form Phillips curve. Split-sample regressions indicate that inflation became less sensitive to fluctuations in the unemployment rate after the FOMC's adoption of a longer-run inflation target. Moreover, Monte Carlo simulations from our theoretical model show that the anchoring of inflation expectations can account for nearly all of the recent flattening of the reduced-form Phillips curve in the United States over the last decade.

In contrast to the US experience, we find little evidence of anchoring in Japan following the BOJ's adoption of a numerical inflation target in 2013. High-frequency evidence suggests that inflation compensation continues to drift with inflation surprises in Japan. Moreover, we fail to find evidence of instability in the reduced-form Phillips curve in Japan after the 2013 announcement. However, by both metrics, 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. These contrasting findings for the United States and Japan show that merely announcing an inflation objective may not be sufficient to better anchor inflation expectations.

2 How Our Work Relates to the Existing Literature

This paper relates to two active areas of macroeconomic research. First, many papers evaluate the efficacy of inflation-targeting policies in anchoring longer-run inflation expectations. For example, Gürkaynak et al. (2007), Gürkaynak, Levin and Swanson (2010) and Beechey, Johannsen and Levin (2011) conduct detailed, cross-country analysis on the behavior of market-based measures of long-run inflation expectations. Similar to our work, these papers exploit high-frequency identification to study the passthrough of unanticipated inflation shocks to these measures of longer-run inflation expectations under different inflation-targeting regimes. These papers conclude that, in estimation samples prior to 2012, inflation expectations were generally better anchored in the Euro Area, Canada, the UK and Sweden when compared to the United States. These authors conjecture that this difference arose because the central banks of these countries publicly adopted numerical targets for inflation, whereas the FOMC had not yet adopted a numerical objective at that time.

Our paper builds on these influential works in several dimensions. First, our results provide external validity to their analysis and conclusions. Indeed, we find a reduction in the sensitivity of far forward measures of nominal compensation to realized inflation surprises in the United States after the FOMC communicated a numerical inflation objective. Thus, as predicted by these previous studies, we find that inflation expectations became better anchored once the FOMC began communicating a numerical objective for longer-run inflation. Second, we subject the event-study models in this previous literature to formal tests of structural instability. Our paper highlights that a key prediction of anchoring inflation expectations is that structural changes in macroeconomic relationships should coincide with the policy change. Therefore, by explicitly applying break tests to the event-study regressions in this prior literature, we are able to provide causal evidence as to whether or not a publicly announced change in monetary policy was the likely source of the change in the behavior of inflation and longer-run inflation expectations. Finally, we provide Monte-Carlo evidence that, despite the small-samples that arise from our focus on recent policy changes, an econometrician can identify structural breaks in the data arising from the anchoring inflation expectations.

Our paper also relates to a second strand of macroeconomic research which explores instabilities in the Phillips curve relationship resulting from changes in the conduct of monetary policy. Recent contributions in this area emphasize the apparent flattening of the reduced-form Phillips curve in the United States since the 1990's (Erceg et al., 2018; Jorgensen and

Lansing, 2019; Del Negro et al., 2020; Hazell et al., 2020). These papers all attribute at least a portion of this flattening to a change in the conduct of monetary policy, either through more aggressive inflation stabilization or a better anchoring of inflation expectations. Others, such as Blanchard (2016); Carvalho et al. (2021), argue that inflation expectations became better anchored following the Volker disinflation of the early 1980s.

In this paper, we instead focus on more recent changes in FOMC communication about its longer-run inflation objective over the past two decades. We show that, even after the improvements in communication and inflation stabilization that took place in the 1980s and 1990s, the communication of a numerical inflation objective between 2009 and 2012 further improved the anchoring of inflation expectations. This change in communication policy results in a further flattening of the Phillips curve, which we document empirically and reproduce in our theoretical model.

In addition, we also draw out the implications for both the reduced-form and structural Phillips curve from a better anchoring of inflation expectations in our model and in the data. Thus, our work connects to Bullard (2018) and McLeay and Tenreyro (2020), which highlight the role that the conduct of monetary policy plays in identifying the reduced-form Phillips curve slope despite *stability* in the underlying structural Phillips curve. Indeed, in our model, the slope of the structural Phillips curve which relates inflation to labor market tightness and near-term inflation expectations is invariant to the degree of anchoring.

We show that, unlike the reduced-form relationship between inflation and unemployment, a Phillips curve regression model augmented with household's one-year-ahead inflation expectations, is stable before and after 2012. On the one hand, this aligns with Coibion and Gorodnichenko (2015) that the finding that near-term household inflation expectations can serve as a reasonable proxy for the near-term inflation expectations of price setters. On the other hand, we argue that inflation expectations have become better anchored since 2012 whereas Coibion and Gorodnichenko (2015) contend that a Phillips curve regression augmented with household's near-term inflation expectations is stable precisely because household's inflation expectations are unanchored.

We resolve this apparent tension by highlighting the disparate behavior between household's near-term and longer-term inflation expectations. In particular, we provide time-series evidence that, similar to market-based measures of longer-term inflation expectations but unlike household's near-term inflation expectations, household's longer-term inflation expectations also became largely unresponsive to unanticipated inflation shocks — including energy price shocks — after 2012. Therefore, one contribution of our work relevant for the growing literature on learning and expectations formation is to demonstrate that credible central bank announcements can anchor longer-term inflation expectations without perfectly stabilizing near-term inflation expectations.

This last result, together with our evidence for Japan, demonstrates how our findings can be used to shed light on the underlying mechanism that drives the anchoring of inflation expectations. Indeed, the fact that we find that longer-term inflation expectations likely remain unanchored in Japan underscores that merely publishing a numerical objective for inflation is not sufficient to anchor inflation expectations at the central bank's target. Instead, for a central bank announcement to effectively anchor longer-term inflation expectations, the announcement likely needs to be accompanied by some combination of a spelled out plan for achieving this target (Eusepi and Preston, 2010; Davig and Foerster, 2021), a recent track record of success in stabilizing inflation (Gáti, 2020; Carvalho et al., 2021), or sufficient credibility (Hausman and Wieland, 2015; De Michelis and Iacoviello, 2016).

3 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.

