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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, first through its Summary of Economic Projections and later through the announcement of a 2 percent target in 2012, better anchored inflation expectations. Moreover, inflation expectations in the United States have remained anchored amid the volatility of the COVID-19 pandemic. In contrast, similar analysis reveals no evidence of anchoring in Japan despite the adoption of a numerical inflation target.

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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. 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."*
>
> – 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 target.

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 a demand-driven decline in the unemployment rate leads to higher inflation. If expectations are not well anchored, then this transitory increase in inflation will raise expectations for inflation far in the future. Forward-looking firms internalize this increase in inflation expectations and set higher prices today. The anchoring of expectations removes these second-round inflationary effects, resulting in a flatter reduced-form Phillips curve.

Using daily 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 communication of a longer-run numerical inflation objective — first through the addition of "Longer-Run" inflation projections to its quarterly Summary of Economic Projections (SEP) in 2009 and later codified by the 2012 announcement — better anchored 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 responded significantly to inflation news contained in the monthly releases of the Consumer Price Index (CPI).

However, once the Federal Reserve began communicating a numerical inflation objective, a battery of econometric tests suggests that far-forward inflation compensation ceased to respond to unanticipated inflation. Moreover, these inflation compensation measures remained unresponsive to the large, upside inflation surprises in the wake of the COVID-19 pandemic, suggesting that longer-term expectations have remained well anchored. 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.

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. 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. One line of this work has studied the time-series properties of survey data, linking the low-frequency changes in the mean or variance of survey forecasts to the conduct of monetary policy (Clark and Davig, 2011; Kozicki and Tinsley, 2012). This research has generally found evidence of lower and more stable long-run survey expectations following the Volcker disinflation. Especially relevant for our study is recent work by Carvalho et al. (2023), which evalutes the degree of anchoring in the US and Japan, among other countries. Carvalho et al. (2023) argue that, in contrast to Japan's experience, professional forecasters' inflation expectations became generally well anchored in the US by the late 1990's.

Despite the aforementioned trends towards improved anchoring since the 1970's, a body of empirical work argued that scope remained to further anchor inflation expectations in the US prior to 2012. For instance, empirical analysis exploiting high-frequency financial market data found potential for further improvements in the anchoring of inflation expectations in the US before 2012. Gürkaynak et al. (2007), Gürkaynak, Levin and Swanson (2010), and Beechey, Johannsen and Levin (2011) conduct cross-country analysis on the behavior of market-based measures of long-run inflation expectations to economic data releases. These papers conclude that, prior to 2012, inflation expectations were generally better anchored in the Euro Area, Canada, the UK and Sweden versus the United States. These authors conjecture that this difference arose because the central banks of these economies publicly adopted a numerical inflation target, whereas the FOMC had not at that time.

Our paper builds on these influential works in several dimensions. First, our results provide external validity to their conclusions regarding the importance of announcing an inflation target. Indeed, we find a reduction in the sensitivity of far-forward measures of inflation compensation to inflation surprises in the United States after the FOMC communicated a numerical inflation objective, suggesting expectations became better anchored. 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 changes in macroeconomic relationships should coincide with the policy change. By applying tests for structural breaks, we are able to provide additional causal evidence on the role that central bank announcements play in shaping inflation dynamics. And, finally, we provide Monte-Carlo evidence that an econometrician can identify structural breaks in the data arising from the anchoring of inflation expectations, even in the small samples that arise from our focus on recent policy changes.

Our paper also relates to a second strand of 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 (Blanchard, 2016; Erceg et al., 2018; Jorgensen and Lansing, 2019; Del Negro et al., 2020; Hazell et al., 2020). These papers attribute at least a portion of this flattening to a change in the conduct of monetary policy, either through more aggressive inflation stabilization or better anchored inflation expectations.

In this paper, we instead focus on more recent changes in FOMC communication about its longer-run inflation objective. We show that, even after the improvements in inflation stabilization that took place in the 1980s and 1990s, the communication of a numerical inflation objective 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 contrast the implications for the reduced-form and structural Phillips curves from better anchoring of inflation expectations. Thus, our work connects to Del Negro, Giannoni and Schorfheide (2015), Bullard (2018), and McLeay and Tenreyro (2020), which highlight that the conduct of monetary policy plays a crucial role in shaping the reduced-form Phillips curve slope despite *stability* in the underlying structural Phillips curve. 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. Consistent with this prediction, we show in US data that a Phillips curve model, augmented with households' one-year-ahead inflation expectations, remains stable before and after 2012.

Finally, we reconcile our findings on the anchoring of inflation expectations and the Phillips curve with the influential Coibion and Gorodnichenko (2015) study. Coibion and Gorodnichenko (2015) argue, like us, that a Phillips curve augmented with one-year-ahead Michigan survey expectations has remained stable in recent decades. However, unlike us, they attribute this stability to unanchored household inflation expectations, particularly in the face of large swings in energy prices.

