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Macroeconomic Drivers and the Pricing of Uncertainty, Inflation, and Bonds*

Brandyn Bok[†], Thomas M. Mertens[‡] and John C. Williams[§]

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Abstract

This paper analyzes a new stylized fact: According to financial market prices, the correlation between uncertainty shocks, as measured by changes in the VIX, and changes in break-even inflation rates has declined and turned negative over the past quarter century. It rationalizes this uncertainty-inflation correlation within a standard New Keynesian model with a lower bound on interest rates combined with a decline in the natural rate of interest. With a lower natural rate, the likelihood of the lower bound binding increased and the effects of uncertainty on the economy became more pronounced. In such an environment, increases in uncertainty raise the possibility that the central bank will be unable to eliminate inflation shortfalls following negative demand shocks. As a result, the observed decline in the correlation between uncertainty and inflation expectations emerges. Average-inflation targeting policies can mitigate the longer-run effects of increases in uncertainty on the real economy.

JEL Classification System: E52

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I Introduction

Over the past quarter century, there has been a marked downward trend in the correlation between changes in market-based measures of uncertainty and changes in expected inflation in the United States. At the beginning of this century, this correlation was essentially zero, but it has since turned negative, averaging close to -0.5 in recent years. This emergence of a negative relationship between uncertainty shocks and expected inflation implies that uncertainty shocks now have a more pronounced effect on the economy than two decades ago and thereby pose greater challenges for the conduct of monetary policy.

In this paper, we analyze the theoretical implications of a decline in the natural rate of interest on the correlation between uncertainty shocks and changes in expectations of future inflation and test these predictions empirically using financial market data. In addition to documenting this stylized fact, the main contributions of this paper are two-fold. The first is a theoretical analysis of the effects of uncertainty shocks on expected inflation and interest rates within a standard New Keynesian model. We highlight how these relationships are affected by the lower bound on interest rates. The second is an empirical analysis of the predictions of the theory, which indicate that the decline in the natural real rate of interest to low levels has contributed to the emergence of the negative relationship between uncertainty and inflation expectations evident in the data.

In the theoretical section of the paper, we use a standard New Keynesian macroeconomic model that incorporates a lower bound on nominal interest rates to derive analytically the theoretical implications of a change in the natural rate of interest on the joint behavior of uncertainty and expectations of inflation and interest rates. We prove that the presence of the lower bound implies the existence of two distinct ergodic distributions for endogenous variables that are delineated by the probability of a binding lower bound, as in Mertens and Williams (2021). We refer to the ergodic distribution associated with a lower probability of being at the lower bound as the target equilibrium, and the other distribution as the liquidity trap equilibrium.

In the model without a lower bound on interest rates, a mean-preserving increase in uncertainty has no systematic effect on the unconditional means of expected inflation and interest rates. In contrast, in the presence of an occasionally binding lower bound on interest rates, a mean-preserving spread to the distribution of demand shocks affects the unconditional mean rate of inflation, decreasing it in the target equilibrium and increasing it in the liquidity trap equilibrium. In the case of the target equilibrium, an increase in uncertainty implies that policy is constrained in a larger region of the state space. This asymmetric constraint on the ability to respond to negative shocks results in a reduction in the unconditional means of

expected inflation and interest rates. In the case of the liquidity trap equilibrium, an increase in uncertainty contracts the region in which the lower bound binds and thereby increases the unconditional means of inflation and interest rates.

In the model, a decline in the natural rate of interest leads to the observed effects of changes in uncertainty on expectations of inflation, at least in the tails of general distributions. Under additional assumptions such as Gaussian distributions, these results hold globally. We show that a decline in the natural rate of interest increases the magnitude of the relationship between uncertainty and expected inflation.

Second, we empirically assess these theoretical predictions using data from the United States. Therefore, we construct rolling-window correlations between changes in market-based measures of uncertainty, such as the VIX, and changes in expectations of far-forward future inflation and interest rates. The New Keynesian model predicts that, analogously to the correlation for inflation expectations, the correlation between uncertainty shocks and changes in expected interest rates should turn negative as well. Indeed, this correlation fell from close to zero to around -0.5 over the same time period. These changes took place as estimates of the natural rate of interest declined and the incidence of monetary policy being near or at the lower bound increased.

To test this prediction, we regress the correlations for interest rate and inflation expectations on estimates of the natural rate of interest and find that the empirical evidence is qualitatively consistent with the U.S. economy being in the target equilibrium, in line with evidence reported in Mertens and Williams (2021). We furthermore establish the robustness of the results in Figure 1. We show that we obtain consistent results at daily and monthly frequencies, removing term premia, or using alternative measures of uncertainty shocks and inflation expectations. For the latter, we show that our results are robust to using survey-based measures of expected inflation from the Survey of Professional Forecasters and to using both an aggregate econometric measure of economic financial uncertainty and a news-based economic policy uncertainty index for uncertainty shocks (see Ludvigson et al. (2021) and Baker et al. (2016)).

Finally, we assess the implications of greater uncertainty about aggregate demand for the design of monetary policy. We show that, in the target equilibrium, the benefits of average-inflation targeting relative to standard inflation targeting increase with greater macroeconomic uncertainty. We implement average-inflation targeting through an adjustment in the level of the policy rule for nominal interest rates. Lowering the intercept leads inflation to overshoot the target inflation rate during times when policy is unconstrained. This policy framework thereby makes up for shortfalls in inflation at times when the lower bound on interest

rates is strictly binding.

We highlight one caveat to our analysis. To maintain analytical tractability, we do not model time variation in correlations explicitly. Instead, we study changes in the derivative of expectations with respect to uncertainty when there is an exogenous shift in the natural real rate of interest. Therefore, we do not generalize these findings to models of monetary policy with richer dynamics nor address the potential endogeneity of the decline in the natural real rate.

Our analysis relates to several strands of the literature. First, it relates to models that incorporate a lower bound on interest rates into a New Keynesian economy, such as Fuhrer and Madigan (1997), Eggertsson and Woodford (2003), Adam and Billi (2006), Campbell et al. (2012), and Cochrane (2018). Papers in this literature that deal with the multiplicity of equilibria include Benhabib et al. (2001), Mendes (2011), Hills et al. (2016), Nie and Roulleau-Pasdeloup (2021), Cuba-Borda and Singh (2022), and Bilbiie (Forthcoming). Our analysis relates to the work by Swanson and Williams (2014) and Mertens and Williams (2021) in that it measures the effects of the lower bound on interest rates from financial market data. Second, it relates to a literature that studies changing correlations in financial markets and links them to macroeconomic outcomes. Baele et al. (2010) and Duffee (2021) document changing correlations between stocks and bonds. Campbell et al. (2020) point out macroeconomic drivers behind this phenomenon. And Bilal (2017) argues that the lower bound on interest rates might be behind this finding. Third, this paper contributes to the large and fast-growing literature on uncertainty shocks and their effects on the economy. Seminal contributions on the measurement include Jurado et al. (2015) and Baker et al. (2016). Leduc and Liu (2016) study the effects of uncertainty shocks in a New Keynesian economy while Basu and Bundick (2017), Plante et al. (2017) and Nakata (2017) take the lower bound on interest rates into account.