3.1 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:

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

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. 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. π^* represents the central bank's inflation objective, if one has been announced. Alternatively, our specification also allows for an undefined inflation objective if $\rho^\pi = 1$.[1] Thus, this specification allows for three possible cases, all of which may be relevant in practice: (1) Drifting longer-run expectations without a formal central bank objective, (2) drifting expectations despite the announcement of an inflation objective, and (3) fully anchored inflation expectations.

In our following analysis, we remain agnostic about the exact microfoundations of Equation (1) and instead focus on the testable implications of anchoring inflation expectations. 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. In their setting, δ^π has a similar interpretation as it governs the signal to noise ratio placed on unanticipated or unforecastable changes in inflation. This specification also builds upon the established 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.

3.2 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

[1] Recent work such as Gáti (2020) and Carvalho et al. (2021) provides an alternative method of modeling longer-term inflation expectations using adaptive learning. In this paper, however, we focus on identifying empirically testable predictions due to the anchoring of inflation expectations following the announcement of a numerical inflation target. Such announcement effects would play no meaningful role in models of adaptive expectations.

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.

3.2.1 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.[2] The household chooses its consumption and bond holdings to maximize its utility subject to its budget constraint each period:

$$C_t + \frac{B_t}{P_t R_t} = \frac{B_{t-1}}{P_t} + W_t N_t + \phi_u (1 - N_t) + D_t - T_t, \qquad \forall t \geq 0,$$

where P_t denotes the aggregate price level, B_t denotes holdings of a nominal risk-free bond, R_t denotes the nominal interest rate, W_t denotes the real wage rate, ϕ_u denotes an unemployment benefit (the replacement ratio), D_t denotes profit income from ownership of intermediate goods producers and of retailers, and T_t denotes a lump-sum tax paid to the government.

3.2.2 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

[2] 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.

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.

3.2.3 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:

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

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:

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

3.2.4 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 in which the parameter b determines the bargaining weight. 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.

3.2.5 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:

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

where ϕ_π denotes the central banks response to inflation deviations. If long-term inflation expectations in Equation (1) follow a random walk ($\rho^\pi = 1$), then inflation and short-term policy rates are nonstationary. However, detrended inflation Π_t / Π_t^{LT} is stationary. Therefore, as in Ireland (2007), we specify a coefficient of one on lagged policy rates to write the policy rule in a stationary form.[3]

3.3 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. Table 1 contains the calibrated values for the model parameters. Since our model shares

[3]See the Appendix for more details. Ireland (2007) also includes a nontrivial response of policymakers to changes in output growth. However, we find that including a response to output growth generates much larger (and likely counterfactual) fluctuations in inflation when we incorporate frictions in the labor market, so we remove this feature to generate more sensible inflation dynamics.

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, with a few important exceptions. 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. Under the drifting inflation expectations calibration, we set the degree of anchoring δ^π to align with our high-frequency estimates before the United States formally adopted its inflation objective. Also, we set $\rho^\pi = 1$ to be consistent with an undefined inflation objective prior to 2012.[4] Then, leaving all other parameters unchanged, we set $\delta^\pi = 0$ to simulate the dynamics under the anchored expectations calibration. We discuss the empirical support for these calibration choices in detail later in the paper. 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.

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 expectations. 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-term inflation expectations. Intuitively, once inflation expectations are anchored, realized inflation no longer informs agent's views of inflation over the longer 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 and thus weakens the reduced-form

[4] The testable predictions that emerge from our model are robust to alternative calibrations of ρ^π, which corresponds to the case in which the central bank has announced a formal inflation objective yet longer-term expectations continue to drift with realized inflation outcomes. For example, we can derive the same testable prediction in Equation (8) with $0 < \rho^\pi < 1$.

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.

3.4 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.

3.4.1 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.

3.4.2 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). In Section 4.2.2, 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.

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.

4 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.

4.1 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 1-year, 9-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 mechanical 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.[5] 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.[6] 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,

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

[5] 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).

[6] 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 1-year, 9-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 2 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 1-year, 9-year forward inflation compensation by nearly three basis points. The finding that far-forward market-based inflation compensation drifts with core CPI surprises in our early sample is consistent with the findings in Beechey, Johannsen and Levin (2011) and Bauer (2015), both of which conclude their estimation in 2007. However, after the FOMC formally adopted its inflation target, the coefficient on the core CPI surprise falls and becomes statistically indistinguishable from zero. In the third column of Table 2, 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.[7]

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

[7]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 considerable information about the food and energy components ahead of the CPI release.

4.1.1 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 Equation (9) 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 resulting from the adoption of an inflation target, we would expect the estimated break date to follow a publicized 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 3 shows the results of Andrews (1993) or 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 the 5% level for both tests, indicating 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. 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. This pattern suggests a one-time structural break in the sensitivity of long-term inflation expectations to core CPI surprises occurring around 2010. More formally, if we split the sample into two subsamples 1999-2010 and 2010-2019, the Andrews-Quandt and Andrews-Ploberger tests indicate no other breaks.

Our 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 publish quarterly projections for "longer-run" inflation. In January of 2009, the Summary of Economic Projections (SEP) added longer-run inflation which, "[...] 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 the view that the numerical SEP projections for longer-run inflation were intended to better anchor inflation expectations. On 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 hypothesized that the addition of longer-run inflation projections to the SEP would largely accomplishing the desired degree of anchoring without needing to adopt a Committee-wide longer-run numerical target.[8] 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." Our results appear to largely support her prediction. In particular, our break date suggests that the modifications to the SEP — which pre-dated the Committee-wide adoption of a formal 2-percent target in 2012 — played an instrumental role in anchoring inflation expectations.

4.1.2 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

[8]Our event-study regressions do not identify the level at which inflation expectations may have become anchored. Bundick and Smith (2021) show that the initial SEP projections from 2009 through 2011 were centered below 2 percent, which they argue may have served to anchor inflation expectations below 2 percent. Similarly, Shapiro and Wilson (2019) perform text analysis of FOMC communication to estimate the FOMC's objective function and also find that the FOMC's implicit objective for inflation may have resided below 2 percent.

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 1-year, 9-year forward nominal 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*.

Rather than using inflation compensation measured from inflation-indexed bonds, this alternative model uses far forward measures of nominal interest rates 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 4 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

[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).