We resolve this apparent tension by drawing a distinction between near- and longer-term inflation expectations. In particular, we provide time-series evidence that, similar to market-based measures of longer-term inflation expectations (and unlike households' near-term inflation expectations) households' 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 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 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 longer-term inflation expectations appear to 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., 2023), or sufficient credibility (Hausman and Wieland, 2015; De Michelis and Iacoviello, 2016).

3 Testable Predictions of Anchoring Inflation Expectations

We now study the implications of anchoring inflation expectations in a theoretical model which guides our empirical tests for anchoring in the US and Japan. The central feature of our model is the potential for long-term inflation expectations 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 shift longer-term inflation expectations, which captures the notion of drifting or unanchored inflation expectations (e.g. Bernanke, 2007). 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. If $\delta^\pi = 0$, then long-term inflation expectations are fully anchored in the sense that they are invariant to realized inflation. If instead $\delta^\pi > 0$, then inflation expectations are unanchored and drift with realized inflation. π^* represents the central bank's inflation objective, if one has been announced. Our specification also allows for an undefined inflation objective if $\rho^\pi = 1$.[1] Thus, three relevant cases for our study are: **(1)** Drifting longer-run expectations without a formal central bank objective ($\delta^\pi > 0$, $\rho^\pi = 1$), **(2)** drifting expectations despite an announced, or perceived, inflation objective ($\delta^\pi > 0$, $\rho^\pi < 1$), and **(3)** fully anchored inflation expectations ($\delta^\pi = 0$).

In our following analysis, we remain agnostic about the 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 evolution of the Federal Reserve's implicit inflation target. Our 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 changes in inflation. This specification also 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.

3.2 A Model with Nominal Rigidities & 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,

[1] Recent work such as Gáti (2020) and Carvalho et al. (2023) 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.

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.

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, 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 allow the model to generate a downward sloping reduced-form Phillips curve. McLeay and Tenreyro (2020) highlight the difficulties that arise in identifying the Phillips curve when cost-push shocks drive economic fluctuations.

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 $\Pi_t = P_t/P_{t-1}$, $\Pi_t^{LT} = \exp(\pi_t^{LT})$ is the gross rate of long-term inflation expectations, ϕ_P governs the magnitude of the adjustment costs, and Y_t is 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:

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

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. Though not essential for our results, we follow Hall (2005) and Blanchard and Galí (2010) and assume that 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 Nash bargained wage and $\gamma \in (0,1)$ reflects 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 inflation deviations from 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 bank's response to inflation deviations. Similar to Gürkaynak, Sack and Swanson (2005), we assume the central bank responds to deviations from long-term expectations rather than its target Π^*, even if a target has been announced.[3] While we remain agnostic as to why the central bank may behave this way, one possible explanation is that policymakers perceive the costs of driving inflation back to target after a large shock as larger than the benefits, enabling expectations to drift (Gürkaynak, Sack and Swanson, 2005). Finally, as in Ireland (2007), we specify a coefficient of one on lagged policy rates to ensure the policy rule is stationary even when $\rho^\pi = 1$.[4]

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 parameter values designed to mimic key features of the US economy. We take most parameter values from either Ireland (2007) or Leduc and Liu (2016), with a few important exceptions. Beginning with the drifting inflation expectations calibration, we set the degree of anchoring δ^π to align with our high-frequency

[3] If instead the central bank responds to deviations from its target, Equation (1) plays no meaningful role in shaping inflation dynamics.

[4] 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.

estimates before the United States formally adopted its inflation objective.[5] Conditional on this value for δ^π, we then calibrate the persistence of longer-term inflation expectations ρ_π, the degree of price rigidity Φ_P, and the central bank's inflation objective $\pi^* = \log(\Pi^*)$ such that the model matches the constant and slope of the reduced-form Phillips curve regressions and the slope of the expectations-augmented Phillips curve regressions prior to 2012. Our calibrated value of $\phi_P = 270$ implies a coefficient on marginal cost of about 0.02, well within the common range of calibrated values from the literature. A calibration of $\rho_\pi = 0.93$ allows our model to closely match the slope of both the reduced-form and structural Phillips curves during the 1999–2011 sample period. This value is consistent with the central bank having a perceived inflation objective during this time but still allowing fluctuations in longer-term expectations (similar to Erceg and Levin, 2003). To simulate the dynamics under the anchored expectations calibration, we leave all other parameters unchanged but set $\delta^\pi = 0$. 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 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 expectations of inflation over the longer run. Instead, when a credible inflation target is announced, long-run inflation is pinned down solely by the central bank's communication. This prediction is somewhat axiomatic from how we define the notion of anchoring. However, as we discuss momentarily, Equation (1) can not be directly estimated, so we need to draw out the general-equilibrium consequences of this first prediction.