II Stylized Fact

This section discusses a new stylized fact: Since the mid-2000s, the correlation between uncertainty shocks and changes in long-run inflation expectations has turned negative from a level of close to zero. While we purposefully keep this section brief, Section V discusses the robustness of these findings.

To demonstrate these patterns, we measure uncertainty shocks as changes in expected volatility in the stock market via the VIX volatility index. For expected inflation, we use market-implied breakeven inflation (BEI) rates, measured as the difference between yields of nominal and inflation-protected Treasury bonds. We use five-year inflation expectations five years ahead to focus on expectations of inflation in the longer term.

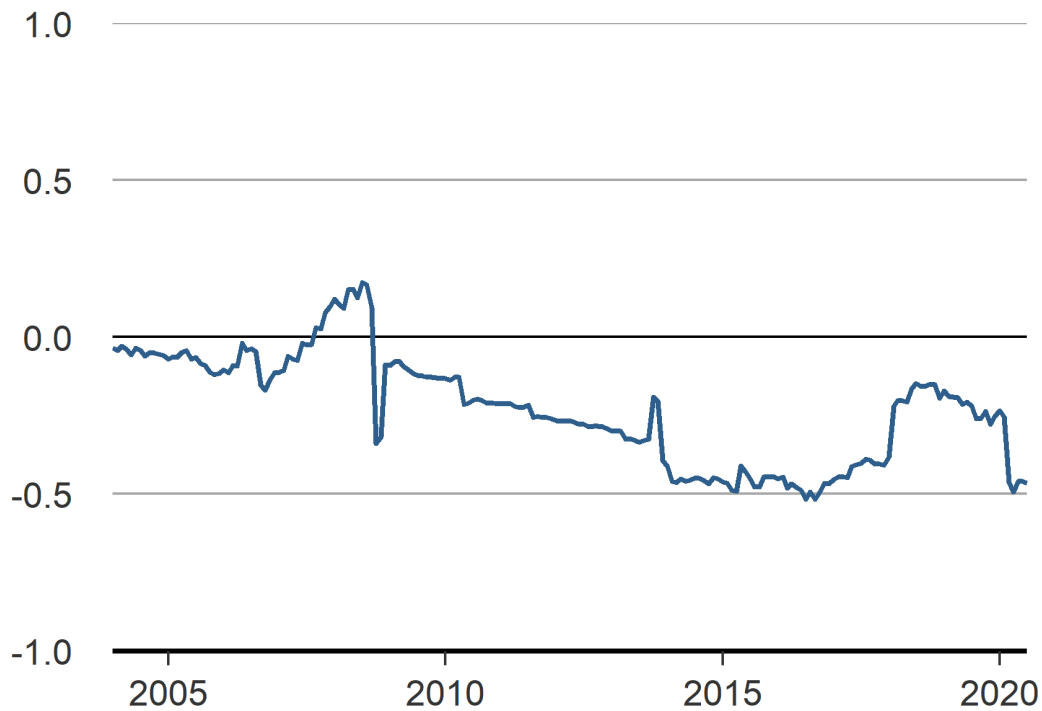


Figure 1: Five-year trailing-window correlation of monthly changes in the VIX with monthly changes in the five-year, five-year-forward breakeven inflation (BEI) rate. The underlying variables are monthly averages of daily market-close values.

We retrieve daily time series on each of these variables from Bloomberg.¹ To get monthly observations, we construct changes in monthly averages of the daily market-close values and compute the correlation between these monthly changes in the VIX and changes in inflation expectations for five-year rolling windows. The overall sample for the monthly changes is restricted by the introduction of Treasury inflation-protected securities (TIPS) in February of 1999, such that the five-year trailing-window correlations begin in February 2004. We end the sample with the Federal Reserve’s announcement of the average-inflation targeting framework.

Figure 1 plots the correlation between these measures of uncertainty shocks and changes in inflation expectations at the monthly frequency. The decline in this correlation occurred during the mid-2000s and accelerated as the federal funds rate was lowered to the effective lower bound. It remained highly negative throughout the initial lower bound period and increased after liftoff from the lower bound. After the onset of the COVID-19 pandemic and the ensuing recession, the correlation turned highly negative again.

We interpret an increase in the VIX as reflecting an increase in uncertainty. There is a literature documenting a negative relationship between increases in the VIX and stock prices. One possible channel for

¹Appendix B summarizes the sources and details of all data used in this paper.

this connection might be that the increase in uncertainty leads to an increased risk premium, rather than there being a negative correlation between the level and volatility of shocks. However, the mechanism in our model would still apply if the mean and volatility of shocks are negatively correlated. As a robustness check, we also analyze other measures of uncertainty.

In the following section, we provide an explanation for this stylized fact. We argue that the lower bound on interest rates can lead to the observed pattern within a basic New Keynesian model. In such an environment, increases in uncertainty raise the possibility that monetary policy is unable to eliminate inflation shortfalls.

III Model

In this section we study a simple version of the standard log-linearized three-equation New Keynesian model, as in, e.g., Woodford (2003), to which we add a lower bound on interest rates. For simplicity and analytical tractability, our model features only demand shocks that are independent and identically distributed (i.i.d.) over time.

The Phillips curve describes the evolution of inflation π_t

$$\pi_t = \kappa x_t + \beta \mathbb{E}_t \pi_{t+1}, \quad (1)$$

where $\kappa > 0$ is the sensitivity of inflation to changes in the output gap, $\beta \in (0, 1)$ is the discount factor, and inflation expectations are based on the information set available at time t .

The IS relation describes the behavior of the output gap x_t

$$x_t = \epsilon_t - \alpha(i_t - \mathbb{E}_t \pi_{t+1} - r^*) + \mathbb{E}_t x_{t+1}, \quad (2)$$

where $\alpha > 0$ is the responsiveness of the output gap to deviations of the real interest rate from its long-run natural rate, r^* .

The demand shock ϵ_t is assumed to be i.i.d. over time with zero mean and follows an arbitrary distribution that can be described by a probability density function $g(\cdot)$ with infinite support.² We denote the corresponding cumulative density function by $G(\cdot)$.

The central bank sets the nominal rate i_t optimally under discretion but is constrained by a lower bound

²While the results hold more generally, we work with an infinite support of the shock to avoid the distinction of the cutoff for a binding lower bound lying within or outside the support of the distribution.

$i^{LB} < r^*$ that lies below the natural real rate of interest. It aims to minimize a loss function \mathcal{L} that penalizes output gaps as well as deviations of inflation from its target, which is normalized to zero,³

$$\min_{i_t \geq i^{LB}} \mathcal{L} = \min_{i_t \geq i^{LB}} (1 - \beta) \mathbb{E}_0 \left[\sum_{t=0}^{\infty} \beta^t (\pi_t^2 + \lambda x_t^2) \right], \quad (3)$$

where $\lambda \geq 0$ is the relative weight the central bank places on the output gap relative to deviations of inflation from target. We focus on optimal monetary policy under discretion, when the central bank lacks the ability to commit to future actions and thus takes inflation expectations as given. The central bank and private sector are assumed to have full knowledge of the model, including the distribution of the shock processes.