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

[12] While the Chow test statistic for a break after 2012 is slightly below the critical value, the p-value of the Chow test for δ^π is 0.1003, indicating that our findings are generally robust.

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 for Japan.

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.

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 4 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 4 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 declines to around 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.

4.2 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 5 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. As we show in Section

3.4.2, 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.

4.2.1 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 2 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 0.27, our high-frequency coefficient on core CPI from Table 2 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 95 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.[13] 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-

[13]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. We also find similar results if we instead use annualized quarterly inflation rather than year-over-year inflation in our simulated Phillips curve exercise.

form Phillips curve slope (-0.19) that we observe in the data during the 1999-2011 period.[14]

The far right columns of Table 5 shows the resulting regression coefficients and their associated 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 due to the better anchoring of inflation expectations.

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 -0.07. 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 (0.12), the 90% model-implied confidence interval of the post-anchoring flattening is (0.06, 0.18), which contains the empirical estimate of 0.14. Thus, even at the 10% 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 over the past decade.

4.2.2 Reconciling our Results with Coibion and Gorognichenko (2015)

Coibion and Gorodnichenko (2015) show that the Phillips curve has remained stable in recent decades once it is augmented with household's near-term inflation expectations. In sharp contrast to our results, the stability of their expectations-augmented Phillips curve results from unanchored rather than anchored household inflation expectations. In this section, we aim to reconcile our results with their findings. We first show that their Phillips curve specification exhibits stability amid the anchoring of longer-run inflation expectations because it proxies well the underlying structural Phillips curve in our model, which is also invariant to the degree of anchoring. Then, we provide evidence that while household's near-term inflation expectations may appear unanchored, household's *longer-term* inflation expectations

[14] Our calibrated value of $\phi_P = 280$ implies a coefficient on marginal cost of about 0.02, well within the common range of calibrated values from the literature. See Section 4.2.2 for further discussion on the reduced-form versus structural Phillips curves implied by our model.

appear to have become better anchored since 2012. These findings suggest no disconnect between our results and the results in Coibion and Gorodnichenko (2015).

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 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.[15] 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

[15] 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.

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.[16]

One remaining tension between our analysis and the conclusion in Coibion and Gorodnichenko (2015) arises from our claim that longer-term inflation expectations have become better anchored since 2012. Coibion and Gorodnichenko (2015) argue instead that their Phillips curve specification exhibits stability precisely because household's near-term inflation expectations are largely unanchored. However, simulations from our model show that the anchoring of long-term inflation expectations does not eliminate fluctuations in near-term inflation expectations. Therefore, household's longer-run inflation expectations may have indeed become better anchored since 2012 even if their near-term inflation expectations continue to vary in response to unanticipated inflation. To examine this hypothesis, we return to the data to study whether there is a notable change in the pass through from inflation surprises to household's longer-term inflation expectations since 2012.

Split-sample VAR estimates suggest that household's longer-run inflation expectations (measured again by the University of Michigan Survey of Consumers) have become better anchored since 2012.[17] The left panel of Figure 6 shows that prior to 2012, increases in energy inflation as well as core inflation spilled over to household's longer-run inflation expectations. However, the right panel shows that household's longer-run inflation expectations ceased to

[16] 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 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. This conclusion is consistent with recent work by Barnichon and Mesters (2020) which argues that better anchoring of inflation expectations helps explain the decline in the Phillips multiplier following monetary policy shocks.

[17] The VAR model is identified through a recursive ordering that assumes that household's observe salient energy and food prices before submitting their longer-run inflation expectations but, consistent with the survey dates and the timing of the CPI release, they only learn of the realized value of core CPI after they submit their longer-run inflation expectations. Leveraging the timing of surveys and data releases follows the identifications strategy in Leduc, Sill and Stark (2007), Clark and Davig (2011), and Leduc and Sill (2013). More details on the VAR model and identification are provided in the appendix.

respond to those inflationary shocks in the post-2012 sample.[18] Interestingly, if we replace household's longer-term inflation expectations with their near-term (1-year ahead) inflation expectations, we observe significant pass through to 1-year ahead inflation expectations from energy prices in both sample periods.[19] Therefore, by distinguishing between near-term and longer-term inflation expectations, our results appear fully consistent with the findings of Coibion and Gorodnichenko (2015).

5 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 a "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 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 these policy announcements better anchored inflation expectations in Japan.

5.1 Inflation Compensation & Inflation News in Japan

Guided by the predictions of our theoretical model in Section 3, 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:

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

[18] Our results contrasts from those in Binder (2017) likely because she focuses on whether the level of expectations moved nearer to 2 percent after the FOMC's 2012 adoption of a numerical inflation target. In contrast, our approach allows for the possibility the household's inflation expectations remain biased, on average, but nevertheless became more stable.

[19] See the appendix for these additional results.

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.[20] 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 further indicate *stability* in the regression model in Equation (12) over time. The Andrews-Quandt 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 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 4 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 publicized by the BOJ, inflation expectations remain unanchored in Japan.

[20] 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 4.1.2. 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 $\delta^\pi = 0$.

5.2 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 5 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 anchor inflation expectations in Japan. This conclusion compliments the findings in Hausman and Wieland (2015); De Michelis and Iacoviello (2016) which suggest that poor credibility on the part of the BOJ may be impeding the transition to a higher inflation regime.

6 Conclusion

Though central banks have recently started to adjust and refine their monetary policy frameworks, inflation targeting remains the global benchmark for promoting price stability mandates. Many central banks have implemented this framework, in part, by communicating a numerical inflation target. Research on the efficacy of such communication, by Gürkaynak et al. (2007), Gürkaynak, Levin and Swanson (2010) and Beechey, Johannsen and Levin (2011), predicted that, by announcing a numerical inflation target, the Federal Reserve could enhance the degree to which inflation expectations are anchored. In this paper, we provide a detailed analysis of whether inflation expectations in the United States indeed became better anchored after the Federal Reserve moved towards publicizing its longer-run numerical objective for inflation between 2009 and 2012. Motivated by the predictions of a theoretical model, we test economic relationships for structural instability that should result upon anchoring expectations. Our empirical results suggest that the Federal Reserve's communication of a numerical inflation target has in fact lead to better anchored inflation expectations in the United States.