The second model prediction is that the adoption of a credible inflation target mutes

[5] There exists some uncertainty on how to map δ^π from un-annualized monthly CPI surprises to our quarterly model. However, our baseline calibration of δ^π delivers long-term inflation effects close to those from Gürkaynak, Sack and Swanson (2005). In the Appendix, we show that our conclusions are also robust to smaller values of δ^π. The testable predictions that emerge from our model are also robust to alternative calibrations of ρ^π, including setting ρ^π near 1.

the response of inflation to changes in unemployment and thus weakens 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.

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

Our 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 direct 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) motivates a high-frequency event-study approach to recovering δ^π. 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 by regressing the change in far-forward inflation compensation on the surprise or unexpected component of the monthly CPI report. If we find that $\delta^\pi > 0$, such that forward inflation compensation responds to news about current inflation, then this would suggest that inflation expectations are unanchored. In the following sections, we 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 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 4.5, we will return to some of these issues when we contrast the breakdown of reduced-form Phillips curve with potential stability of structural Phillips curve estimates amid the anchoring of long-run inflation expectations.

For now, we 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 induced by anchoring inflation expectations. To generate the model-implied Phillips curves, we simulate the model with the drifting inflation target economy for 156 periods, the same length as our empirical 1999–2011 sample period we use later in the paper. 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 (the length of our 2012–2019 sample period). 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.[6] We repeat this exercise 1000 times and generate small-sample bootstrapped confidence intervals.

Even in a small sample, our simulation evidence suggests an econometrician would be able to detect a break in the reduced-form Phillips curve due to the better anchoring of inflation expectations. The left three columns of Table 2 show the resulting regression coefficients and their associated bootstrapped standard errors. When longer-term inflation expectations drift, the first column of Table 2 illustrates a negative, statistically significant relationship between inflation and unemployment. Once anchored, however, the slope of the Phillips curve becomes much flatter, declining in magnitude by over 50%. Finally, the third column shows results of a Chow (1960) test for a break in the slope coefficient, confirming that

[6]Specifically, we regress year-over-year inflation onto a constant and the unemployment rate in our model-implied Phillips curves. 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.

the flattening in the Phillips curve is statistically significant. In the following sections, we perform this same test to detect whether there was a flattening in the reduced-form Phillips curve following the adoption of numerical inflation targets in the US and Japan.

4 Are Inflation Expectations Anchored in the US?

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.[7] 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. For the US, Bloomberg provides these forecasts and the actual release for the month-over-month percent change in CPI for both headline and core inflation. Using this data, along with the weight of core components in the CPI basket, we construct both a core and an implied food and energy inflation surprise. Our sample period is limited by the availability of TIPS yields, which starts in 1999.

We estimate the following event-study regression to measure δ^{π}, the sensitivity of long-term inflation compensation to news about inflation,

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

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

[7] Our regression model is therefore very similar to the model in Gürkaynak, Levin and Swanson (2010) except we focus exclusively on CPI reports, as prescribed from Equation (8).

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 3 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 3, we 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 an explicit inflation target.[8]

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 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 food and energy commodity markets, bond investors already have considerable information about the food and energy prices ahead of the CPI release.

[8] 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.

4.2 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 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 4 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}. Panel A.1 of Figure 2 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 inflation over the "longer run." In January of 2009, the FOMC added inflation to their longer-run Summary of Economic Projections (SEP) forecasts 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 long 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 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 accomplish the desired degree of anchoring without needing to adopt a Committee-wide longer-run numerical target.[9] 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 support her prediction. In particular, our break date suggests that the modifications to the SEP — which pre-dated the adoption of an explicit 2 percent target in 2012 — played an instrumental role in anchoring inflation expectations.

Rolling-window regressions further suggest that the 2009 changes to the SEP served as a catalyst for anchoring inflation expectations. Panel A.2 of Figure 2 illustrates the time variation in δ^{π} from the estimates of 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 2009, 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 after 2012.[10]

4.3 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 despite no change in the slope of the underlying structural Phillips curve. Empirically examining this prediction demonstrates the 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.

[9] 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 (2022) perform text analysis of FOMC communication and also find that the FOMC's implicit objective for inflation may have resided below 2 percent.

[10] In the Appendix, we show that our findings are robust to different measures and horizons of nominal compensation, as well as removing the financial crisis from our sample period.

The reduced-form Phillips curve in the US appears to have significantly flattened over the last two decades. Table 2 contains the results from a regression of inflation on a constant and the unemployment rate before and after January 2012. We highlight two key findings. First, the slope of the Phillips curve has flattened since 2012. Prior to 2012, Table 2 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. As we discussed 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.