The state space of the system consists only of the current realization of the shock ϵ_t . Due to the shock being purely temporary, its current realization has no predictive content for future realizations of the shock process. As a result, conditional expectations of future variables based on the current information set coincide with unconditional expectations. This is reflected in the proceeding analysis with the conditional time subscript t dropped from the expectations operator.

It is convenient to define a shadow rate \mathbf{i}_t as

$$\mathbf{i}_t = \theta + r^* + \psi \mathbb{E} \pi_{t+1} + \frac{1}{\alpha} \epsilon_t,$$

which, for $\theta = 0$, is the optimal policy rate under discretion without the lower bound constraint. In a later section, we use the parameter θ to introduce different monetary policy frameworks.

In the presence of a lower bound, the optimal discretionary policy is to set the interest rate equal to the shadow rate whenever feasible and to the lower bound otherwise:

$$i_t = \max\{\mathbf{i}_t, i^{LB}\} = \mathbf{i}_t + i_t^\Delta.$$

We furthermore define the interest rate wedge i_t^Δ as the difference between the policy rate and the shadow rate as in the second part of the above equation. Since the policy rate i_t is either equal to the shadow rate \mathbf{i}_t or set above the shadow rate at i^{LB} when the shadow rate falls below the lower bound, the interest rate wedge i_t^Δ is always nonnegative.

Equipped with this notation, we rewrite the model equations in terms of the interest rate wedge by

³More generally, the same results follow for nonzero inflation target, with π_t then representing deviations in inflation from target.

substituting $i_t + i_t^\Delta$ for the nominal rate i_t . The system of equations (1) and (2) then becomes

$$\pi_t = \frac{\lambda}{\kappa^2 + \lambda} \beta \mathbb{E}_t \pi_{t+1} - \alpha \kappa (i_t^\Delta + \theta) \quad (4)$$

and

$$x_t = -\frac{\kappa}{\kappa^2 + \lambda} \beta \mathbb{E}_t \pi_{t+1} - \alpha (i_t^\Delta + \theta)$$

where $\psi = 1 + \frac{1}{\alpha \kappa} - \frac{\beta \lambda}{\alpha \kappa (\kappa^2 + \lambda)} > 1$.

In these equations for inflation and the output gap defined in terms of the interest rate wedge, the first terms reflect what holds when the lower bound is not binding and when $\theta = 0$, as in the standard New Keynesian model (see Clarida et al. (1999)). The demand shock is fully offset and there is no tradeoff between stabilizing inflation and output, i.e., “divine coincidence” holds. However, the final term in each of these equations shows that a binding lower bound leads, *ceteris paribus*, to lower inflation and a negative output gap. Monetary policy can raise inflation on average by lowering interest rates on average, that is, by setting θ to be negative.

III.A Model Solution

Absent a lower bound, equation (4) shows that there is a unique equilibrium. Taking unconditional expectations on both sides results in an equilibrium where inflation is at target, $\mathbb{E} \pi_t = 0$, and the nominal interest rate equals the long-run natural rate, $i_t = r^*$.

The presence of a lower bound on interest rates in the model gives rise to the possibility of a second equilibrium. To see this, we split equation (4) into parts depending on whether the lower bound binds in the current period or not

$$\pi_t = \begin{cases} \frac{\beta \lambda}{\kappa^2 + \lambda} \mathbb{E}_t \pi_{t+1} - \alpha \kappa \theta & \text{if } i_t \geq i^{LB} \\ \kappa \epsilon_t - \alpha \kappa (i^{LB} - \theta - r^*) + (1 - \alpha \kappa) \mathbb{E}_t \pi_{t+1} & \text{if } i_t < i^{LB}. \end{cases} \quad (5)$$

In a deterministic version of our model, each part of equation (5) delivers one equilibrium. The deterministic steady state stemming from the upper equation is characterized by inflation at its target of zero ($\pi = 0$) and the nominal interest rate equal to the long-run natural real rate ($i_t = r^*$). The lower equation gives rise to an equilibrium in which monetary policy is constrained such that the interest rate is permanently at the

lower bound, $i_t = i_t^{LB}$, as in Benhabib et al. (2001). The corresponding steady-state inflation is equal to the difference between the lower bound and the long-run natural real rate, $\pi = i^{LB} - r^* < 0$, and thus below its target of zero.

Characterizing the equilibrium in the stochastic environment with demand shocks requires solving for expectations of future inflation, the output gap, and the interest rate wedge. To this end, we recognize that these expectations are linked in the model via

$$\mathbb{E}\pi_{t+1} = \frac{\kappa}{1-\beta}\mathbb{E}x_{t+1} = -\frac{1}{\psi-1}(\mathbb{E}i_{t+1}^\Delta + \theta). \quad (6)$$

Furthermore, the Fisher equation holds

$$\mathbb{E}\pi_t = \mathbb{E}i_t - r^*.$$

Appendix A contains details on the derivation of these equations.

Equation (6) and the Fisher equation imply that it is sufficient to characterize one of the expectations. In the following analysis, we focus on solving for the expected interest rate wedge $\mathbb{E}i^\Delta$, which translates into inflation expectations and, by the Fisher equation, into expected interest rates.

Computing the expected interest rate wedge results in

$$\mathbb{E}i^\Delta = \int_{-\infty}^{\bar{\epsilon}^{LB}} (i^{LB} - i_t)g(\epsilon_t)d\epsilon_t = \frac{1}{\alpha}\mathcal{G}(\bar{\epsilon}^{LB}), \quad (7)$$

where

$$\bar{\epsilon}^{LB} = \alpha(i^{LB} - \theta - r^* + \frac{\psi}{\psi-1}(\mathbb{E}i^\Delta + \theta)) \quad (8)$$

defines the cutoff for the demand shock below which the lower bound binds. $\mathcal{G}(\epsilon_t)$ is the super-cumulative distribution function of the demand shock ϵ_t , which is obtained by integrating the cumulative distribution function. This is to say that $\mathcal{G}'(\bar{\epsilon}) = G(\bar{\epsilon})$ where $G(\cdot)$ is the cumulative distribution function associated with the probability density function $g(\cdot)$.

Using the definition of the cutoff, we plug in for the expected interest rate wedge using equation (7) to get the equilibrium correspondence

$$\mathcal{G}(\bar{\epsilon}^{LB}) = \frac{\psi-1}{\psi}\alpha\left(\frac{1}{\alpha}\bar{\epsilon}^{LB} - i^{LB} - \frac{1}{\psi-1}\theta + r^*\right). \quad (9)$$

Solving this equation for the cutoff value delivers the solution to the model as the cutoff determines expected interest rate wedges via equation (7). Appendix A.A contains the derivations and more details.

According to equation (9), an equilibrium of the model occurs when the super-cumulative distribution function equals the right-hand side of the equation. This right-hand side is linear in the cutoff $\bar{\epsilon}^{LB}$ and depicted in the black line in Figure 2.