However, our analysis of Japan underscores the limitations of announcement effects in anchoring inflation expectations. Although the BOJ adopted and publicized an explicit numerical inflation objective about one year after the Federal Reserve, we find no statistically significant evidence that inflation expectations subsequently became better anchored 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 the announcement of a numerical inflation target. The contrasting experiences of the US and Japan that we document provide useful evidence to inform and study the mechanisms that lead to the successful anchoring of inflation expectations.

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Table 1: Calibrated Model Parameters

Parameter	Description	Value	Source
β	Household Discount Factor	0.9995	Ireland (2007)
χ	Disutility of Working Scalar	0.476	Leduc and Liu (2016)
θ	Elasticity of Substitution Intermediates	6.0	Ireland (2003)
α	Share Parameter in Matching Function	0.5	Blanchard and Galí (2010)
μ	Matching Efficiency	0.645	Leduc and Liu (2016)
ρ	Job Separation Rate	0.1	Monthly Separation Rate of 3.5%
ϕ_u	Flow Benefit of Unemployment	0.25	Hall and Milgrom (2008)
κ	Vacancy Cost	0.14	Leduc and Liu (2016)
b	Nash Bargaining Parameter	0.5	Blanchard and Galí (2010)
γ	Real Wage Rigidity	0.8	Gertler and Trigari (2009)
ϕ_P	Cost of Adjusting Nominal Prices	280	Calibrated to Match Phillips Curve
ϕ_π	Central Bank Response to Inflation	0.8594	Ireland (2007)
ρ_a	Preference Shock Persistence	0.9097	Ireland (2007)
σ_a	Preference Shock Volatility	0.01	Implies 1% Demand Shock

Table 2: Chow Test: US Inflation Compensation Model

	Estimation Sample		
Δ 1-Year, 9-Year Fwd Breakeven Inflation	1999-2011	2012-2019	1999-2019
Constant	0.00	0.00	0.00
	(0.01)	(0.00)	(0.00)
Core CPI surprise	0.27^{***}	-0.03	0.27^{***}
	(0.07)	(0.06)	(0.10)
Food & Energy CPI surprise	-0.02	-0.02	-0.02
	(0.05)	(0.04)	(0.05)
Constant $\times \mathcal{I}_{t \geq 2012}$			0.00
			(0.01)
Core CPI surprise $\times \mathcal{I}_{t \geq 2012}$			-0.30^{***}
			(0.11)
Food & Energy CPI surprise $\times \mathcal{I}_{t \geq 2012}$			0.05
			(0.06)
Observations	156	95	251
R^2	0.06	0.01	0.06

Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$,$^{***}p < 0.01$

Table 3: Structural Break Tests at an Unknown Date: US Inflation Compensation Model

		Structural Break Test	
Δ 1-Year, 9-Year Fwd Breakeven Inflation		Andrews-Quandt	Andrews-Ploberger
	Break Date	Test Statistic	Test Statistic
Constant	2003:03	1.30	0.11
		[0.96]	[0.99]
Core CPI surprise	2010:05	10.77^{**}	2.85^{**}
		[0.02]	[0.02]
Food & Energy CPI surprise	2011:08	3.12	0.38
		[0.53]	[0.55]
All Coefficients	2010:05	14.15^{**}	4.31^{**}
		[0.05]	[0.05]
Residual Variance	2003:04	3.21	1.05
		[0.51]	[0.18]

Note: Approximate asymptotic p-values from Hansen (1997) in brackets.
Observations: 251. $^{*}p < 0.10$,$^{**}p < 0.05$,$^{***}p < 0.01$

Table 4: Chow Test: US Forward Rate Model

	Estimation Sample		
Δ 1-Year, 9-Year Fwd Nominal 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	0.12	0.06

Note: Eicker-White standard errors in parenthesis. ${}^{*}p < 0.10$, ${}^{**}p < 0.05$, ${}^{***}p < 0.01$

Table 5: Chow Test: 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)	(0.19)	(0.07)	(0.19)
Unemployment Rate	-0.19^{***}	-0.04	-0.19^{***}	-0.19^{***}	-0.07^{***}	-0.19^{***}
	(0.03)	(0.04)	(0.03)	(0.03)	(0.02)	(0.03)
Constant $\times\ \mathcal{I}_{t\geq 2012}$			-0.96^{***}			-0.01
			(0.28)			(0.19)
Unemployment Rate $\times\ \mathcal{I}_{t\geq 2012}$			0.14^{***}			0.12^{***}
			(0.04)			(0.04)
Observations	156	95	251	156	95	251
R^2	0.49	0.09	0.45	0.56	0.29	0.59

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-dimulated data, we show bootstrapped standard errors. See Section 4.2 for more details. $^{*}p < 0.10$,$^{**}p < 0.05$,$^{***}p < 0.01$

Table 6: Chow Test: 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)	(0.70)	(0.24)	(0.70)
Unemployment Rate	-0.18^{***}	-0.20^{***}	-0.18^{***}	-0.25^{**}	-0.19^{***}	-0.25^{**}
	(0.07)	(0.04)	(0.07)	(0.12)	(0.05)	(0.12)
Constant $\times\ \mathcal{I}_{t\geq 2012}$			0.12			0.01
			(0.46)			(0.73)
Unemployment Rate $\times\ \mathcal{I}_{t\geq 2012}$			0.02			0.06
			(0.08)			(0.13)
Observations	156	95	251	156	95	251
R^2	0.17	0.51	0.22	0.17	0.43	0.44

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, we show bootstrapped standard errors. See Section 4.2 for more details. ${}^{*}p < 0.10$, ${}^{**}p < 0.05$, ${}^{***}p < 0.01$

Table 7: Chow Test: Japan Forward Rate Model

	Estimation Sample		
Δ 1-Year, 9-Year Fwd Nominal 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