4.4 Accounting for the Flattening of the US Phillip Curve

The right three columns in Table 2 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 attributed to the better anchoring of inflation expectations? To answer this question, we use our theoretical model to quantitatively assess changes in the slope of the reduced-form Phillips curve given the observed changes in the degree of anchoring of inflation expectations. To facilitate comparison with our empirical evidence, we calibrate the model under the drifting inflation target specification such that it generates the same reduced-form Phillips curve slope (-0.19) that we observe in the data during the 1999–2011 period.

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 -0.09 after anchoring. This predicted flattening in the slope coefficient is smaller in magnitude than we observe in the data. However, the 90% model-implied confidence interval of the post-anchoring flattening (0.03, 0.17) contains the empirical estimate of 0.15. Thus, at the 10% level, we cannot reject the null hypothesis that the better of anchoring of inflation expectations explains the observed flattening in the US reduced-form Phillips curve over the past decade.

4.5 Reconciling our Results with Coibion and Gorodnichenko (2015)

Coibion and Gorodnichenko (2015) show that the Phillips curve has remained stable in recent decades once it is augmented with households' near-term inflation expectations. They argue that this stability 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-term inflation expectations because it proxies well the underlying structural Phillips curve in our model, which is invariant to the degree of anchoring. Taking a first-order approximation to the optimal pricing decision of retail firms implies the following relationship:

$$\pi_t - \pi_t^{LT} = \beta \, \mathrm{E}_t \left\{ \pi_{t+1} - \pi_{t+1}^{LT} \right\} + \Psi \, \Xi_t, \tag{10}$$

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 largely results from the failure to account for firm inflation expectations. Intuitively, anchoring removes the second-round inflation effects through expectations that, if unaccounted for, make the reduced-form Phillips curve unstable.

We now use simulations from our theoretical model to demonstrate that a Phillips curve augmented with near-term inflation expectations is stable amid changes in the degree of anchoring. Motivated by the Coibion and Gorodnichenko (2015) specification, we regress year-over-year inflation less one-year ahead inflation expectations on the unemployment rate using model-generated data. The left columns of Table 5 show the resulting regression coefficients when we estimate this expectations-augmented Phillips curve using a sample size of 252 periods with a break in the degree of anchoring in period 157 and repeat this exercise 1000 times. Despite estimating a misspecified Phillips curve relative to the actual data generating process in Equation (10), we find that the slope of the 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. 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 one-year ahead household inflation expectations (measured by the University of Michigan Survey of Consumers) as the dependent variable. The far right columns of Table 5 reveal no evidence of a change in the underlying slope of the Coibion and Gorodnichenko (2015) Phillips curve after January 2012.[11]

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 households' 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, households' 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 households' longer-term inflation expectations since 2012.

Split-sample VAR estimates suggest that households' longer-run inflation expectations (measured again by the University of Michigan Survey of Consumers) have become better anchored since 2012.[12] The left panel of Figure 3 shows that prior to 2012, increases in energy inflation as well as core inflation spilled over to households' longer-run inflation expectations. However, the right panel shows that households' longer-run inflation expectations ceased to respond to those inflationary shocks in the post-2012 sample.[13] Interestingly, in the Appendix we show that if we replace households' longer-term inflation expectations with their near-term (one-year ahead) inflation expectations, we observe significant pass-through to one-year ahead inflation expectations from energy prices in both sample periods. Therefore, by distinguishing between near-term and longer-term inflation expectations, our results can

[11] The Appendix offers further evidence that this expectations-augmented Phillips curve has remained stable since 2012.

[12] See the Appendix for more details on the VAR model and its identification strategy which is based on the timing of survey periods relative to CPI releases.

[13] Our results contrast 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 that households' inflation expectations remain biased, on average, but nevertheless became more stable.

be reconciled with the findings of Coibion and Gorodnichenko (2015).

Aside from the relation to the results in Coibion and Gorodnichenko (2015), this VAR evidence is interesting in its own right because of the debate around whose inflation expectations are most relevant for price dynamics. 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. While we do not to take a stand on this important issue, we note that it is comforting that our results based on market-based expectations carry over to households' longer-term inflation expectations. This household survey evidence also suggests that changes in the behavior of risk premia embedded in inflation compensation are not driving our baseline regression results.[14]

4.6 The Durability of Anchoring in the COVID-19 Period

The onset of the COVID-19 pandemic brought about unprecedented swings in economic conditions, including extreme volatility in employment and inflation dynamics. Anchoring longer-run inflation expectations should distill benefits precisely in such circumstances, as it affords Federal Reserve policymakers the ability to focus on restoring maximum employment without significant costs to their price stability mandate. The COVID-19 period therefore provides an ideal test of the durability of anchoring.