Since one can obtain this function by integrating over the cumulative distribution function, it becomes clear that the function converges to zero for $\bar{\epsilon}^{LB} \rightarrow 0$ and to infinity for $\bar{\epsilon}^{LB} \rightarrow \infty$. Furthermore, since its second derivative is the probability density function, it is nonnegative and \mathcal{G} is convex.

Due to the convexity of the super-cumulative distribution function, there can be either zero, one (in a knife-edge case), or two equilibria. As we saw above, the deterministic case always features two equilibria. We label the equilibrium to the left as the *target equilibrium*. In the deterministic case, it turns into the equilibrium in which the lower bound does not bind. With uncertainty, however, the lower bound becomes occasionally binding. The equilibrium to the right is a *liquidity trap equilibrium* and is associated with the deterministic equilibrium in which the lower bound binds. In the stochastic version, there can also be an occasionally binding constraint.

The probability with which the lower bound binds is

$$P^{LB} = G(\bar{\epsilon}^{LB}) = \int_{-\infty}^{\bar{\epsilon}^{LB}} g(\epsilon) d\epsilon. \quad (10)$$

With this definition, we get the following lemma.

Lemma 1 *Depending on the amount of uncertainty, there exist two, one, or zero equilibria in this economy. Furthermore, in the case of two distinct equilibria, there is a threshold $\frac{\psi-1}{\psi}$ such that the probability of a binding lower bound is always below the cutoff in the target equilibrium and always above the cutoff in the liquidity trap equilibrium.*

The proof is in Appendix A.A.2.

In the presence of uncertainty, the lower bound on interest rates is occasionally binding in both equilibria. When there are two equilibria, the probability of a binding lower bound in the target equilibrium always falls below the cutoff value whereas it exceeds it in the liquidity trap equilibrium. In the knife-edge case of a unique equilibrium, the probability of a binding lower bound is $\frac{\psi-1}{\psi}$. We denote the corresponding cutoff by $\bar{\epsilon}^*$.

As an illustration of the solution of the model, we consider the case of a normal distribution for the shock.

There, the super-cumulative function takes the form

$$\mathcal{G}^{\text{norm}}(\bar{\epsilon}^{LB}) = \bar{\epsilon}^{LB} \Phi\left(\frac{\bar{\epsilon}^{LB}}{\sigma}\right) + \phi\left(\frac{\bar{\epsilon}^{LB}}{\sigma}\right) \sigma,$$

where $\Phi(\cdot)$ denotes the cumulative distribution function (c.d.f.) of the standard normal distribution, $\phi(\cdot)$ the corresponding probability density function (p.d.f.), and σ is the standard deviation of the demand shock.

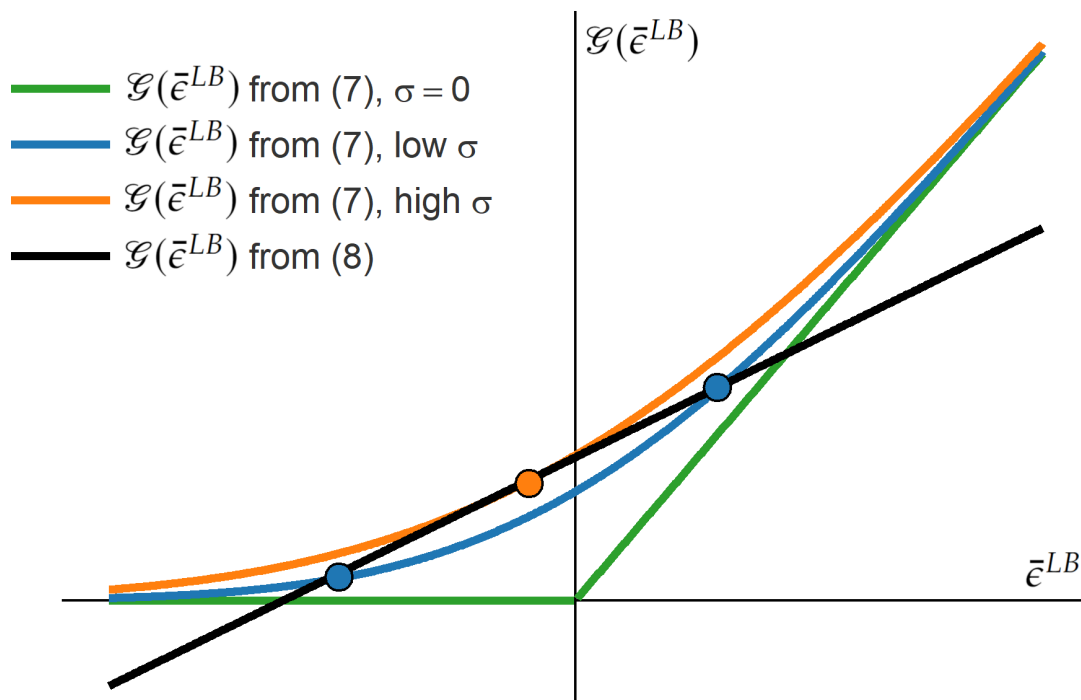


Figure 2: Equilibrium correspondence. Solid straight black line plots $\mathcal{G}(\bar{\epsilon}^{LB})$ as a linear function of $\bar{\epsilon}^{LB}$, as given by equation (9). The other lines plot the super-cumulative distribution function $\mathcal{G}^{\text{norm}}(\bar{\epsilon}^{LB})$ for the case of a normal distribution, with mean zero and varying standard deviations (σ). The intersection points with the solid black line indicate equilibria for these various values of σ under normality.

Figure 2 depicts this super-cumulative distribution function and the condition leading to an equilibrium. The blue line shows the super-cumulative distribution for a normally distributed demand shock as a function of the cutoff $\bar{\epsilon}^{LB}$.

As can be seen in Figure 2, the equilibrium cutoffs and thus expected interest rate wedges are closer to each other in the case of uncertainty compared to the deterministic counterpart. In the deterministic target equilibrium, the lower bound on interest rates does not bind. In the case with uncertainty, however, there is a cutoff for the demand shock below which the central bank becomes constrained by the lower bound. This is because the central bank aims at offsetting the demand shock and lower realizations cause it to cut interest

rates. With the demand shock being sufficiently low, the shadow rate falls below the lower bound.

In the liquidity trap equilibrium, the logic is reversed. The lower bound always binds in the deterministic steady state. Adding uncertainty results in situations where the demand shock is so high that the shadow rate is above the lower bound and the central bank finds itself unconstrained. Since there are now situations in which the central bank stabilizes the economy, average inflation—and thus inflation expectations—rise.

Further increasing uncertainty eventually leads the two equilibria to be identical. Beyond this knife-edge case, any further increase in uncertainty leads to non-existence of equilibria.

To summarize, without a lower bound on interest rates, the New Keynesian model results in a unique equilibrium where inflation expectations are anchored at the target level. Introducing a lower bound adds a liquidity trap equilibrium that is characterized by a higher probability of a binding lower bound compared to the target equilibrium. We now turn to analyzing the effects of changes in uncertainty on the economy.