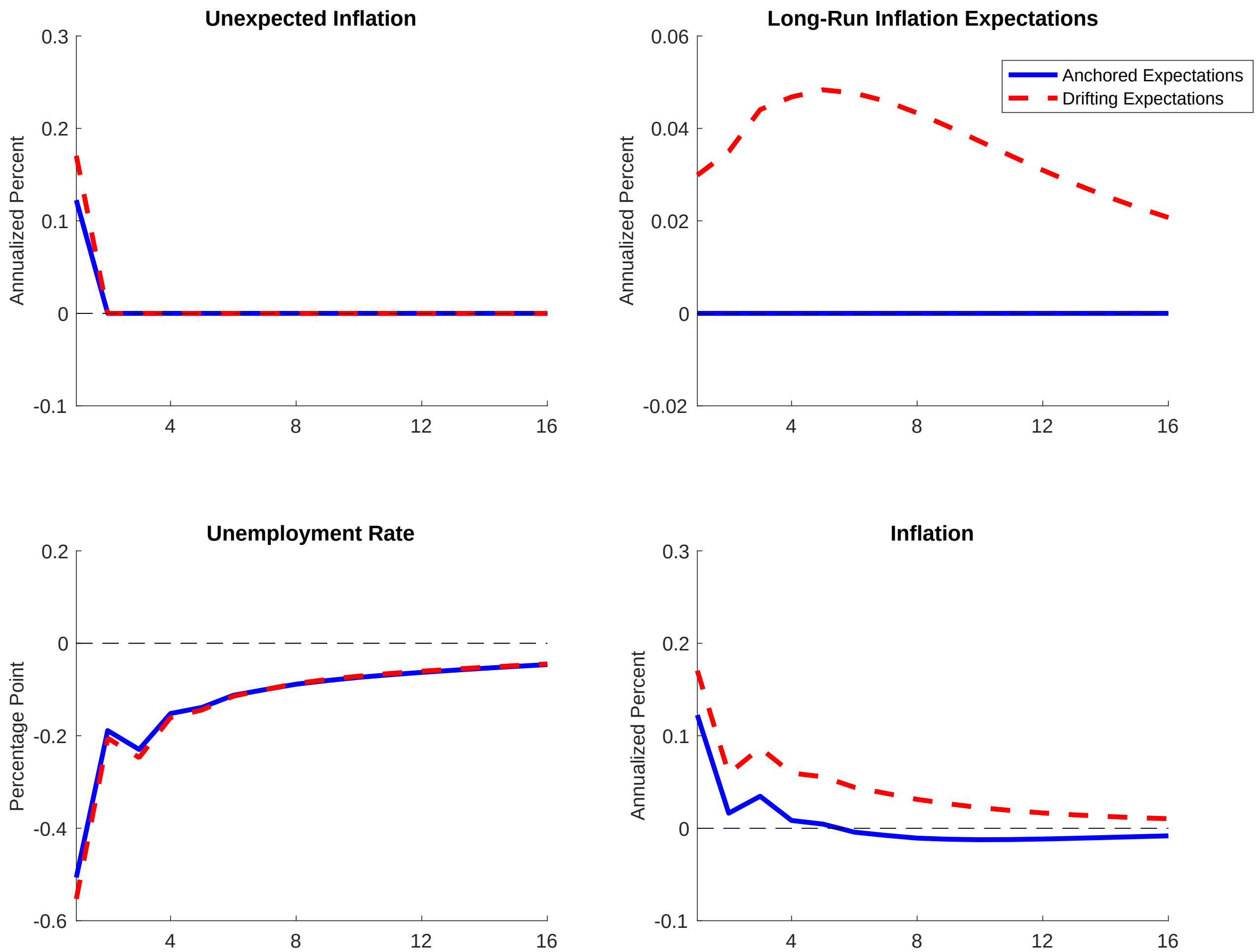
Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

Table 8: Chow Test: Japan Reduced-Form Phillips Curve Regressions

	Estimation Sample		
Core Inflation	2001-2012	2013-2019	2001-2019
	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

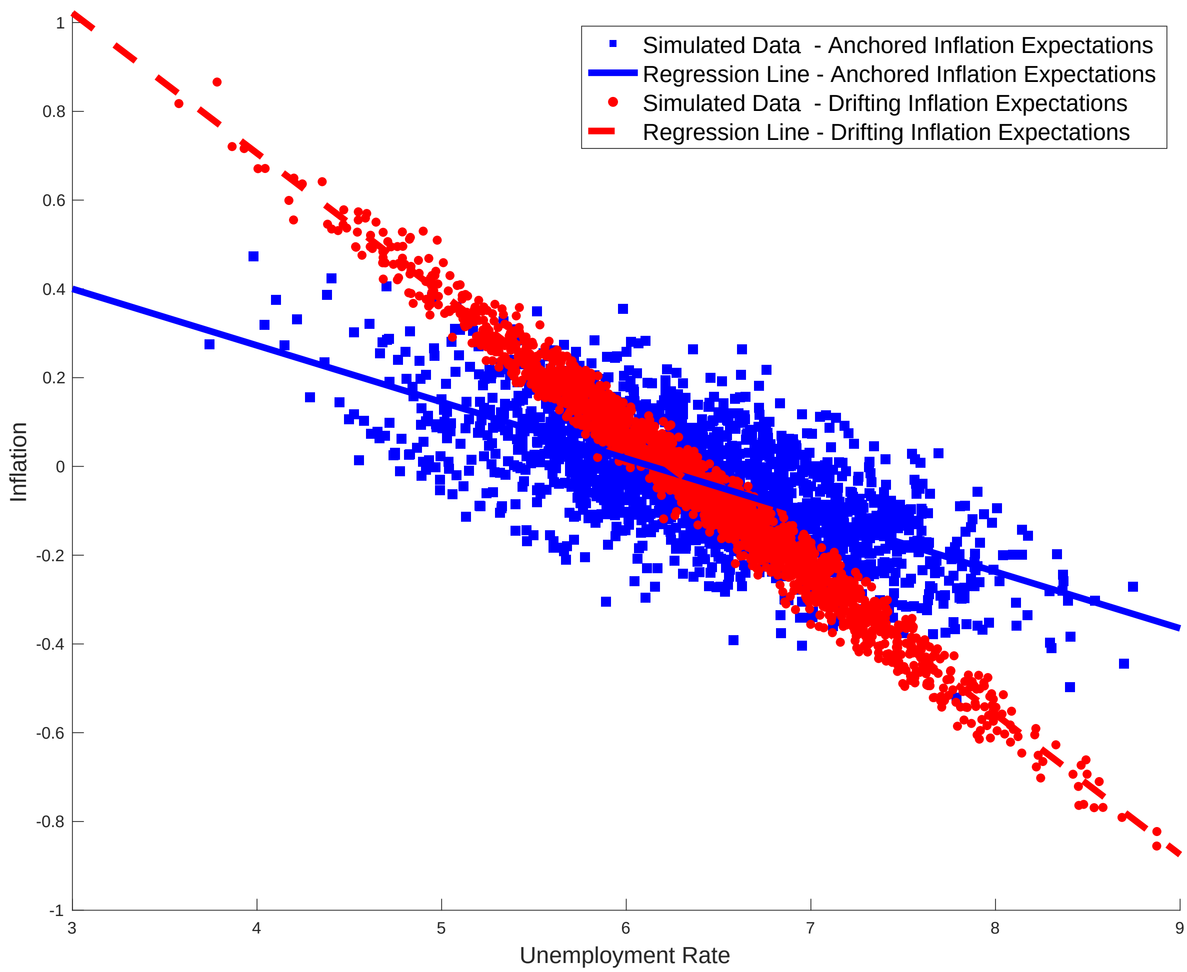
Note: Core Inflation is measured as the year/year percent change in the CPI excluding food, energy, and sales tax changes. Newey-West standard errors with 12 lags in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$, $^{***}p < 0.01$