Table 6 extends our inflation compensation regression model to test for a break in the sensitivity of inflation expectations to inflation surprises after December 2019. The first column replicates results from Table 3 for the Jan-2012 through Dec-2019 period and the second column tests for a break in the pass-through of inflation surprises (δ^{π}) after Dec-2019. Despite large and persistent inflation surprises since the start of the pandemic, available evidence suggests that inflation expectations in the US have remained well anchored.

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

[14]The results based on individual SPF respondents in Binder, Janson and Verbrugge (2022) further suggest that, rather than changes in risk premia, expectations have indeed become better anchored since 2012.

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. Governor Kuroda then 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. Given this narrative evidence of several regime 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 a forward rate model:

$$\Delta y_t^{LT} = \delta^0 + \delta^\pi \pi_t^{core} + \delta^f \pi_t^{food} + \varepsilon_t, \tag{11}$$

where Δy_t^{LT} is the one-day change in a 1-year, 9-year nominal 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.[15] 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 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

[15] 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 this forward rate model. In the Appendix, we show that this forward model and our preferred inflation compensation model deliver similar findings for the US. 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 $\delta^\pi = 0$.

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 (11) over time. The Andrews-Quandt and Andrews-Ploberger tests indicate no evidence of significant time variation in δ^π. Panel B.1 of Figure 2 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 show no evidence of structural change. Panel B.2 of Figure 2 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.

5.2 Phillips Curve Estimates in Japan

Corroborating our high-frequency evidence that inflation expectations remain unanchored, we also find evidence of a *stable* reduced-form Phillips curve in Japan despite the adoption of an inflation target. Table 8 shows the output from regressing core inflation on the unemployment rate in 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 can observe some flattening of the reduced-form Phillips curve. 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 early signs of anchoring in Japan. However, at this time, the combined evidence suggest that adopting an inflation objective has yet to anchor inflation expectations in Japan. This conclusion compliments the findings in Hausman and Wieland (2015) and De Michelis and Iacoviello (2016), which suggest that poor credibility on the part of the BOJ may be preventing expectations from anchoring at the BOJ's target.

6 Conclusion

Though central banks regularly 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 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 led to better anchored inflation expectations in the United States, an anchoring that has remained durable through the COVID-19 pandemic.

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	270	Calibrated to Match Phillips Curve
ρ^π	Persistence of Inflation Expectations	0.93	Calibrated to Match Phillips Curve
Π^*	Central Bank Inflation Target	1.008	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 Reduced-Form Phillips Curve Regressions

	Model Simulations			**US Data**		
	Inflation			Core Inflation		
	1999-2011	2012-2019	1999-2019	1999-2011	2012-2019	1999-2019
Constant	3.18^{***}	3.18^{***}	3.18^{***}	3.18^{***}	2.22^{***}	3.18^{***}
	(0.02)	(0.02)	(0.02)	(0.23)	(0.22)	(0.22)
Unemployment Rate	-0.19^{***}	-0.09^{***}	-0.19^{***}	-0.19^{***}	-0.04	-0.19^{***}
	(0.04)	(0.02)	(0.04)	(0.04)	(0.04)	(0.03)
Constant $\times\ \mathcal{I}_{t\geq 2012}$			0.00			-0.96^{***}
			(0.03)			(0.28)
Unemployment Rate $\times\ \mathcal{I}_{t\geq 2012}$			0.10^{**}			0.15^{***}
			(0.04)			(0.05)
Observations	156	96	252	156	96	252
R^2	0.35	0.37	0.35	0.49	0.09	0.45

Note: For the regressions on model-simulated data, we show bootstrapped standard errors. 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. See Sections 3.4.2 and 4.3 for more details. $^{*}p < 0.10,^{**}p < 0.05,^{***}p < 0.01$

Table 3: Chow Test: US Inflation Compensation Model

	Estimation Sample		
Δ 1-Year, 9-Year Forward 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^{***}
	(0.10)	(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 4: Structural Break Tests at an Unknown Date: US Inflation Compensation Model

		Structural Break Test	
Δ 1-Year, 9-Year Forward 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 5: Chow Test: US Expectations-Augmented Phillips Curve Regressions

	Model Simulations			**US Data**		
	Inflation Less Expected Inflation			Core Inflation Less Expected Inflation		
	1999-2011	2012-2019	1999-2019	1999-2011	2012-2019	1999-2019
Constant	-0.00	0.00	-0.00	0.09	0.21	0.09
	(0.04)	(0.02)	(0.04)	(0.41)	(0.21)	(0.41)
Unemployment Rate	-0.18^{***}	-0.12^{***}	-0.18^{***}	-0.18^{***}	-0.20^{***}	-0.18^{**}
	(0.04)	(0.02)	(0.04)	(0.07)	(0.04)	(0.07)
Constant $\times$ $\mathcal{I}_{t\geq 2012}$			0.00			0.12
			(0.05)			(0.46)
Unemployment Rate $\times$ $\mathcal{I}_{t\geq 2012}$			0.05			-0.02
			(0.04)			(0.08)
Observations	156	96	252	156	96	252
R^2	0.31	0.52	0.36	0.17	0.51	0.22