III.B The effects of uncertainty and demand shocks

This section studies the impact of an increase in uncertainty on the economy. It shows that the lower bound plays a crucial role: Without it, an increase in uncertainty has no effect on inflation and output. With a lower bound, higher uncertainty increases the region of the state space in which the lower bound is binding in the target equilibrium while it shrinks the region in a liquidity trap equilibrium.

In the absence of a lower bound, uncertainty has no effect on the economy since demand shocks can be fully offset. Optimal interest rate policy shields inflation and output from the effect of demand shocks. More volatile demand leads to more volatile interest rates but does not transmit to prices or the real economy.

The lower bound on interest rates changes these predictions. We show that the expected interest rate wedge increases in the target equilibrium and falls in a liquidity trap equilibrium. We demonstrate these effects for general distributions with unbounded support. Therefore, we show that a mean-preserving spread alters the probability of a binding lower bound and, in turn, expected interest rate wedges and inflation expectations.

Let $\varphi_t = \epsilon_t + s\nu_t$ be a series of mean-preserving spreads of ϵ_t , indexed by $s \geq 0$, where $\mathbb{E}[\nu_t|\epsilon_t] = 0$. A larger value of s implies a higher variance of the mean-preserving spread. Denote the super-cumulative distribution function of φ_t by $\mathcal{G}(s, \varphi_t)$. Then we get the following lemma.⁴

⁴For a proof, see Rothschild and Stiglitz (1970).

Lemma 2 (Mean-preserving spread and second-order stochastic dominance)

The following statements are equivalent:

1. φ_t is a mean-preserving spread of ϵ_t with p.d.f. $g(s, \varphi_t)$ and c.d.f. $G(s, \varphi_t)$.
2. $G(s, \varphi_t)$ second-order stochastically dominates $G(\epsilon_t)$ for all $s > 0$.
3. $\mathcal{G}(s, \varphi) \geq \mathcal{G}(\varphi)$ for all φ and $s > 0$.
4. There exist $\underline{\varphi}$ and $\bar{\varphi}$ such that $G(s, \varphi_t) \geq G(\varphi_t)$ for all $\varphi_t \leq \underline{\varphi}$ and $G(s, \varphi_t) \leq G(\varphi_t)$ for all $\varphi_t \geq \bar{\varphi}$.

A general increase in uncertainty through a mean-preserving spread pushes mass into the tails. In that case, the distribution for the more volatile demand shock is said to second-order stochastically dominate the distribution without the mean-preserving spread. As a consequence, the cumulative distribution function rises in the left tail, thereby increasing its integral, the super-cumulative distribution function, as shown in Lemma 2.

Figure 2 depicts these effects for the case of the normal distribution. While the linear relationship between expected interest rate wedges and the cutoff $\bar{\epsilon}^{LB}$ in the black line is unaffected by an increase in uncertainty, the super-cumulative distribution function increases and the two equilibria move closer together. Further increases in uncertainty eventually lead to a unique equilibrium and non-existence for yet higher levels of uncertainty.

From this picture, it is also evident that the cutoff $\bar{\epsilon}^{LB}$ rises with uncertainty in the target equilibrium (the point of intersection to the left) and, as a result, the lower bound binds over a larger portion of the support for the demand shock. The opposite prediction holds in a liquidity trap equilibrium. Proposition 1 summarizes these results.

Proposition 1 (Effects of uncertainty)

A mean-preserving spread has no effect on inflation expectations absent the lower bound; in the presence of a lower bound, it raises (lowers) the expected interest rate wedge in a target (liquidity trap) equilibrium.

We analyze the derivative of inflation expectations with respect to an increase in the mean-preserving spread

$$\frac{d\mathbb{E}i^\Delta}{ds} = \frac{1}{\alpha} \frac{\psi - 1}{\psi} \frac{\mathcal{G}'_1(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - P^{LB}}.$$

We interpret this expression as the model's analogue of the correlation of changes in expectations due to an uncertainty shock, as measured by a change in the mean-preserving spread. This expression formalizes the discussion of the effects of changes in uncertainty. When there is no lower bound, the function $\mathcal{G}(\epsilon_t)$ is zero everywhere for all values of s . Thus, the derivative of the wedge vanishes as well, and changes in uncertainty have no effects on the inflation expectations absent the lower bound.

With a lower bound on interest rates, the sign of the effects of changing uncertainty on expectation is determined by the equilibrium the economy is in. The numerator is positive, $\mathcal{G}_1(s, \epsilon_t) > 0$, according to Lemma 2. We furthermore established that the lower bound probability is below the threshold $(\psi - 1)/\psi$ when the economy is in a target equilibrium and above in a liquidity trap equilibrium. Consequently, increases in uncertainty raise interest rate wedges, and thus lower inflation expectations, in the target equilibrium and have the opposite effect in the liquidity trap equilibrium.

The impact of changes in uncertainty on the probability of a binding lower bound, however, is more subtle. While the region of the state space expands (contracts) with higher uncertainty in the target (liquidity trap) equilibrium, the probability of a binding lower bound does not always change in the same direction. The reason is that there can be two opposing forces. In the target equilibrium, the cutoff $\bar{\epsilon}^{LB}$ increases with uncertainty, suggesting that the probability of a binding lower bound should increase. However, a mean-preserving spread has a direct effect on the distribution function that can result in a lower value at the cutoff. To see this, we differentiate the definition of the lower bound probability with respect to s ⁵

$$\frac{dP^{LB}}{ds} = G_1(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{\mathcal{G}_1(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - G(s, \bar{\epsilon}^{LB})}. \quad (11)$$

The first term in the final expression can be responsible for the lower bound probability and the cutoff moving in opposite directions.

Lemma 3 establishes that, in the tails, the lower bound probability and the cutoff always move in the same direction in response to increases in uncertainty.

Lemma 3 *There are values $\underline{\epsilon}$ and $\bar{\epsilon}$ such that an increase in uncertainty raises the probability of a binding lower bound whenever $\bar{\epsilon}^{LB} < \underline{\epsilon}$ in the target equilibrium and lowers it for all $\bar{\epsilon}^{LB} > \bar{\epsilon}$ in the liquidity trap equilibrium. When the*

⁵Appendix C.B contains a derivation of this expression.

demand shock is normally distributed, these results hold globally, i.e.,

$$\frac{dP^{LB}}{ds} = \begin{cases} > 0 & \text{for } \bar{\epsilon}^{LB} < \bar{\epsilon}^* \quad (\text{target equilibria}). \\ < 0 & \text{for } \bar{\epsilon}^{LB} > \bar{\epsilon}^* \quad (\text{liquidity trap equilibria}). \end{cases}$$

Appendix C.C contains a proof.