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



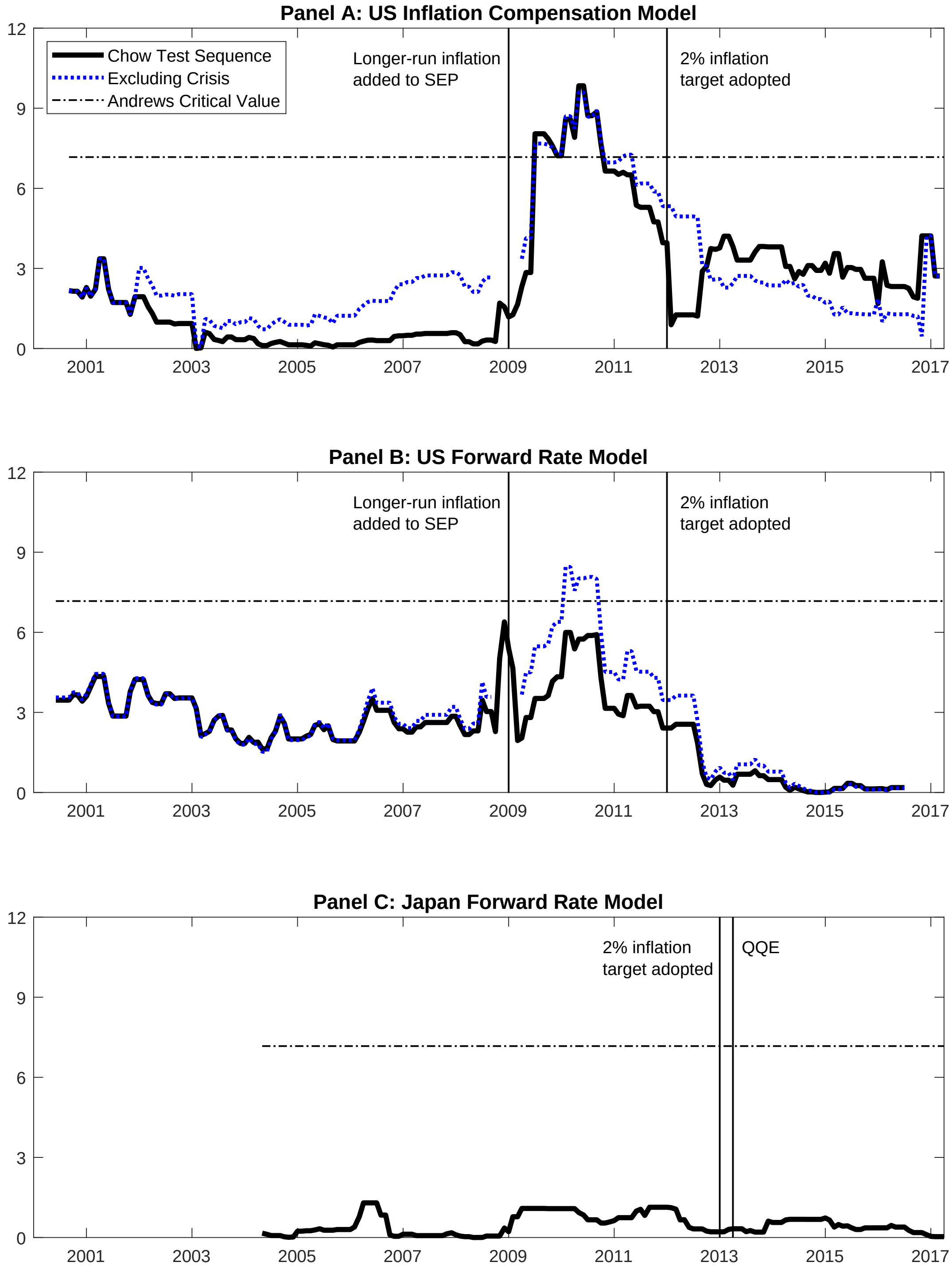
Note: The figure shows the impulse responses in the theoretical model to a one standard deviation aggregate demand (preference) shock under both drifting and anchored coefficient estimates from Table 2. See Section 3.3 for additional details.