For the regressions on model-simulated data, we show bootstrapped standard errors. 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 one-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. See Section 4.5 for more details. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

Table 6: COVID-19 Chow Test: US Inflation Compensation Model

	Estimation Sample		
Δ 1-Year, 9-Year Forward Breakeven Inflation	2012-2019	2020-2022	2012-2022
Constant	0.00	−0.00	0.00
	(0.00)	(0.01)	(0.00)
Core CPI surprise	−0.03	−0.01	−0.03
	(0.06)	(0.05)	(0.06)
Food & Energy CPI surprise	0.02	−0.08	0.02
	(0.04)	(0.12)	(0.04)
Constant $\times \mathcal{I}_{t \geq 2020}$			−0.00
			(0.01)
Core CPI surprise $\times \mathcal{I}_{t \geq 2020}$			0.02
			(0.08)
Food & Energy CPI surprise $\times \mathcal{I}_{t \geq 2020}$			−0.11
			(0.13)
Observations	95	30	125
R^2	0.01	0.01	0.01

Note: Eicker-White standard errors in parenthesis. $^{*}p < 0.10$,$^{**}p < 0.05$,$^{***}p < 0.01$. The 2020-2022 sample ends in September 2022.

Table 7: Chow Test: Japan Forward Rate Model

	Estimation Sample		
Δ 1-Year, 9-Year Forward 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. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

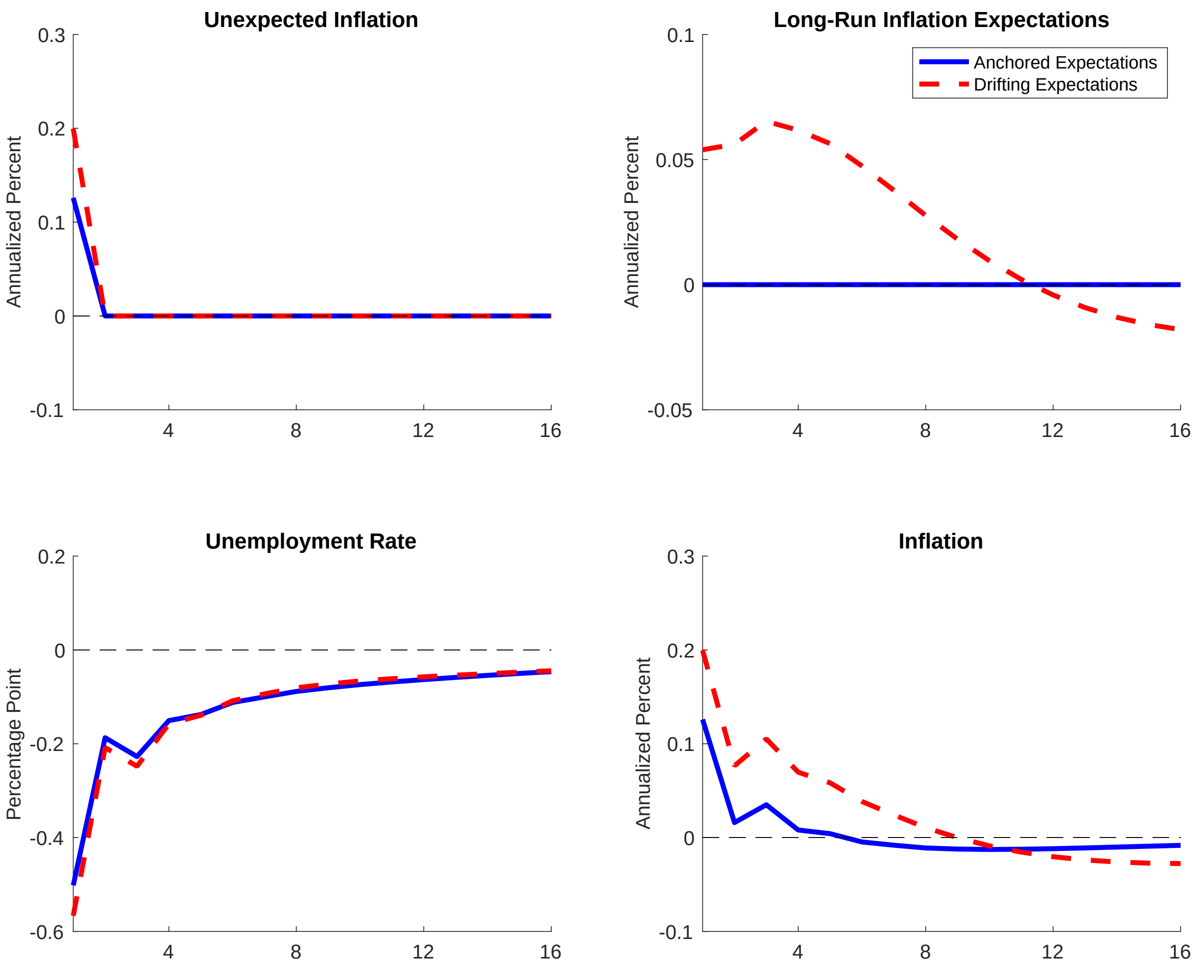
Table 8: Chow Test: Japan Reduced-Form Phillips Curve Regressions