Intuitively, in the target equilibrium, the probability of a binding lower bound increases in the tails with uncertainty since the likelihood of large movements increases. In the liquidity trap equilibrium, the prediction is reversed, since, with higher uncertainty, the probability of a sufficiently positive shock increases in the tails that lifts the interest rate off the lower bound. When the demand shock follows a normal distribution, any increase in volatility raises the probability of a binding lower bound in the target equilibrium, i.e., when the probability of a binding lower bound falls below the cutoff $(\psi - 1)/\psi$; in the liquidity trap equilibrium, the probability falls. Therefore, imposing a normal distribution implies that the sign of the derivative of the probability of being at the lower bound to a mean-preserving spread depends only on which equilibrium the economy is in.⁶

Together these results indicate that changes in uncertainty should have an effect on inflation expectations only in close proximity to the lower bound. Looking at the stylized fact about changing correlations of uncertainty shocks and inflation expectations through this lens, the model is consistent with close to zero correlations in the early 2000s when the lower bound had a low probability of binding. Later in the sample, when the lower bound became a more salient concern, the correlation between uncertainty shocks and inflation expectations turned negative. The following comparative statics show within the model that a decline in the natural real rate of interest is the most likely source of these developments.

III.C Changes in the effects of uncertainty on expectations

This section explores the effects of a change in the natural real rate of interest on the relationship between uncertainty and inflation expectations. We find that, without the lower bound, our model cannot generate the observed change in the effects of increased uncertainty on expected inflation and interest rates. With a

⁶A similar set of results holds under more stringent assumptions. When the distribution is unimodal and symmetric and $\psi = 2$, the results hold globally. When $\psi \in (1, 2)$, we can show that the lower bound probability rises with uncertainty for all target equilibria and, when $\psi > 2$, the lower bound probability declines for higher uncertainty across all liquidity trap equilibria. Appendix C.E contains details.

lower bound, the equilibrium of the economy again determines the sign of the effects. In the general case, we find that the target equilibrium is, in the tails, consistent with the stylized facts of Section II. With normally distributed shocks, the results hold across all target equilibria.

We now explore the changing impact of uncertainty on expectations when the natural real rate of interest falls. Therefore, we continue to interpret the correlation between expectations and changes in uncertainty as the derivative of expected interest rate wedges with respect to uncertainty.

Proposition 2 (Changes in the natural real rate and the effects of changes in uncertainty) *Changes in uncertainty do not affect expected interest rate wedges in the absence of a lower bound on interest rates, independent of the level of the natural rate of interest r^* . In the presence of a lower bound, the following expression determines whether the sensitivity of the expected interest rate wedge to changes in uncertainty increases or falls with the natural real rate of interest*

$$\frac{d^2 \mathbb{E}i^\Delta}{dr^* ds} = - \left(\frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - p^{LB}} \right)^2 \frac{dP^{LB}}{ds}. \quad (12)$$

Appendix C.D contains the proof of these results.

This proposition allows us to sign the direction in which a fall in the natural rate pushes the sensitivity of the expected interest rate wedge to changes in uncertainty. The determining factor is whether uncertainty raises the probability of a binding lower bound. Using Lemma 3, the right-hand side of equation (12) is always negative in the target equilibrium and always positive in the liquidity trap equilibrium, provided that shocks are normally distributed. The results furthermore hold in the tails of any distribution.

As a consequence, the sensitivity of inflation expectations with respect to uncertainty declines in the target equilibrium as the natural real rate of interest falls. In the liquidity trap equilibrium, this sensitivity rises. Measuring this sensitivity through the correlation of changes in inflation expectations and uncertainty shocks therefore shows that the results in Section II are consistent with the target equilibrium and at odds with the liquidity trap equilibrium. We investigate this connection more formally in Section IV.

Intuitively, when the natural real rate falls, the policy rate is closer to the lower bound in the target equilibrium, making it more likely to bind. When shocks display higher uncertainty, the probability of a binding lower bound increases, constraining the ability to offset negative demand shocks. As a consequence, the mean inflation rate declines.

We now turn to the question of whether the monetary policy framework has an impact on the transmission of uncertainty to the real sector. Here, we allow the central bank to permanently commit to any policy

parameter θ . The central bank is given the ability to permanently adjust the level of its interest rate rule through the choice of the intercept parameter θ . Adjusting this parameter amounts to static average-inflation targeting, as laid out in Mertens and Williams (2019, 2020) and Diwan et al. (2020).

Through lowering the interest rate intercept, the central bank provides more stimulus whenever it is not constrained by the lower bound. Consequently, it runs inflation above its target level during times when it is unconstrained to correct for the downward bias in inflation expectations due to the lower bound. That is, the central bank makes up for shortfalls during recessions by overshooting the inflation target during expansions. In that sense, it averages over the business cycle.

We are interested in the effect static average-inflation targeting has on the effects of changes in uncertainty on inflation expectations. We get the following result.

Proposition 3 (Static average-inflation targeting and the effects of changes in uncertainty)

Changes in uncertainty do not affect expected interest rate wedges in the absence of a lower bound on interest rates, independent of the level of the policy intercept θ . In the presence of a lower bound, the following expression determines whether the sensitivity of the expected interest rate wedge to changes in uncertainty increases or falls with the policy intercept θ

$$\frac{d^2 \mathbb{E}i^\Delta}{d\theta ds} = \frac{1}{\psi - 1} \left(\frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - P^{LB}} \right)^2 \frac{dP^{LB}}{ds}.$$

Appendix C.F contains the proof of these results.

Proposition 3 demonstrates that a change in the policy intercept can prevent the effects of a fall in the natural rate. A negative intercept θ lowers the expected interest rate wedge and increases inflation on average. This is because the central bank raises inflation above its target when unconstrained to make up, on average, for shortfalls during episodes of a binding lower bound. In the liquidity trap equilibrium, the sign of these effects is reversed.

As a result, the derivative of the expected interest rate wedge with respect to uncertainty rises when the central bank lowers the intercept of the policy rule. These results hold globally for normally distributed shocks and in the tails for any distribution.

In the following section, we document the empirical patterns around inflation expectations and expected interest rates and link them to movements in the natural real rate of interest.

IV Empirical Implications

We approach the empirical tests of the theoretical model in three steps. First, we recognize that the New Keynesian model predicts that the stylized fact should appear in the correlation of uncertainty shocks and changes in expectations of interest rates. After confirming this, we provide evidence that the natural real rate of interest fell during our sample period. Consistent with the theoretical predictions of our model, we show that the correlations of interest rates fell with it. And, lastly, we perform formal tests that link changes in the correlations with the natural real rate of interest.

IV.A Stylized fact for interest rates

The theoretical model of the previous section shows that the stylized fact for inflation expectations should also apply to interest rate expectations. The reason behind this prediction is the Fisher equation: As uncertainty changes inflation expectations, interest rate expectations adjust accordingly.

Figure 3 demonstrates that, as the correlation between uncertainty shocks and changes in inflation expectations turned negative, so did the correlation between uncertainty shocks and changes in expected interest rates. Here, expected interest rates are measured as the five-year ahead five-year Treasury forward rate (TFR) that we compute using constant maturity Treasury yields.

We interpret this declining correlation as the economy initially being shielded from uncertainty shocks, while over time their impact spills over increasingly to the real sector. With monetary policy serving as the primary tool to guard the economy against adverse shocks, the negative correlation between uncertainty shocks and changes in interest rates and inflation expectations, respectively, suggest that the lower bound affected the ability to respond to negative shocks using the policy rate over the sample period.