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



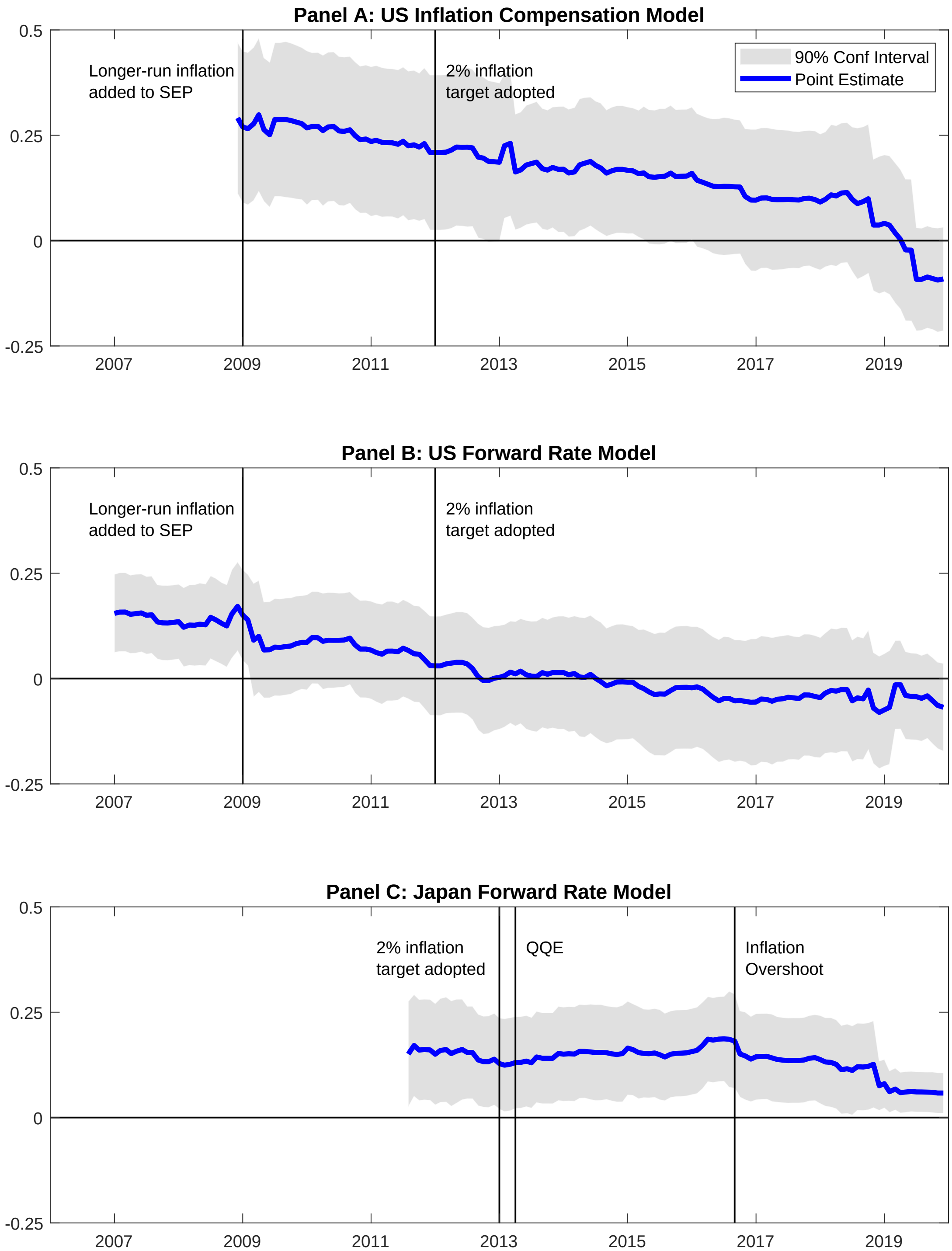
Note: The figure shows the simulated data and model-implied regression lines in the theoretical model under both drifting and anchored coefficient estimates from Table 2. Inflation is measured in annualized percent. See Section 3.4.2 for additional details.

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



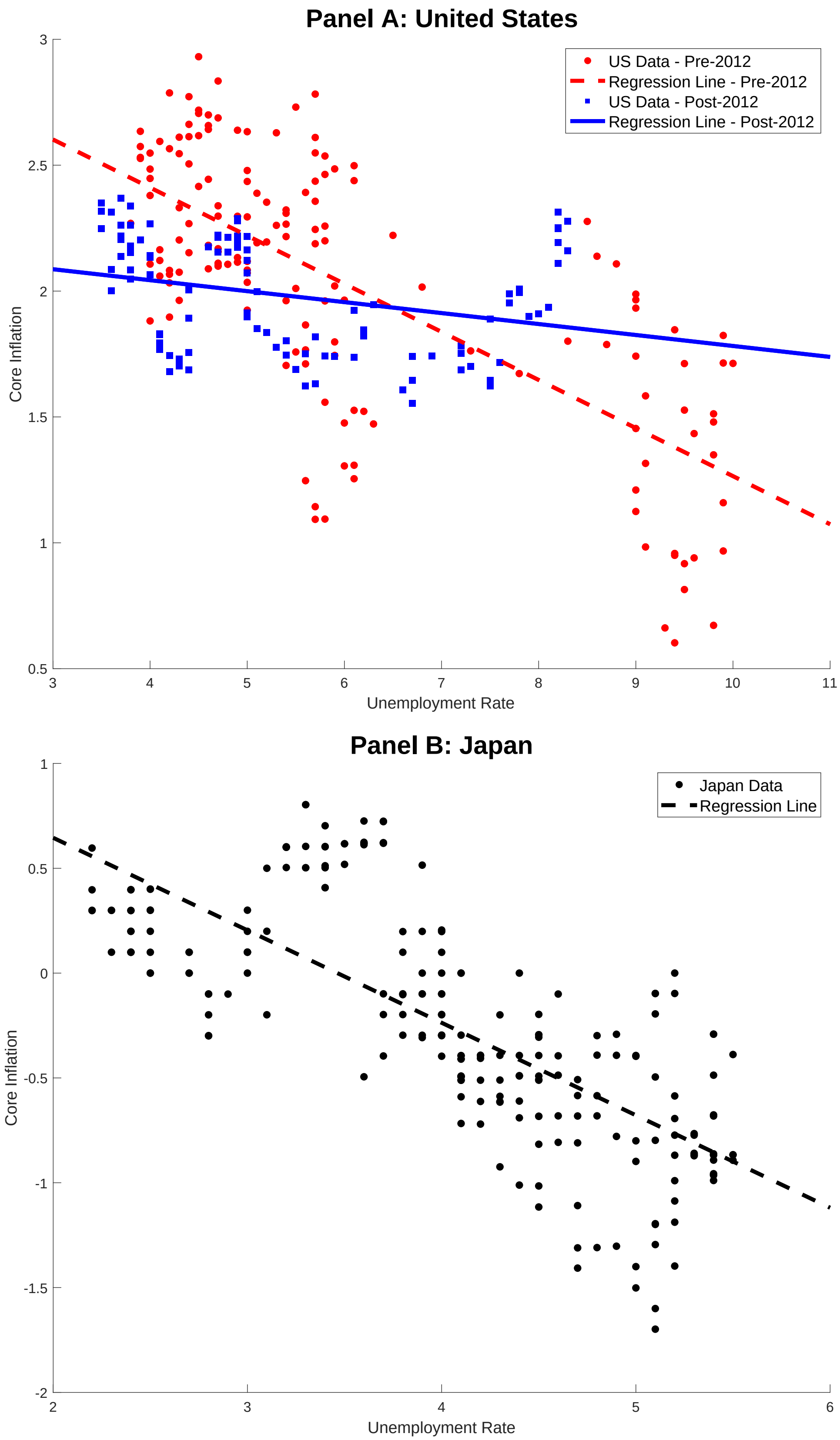
Note: Each panel shows the sequence of Chow test statistics as a function of candidate break dates. For each model, 15% of the observations on the ends of the sample are not examined as break points. 10% critical values are obtained from Andrews (1993) for $\pi_0 = 0.15$ and $p = 1$.