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

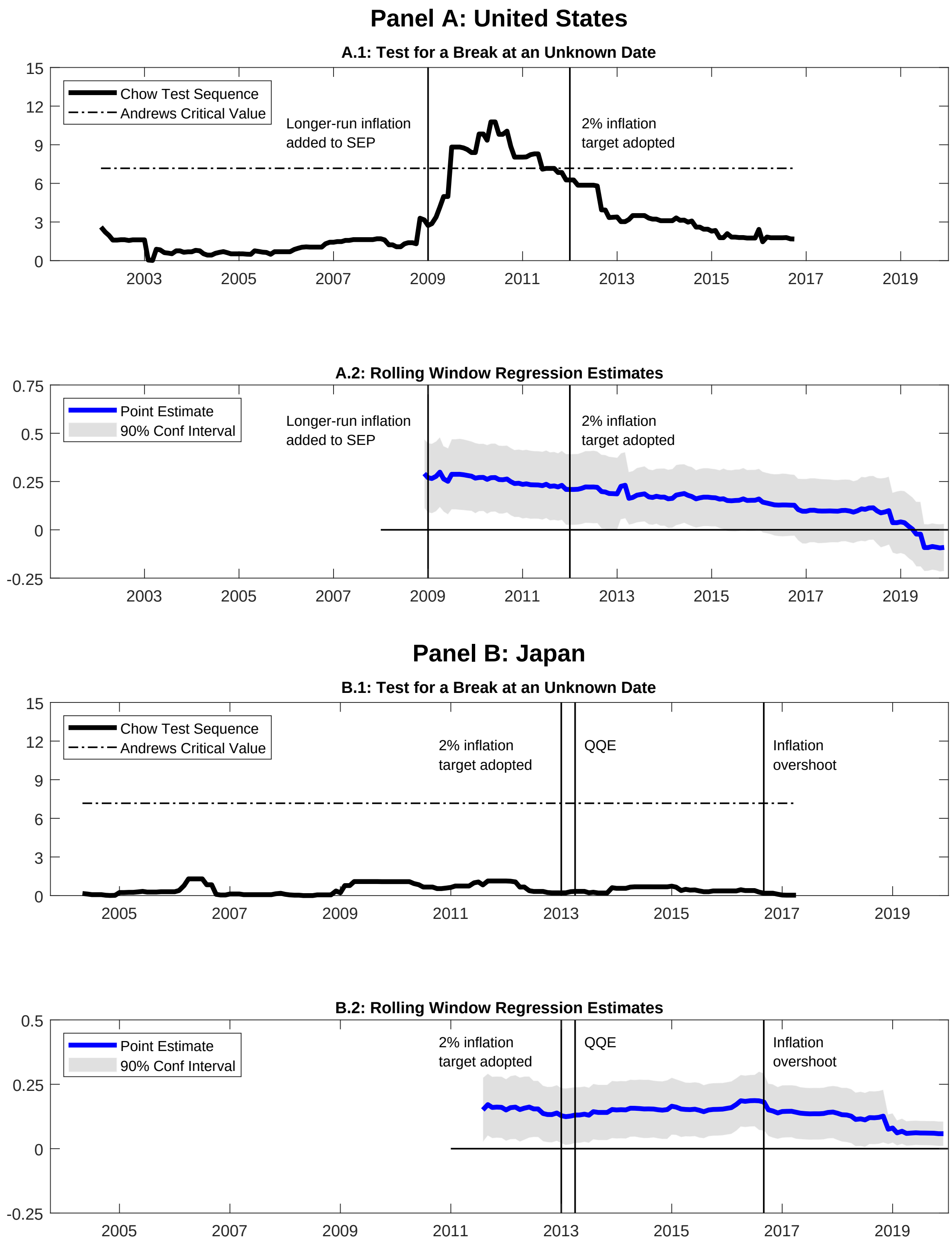
$^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$