IV.B The Decline in the Natural Real Rate of Interest

In recent years, a body of research has documented a decline in the natural real rate of interest in the United States and other advanced economies over the past few decades (Williams, 2017).⁷ Figure 4 plots the natural interest rate from 1999 to 2020, using daily estimates of the real natural rate from Christensen and Rudebusch (2019) and daily five-year forward zero-coupon TIPS yields five years ahead from Gürkaynak et al. (2010) as a proxy for the natural rate. Over the sample period, estimates of the natural real rate of interest from

⁷See, for example, Laubach and Williams (2003), Kiley (2015), Lubik and Matthes (2015), Johannsen and Mertens (2016), Holston et al. (2017), Crump et al. (2017), Del Negro et al. (2017) and Christensen and Rudebusch (2019). Estimates from these models for r^* show a decline consistent with that from the measure used in this paper.

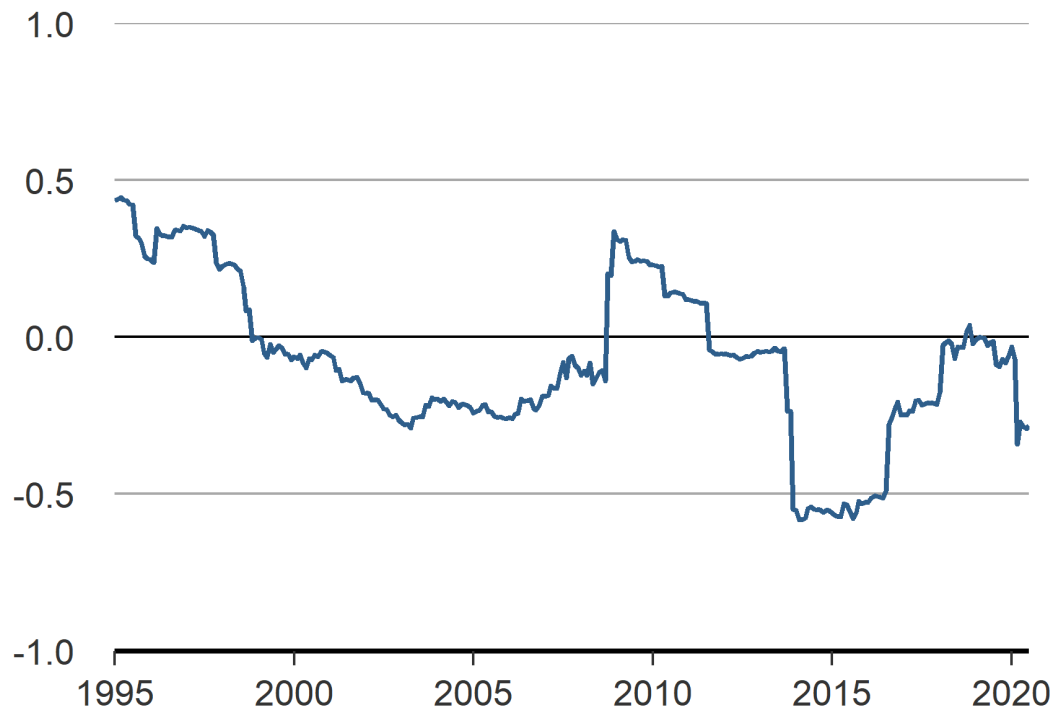


Figure 3: Five-year trailing-window correlation of monthly changes in the VIX with monthly changes in the five-year ahead five-year forward Treasury constant maturity rate.

Christensen and Rudebusch (2019) declined by 8 basis points per year on average and five-year five-year forward zero-coupon TIPS yields started out from a higher level and declined by 18 basis points per year on average. The persistent downward trend over time resulted in historically low levels of the natural rate in recent years.

The low natural rates of interest are likely to persist for an extended period of time for at least three reasons. First, measures of historical levels of r^* display a significant amount of persistence. The estimates aim to capture the low frequency component of real short-term interest rates and, as such, are highly persistent. Holston et al. (2017) show that their estimates of the natural rate are nonstationary, reflecting a high degree of persistence. Second, even well into the recovery from the Great Recession of 2007-2009, r^* had not returned to historically normal levels. Therefore, it is likely that long-term influences are holding the natural rate of interest down. Third, many possible explanations for low r^* , not only in the United States but internationally, reflect highly persistent forces affecting the global supply and demand for savings. For example, one potential explanation for the decline in r^* is a dramatic slowdown in trend real GDP growth in many advanced economies. For a more detailed discussion, see Williams (2017).

To investigate the possibility of a structural break in the level of r^* , we apply the Andrews (1993) test,

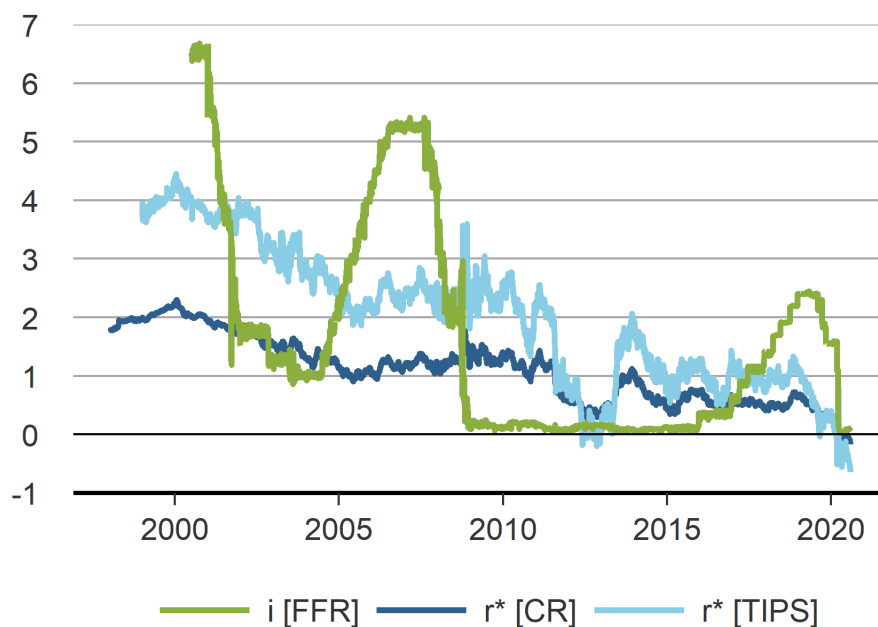


Figure 4: The green line depicts the effective federal funds rate at the daily frequency, while the dark blue line shows daily estimates of the natural real rate of interest from Christensen and Rudebusch (2019) (CR), and the light blue line shows the five-year ahead five-year forward zero-coupon TIPS rate (TIPS) computed by Gürkaynak et al. (2010) from Treasury real yields.

which shows the existence of a structural break in all measures of the natural rate.⁸ The test identifies July 2011 as the breakpoint for both the CR and TIPS measures of the natural rate. Based on this dating of the structural break in r^* , in the remainder of this subsection we split our overall sample into two subsamples: January 1999 through June 2011, and July 2011 through July 2020. We end the sample in July 2020 before the announcement of the flexible average-inflation targeting framework by the Federal Reserve. Our theoretical analysis suggests that the new framework should lead to a different data-generating process for our variables of interest.