Figure 4: Rolling Window Estimates of Core Inflation Coefficient



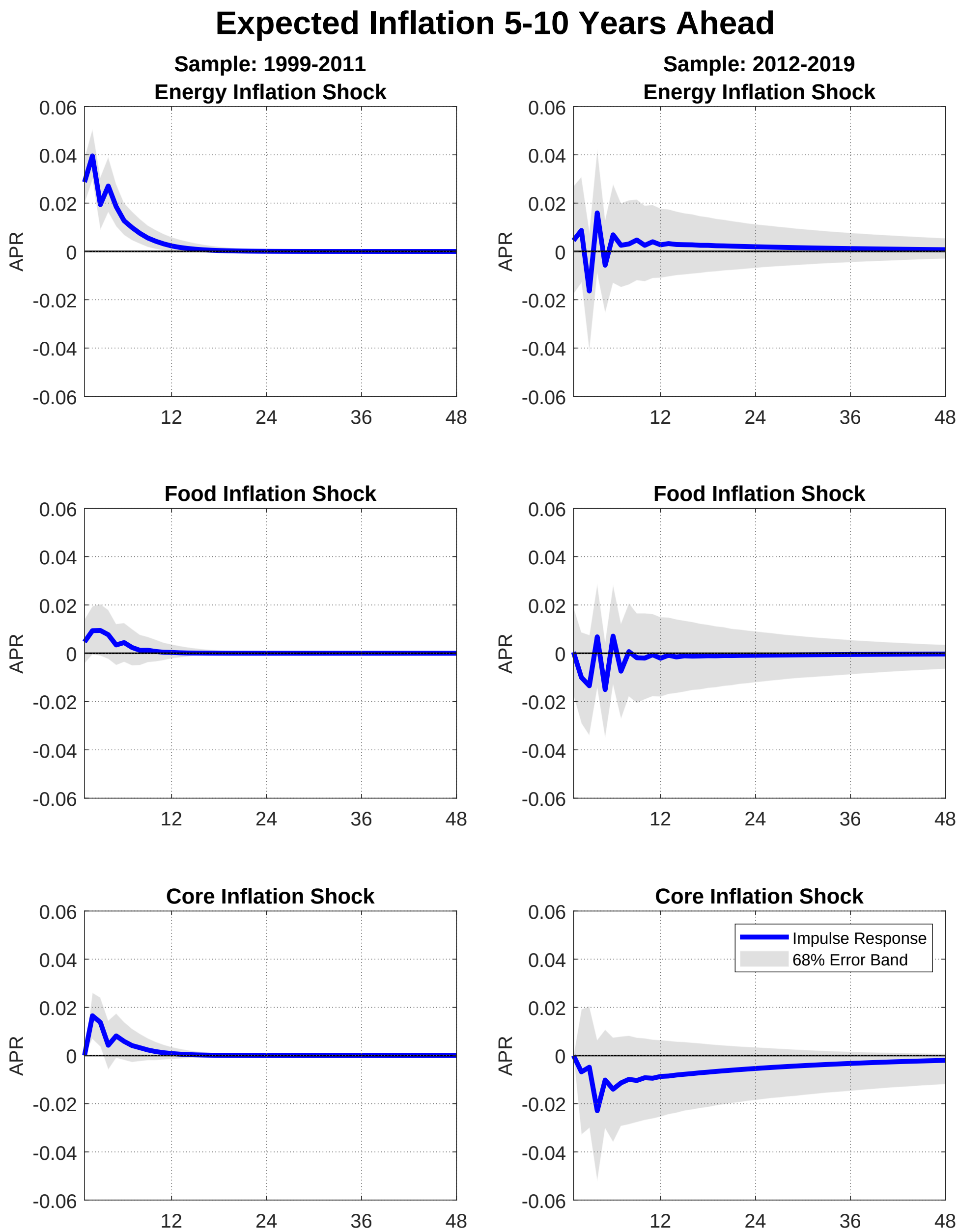
Note: Each panel shows the sequence of estimates of δ^{π} as a function of time. The date on the x-axis denotes the end point of the 10-year rolling sample. The 90% confidence intervals are computed as the point estimate plus or minus 1.645 times the Eicker-White standard error.

Figure 5: 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.

Figure 6: VAR Impulse Responses of Household's Longer-Run Inflation Expectations



Note: The figure shows VAR-estimated impulse responses of household's longer-run inflation expectations in response to various inflationary impulses. The impulse responses in the left column are estimated from 1999-2011. The impulse responses in the right column are estimated from 2012-2019. Each VAR model is estimated on monthly dated comprised of CPI energy inflation, CPI food inflation, longer-run inflation expectations collected from the University of Michigan Survey of Consumers, and CPI inflation. The CPI inflation series enter the VAR in month-over-month inflation rates, annualized. See Section 4.2.2 as well as the appendix for additional details.