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



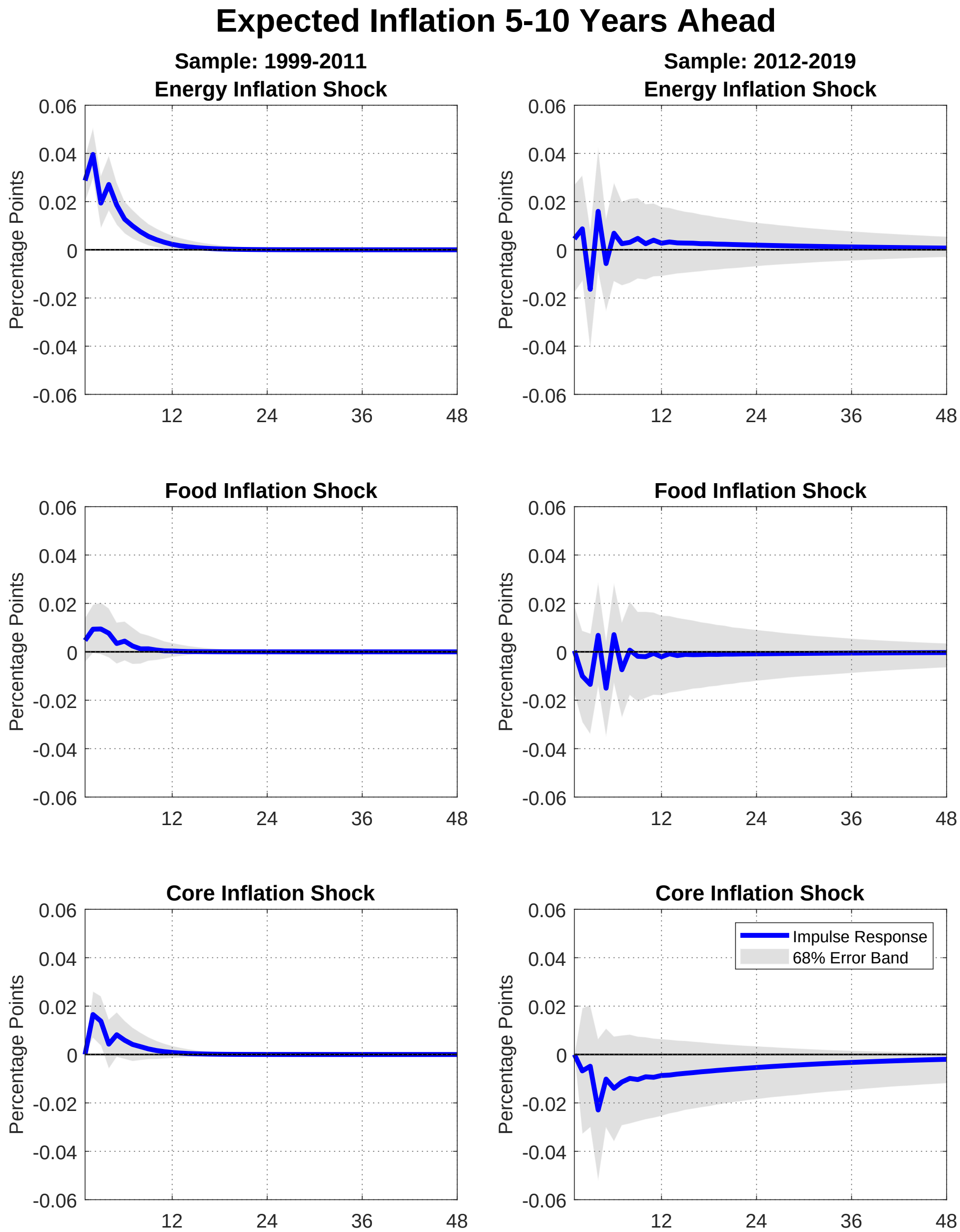
Note: The figure shows the impulse responses in the theoretical model to a one standard deviation aggregate demand (preference) shock under both drifting and anchored coefficient estimates from Table 3. See Section 3.3 for additional details.

Figure 2: Chow Test Sequence & Rolling Window Estimates for Core Inflation Coefficient



Note: Panels A.1 and B.1 show the sequence of Chow test statistics for δ^{π} as a function of candidate break dates. For each country, 15% of the observations on the ends of the sample are not considered for a break. 10% critical values are obtained from Andrews (1993). Panels A.2 and B.2 show the sequence of estimates of δ^{π} as a function of time where the x-axis date denotes the end of the 10-year rolling sample.

Figure 3: VAR Impulse Responses of Households' Longer-Run Inflation Expectations



Note: The figure shows VAR-estimated impulse responses of households' 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 data 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 4.5 as well as the Appendix for additional details.