Table 1 reports the average r^* and the average correlation of changes in expected inflation and interest rates with changes in the VIX for each subsample. To test for differences across subsamples, we compute the t -statistic for the null hypothesis that the estimates are the same across subsamples, controlling for serial correlation using the methodology outlined by Wilks (1997).⁹ In the earlier subsample, the natural real rate of interest is higher than in the later sample. This feature allows us to assess the theoretical predictions of the model. In the target equilibrium of our theoretical model, a decline in the natural real rate lowers the

⁸For the CR measure of the natural rate, the sup F value is 621.55 which occurs in July 2011, and for the TIPS measure, the sup F value is 662.94 which also occurs in July 2011. In both cases, the null hypothesis of structural stability in the level of r^* is rejected at the 1% level.

⁹Appendix D.A outlines the details of constructing the t -statistic as applied in this context.

correlation between uncertainty shocks and changes in expectations about inflation and interest rates further into negative territory, at least in the tails of the distribution. In the liquidity trap equilibrium, the predictions are reversed.

Table 1: Differences in moments across subsamples

frequency	mean r^*		mean correlation with ΔVIX			
	r_{CR}^* monthly	r_{TIPS}^* monthly	BEI daily	BEI monthly	TFR daily	TFR monthly
Jul 2011 - Jul 2020	0.54	0.83	-0.29	-0.35	-0.33	-0.23
Jan 1999 - Jun 2011	1.47	2.88	-0.09	-0.08	-0.16	-0.08
difference	-0.93***	-2.05***	-0.20***	-0.27***	-0.17***	-0.15**

Note: ***, **, and * signify that the difference in means across subsamples is statistically significant at the 1% level, 5% level, and 10% level, respectively, accounting for serial correlation. Correlations are the trailing-window correlation of the change in the variable with the change in the VIX (here referred to as ΔVIX). Correlations are computed over one-year windows at the daily frequency and over five-year windows at the monthly frequency. r_{CR}^* denotes estimates of the natural rate from Christensen and Rudebusch (2019), and r_{TIPS}^* is the five-year ahead five-year forward zero-coupon TIPS rate computed from Gürkaynak et al. (2010) used as a proxy for the natural rate. BEI is the five-year ahead five-year forward breakeven inflation rate, while TFR is the five-year ahead five-year forward Treasury constant maturity yield.

Consistent with the target equilibrium of our theoretical model, the mean daily correlation, measured over a one-year trailing-window, between changes in five-year ahead five-year forward breakeven inflation expectations (BEI) and changes in the VIX declined from -0.09 in the earlier subsample to -0.29 in the later subsample. This decline is statistically significant at the 1% level.

The analogous mean correlation, again measured at the daily frequency over a one-year trailing-window, between changes in five-year five-year forward Treasury constant maturity yields (TFR) and changes in the VIX declined from -0.16 in the earlier subsample to -0.33 in the later subsample. This difference across subsamples is significant at the 1% level.

These changes in the link between uncertainty shocks and interest rates and inflation, respectively, hold up at different frequencies. As Table 1 shows, the corresponding monthly correlation for breakeven inflation rates, measured over a five-year trailing window, fell from -0.08 to -0.35 across subsamples, a statistically significant decline at the 1% level. Similarly, the correlation for interest rates declined as well with the changes being statistically significant at the 5% level.

IV.C Lower Natural Rate and the Effects of Uncertainty Shocks on Inflation

This section tests the empirical prediction of our theoretical model that the correlation between uncertainty shocks and movements in inflation expectations changes as the natural real rate of interest falls. The target equilibrium of our model predicts a fall in the correlation while a liquidity trap equilibrium implies a movement in the opposite direction.

To investigate whether either equilibrium is consistent with the data, we run a time series regression of the correlation between changes in the VIX and changes in inflation expectations on the real natural rate of interest and an intercept.

The correlation between uncertainty shocks and changes in breakeven inflation rates serving as the dependent variable is our baseline correlation. We compute it in three different ways. First, we compute the correlation from daily data over a 14-day rolling window. Despite obtaining a noisier measure, we choose a short window for computing the correlation to avoid having a highly persistent regressor. Second, we compute it at the monthly frequency using a five-year rolling window. And third, we compute a correlation at the monthly frequency from daily data within this month, labeled as “within.” Further details can be found in Section II. Data for the natural rate of interest, r^* , are the daily estimates of Christensen and Rudebusch (2019) and daily data on the five-year ahead five-year forward zero-coupon TIPS yield computed from Gürkaynak et al. (2010); for the monthly and within specifications, we take the monthly average of the daily values of r^* . The sample period for the daily regression and the within regression is January 1999 to July 2020, while the sample for the monthly regression is January 2004 to July 2020.¹⁰

Together, all empirical specifications show a significantly positive coefficient on the natural rate of interest. Table 2 reports the estimated coefficients, accounting for serial correlation using Newey-West standard errors. The left panel – Columns (1), (2), and (3) – uses the CR measure of r^* as a regressor, while the right panel – Columns (4), (5), and (6) – uses the TIPS proxy for r^* . Column (1), the regression of the two-week trailing-window correlation between changes in the VIX and changes in inflation expectations on the CR estimates of r^* at the daily frequency, shows a highly significant positive relationship between the sensitivity of inflation expectations to uncertainty shocks and the natural rate of interest. Column (2), changing the frequency to monthly and computing the correlation over a five-year trailing window, increases the estimated coefficient on the CR r^* measure, which remains significant at the 1% level. Finally, as a robustness check, the within

¹⁰There is a shorter sample with a later start date for the monthly regression because the correlation is computed at the monthly frequency with a long enough trailing window to allow for a meaningful calculation. With daily data, a trailing window correlation can be computed with the same number of data points with fewer calendar months of data being cut off.

Table 2: Regression of correlation between changes in VIX and BEI on the natural rate of interest

	<i>Dependent variable:</i>					
	corr(Δ VIX _t , Δ BEI _t)					
	r^*_{CR}			r^*_{TIPS}		
	(1)	(2)	(3)	(4)	(5)	(6)
r^*	0.192*** (0.038)	0.313*** (0.078)	0.190*** (0.039)	0.092*** (0.018)	0.134*** (0.036)	0.091*** (0.019)
constant	-0.356*** (0.045)	-0.493*** (0.106)	-0.362*** (0.042)	-0.334*** (0.040)	-0.432*** (0.098)	-0.341*** (0.036)
Frequency	daily	monthly	within	daily	monthly	within
Sample Start	Jan 1999	Jan 2004	Jan 1999	Jan 1999	Jan 2004	Jan 1999
Sample End	Jul 2020	Jul 2020	Jul 2020	Jul 2020	Jul 2020	Jul 2020
Observations	5,374	199	259	5,371	199	259
R ²	0.072	0.483	0.120	0.078	0.463	0.132

Note: Newey-West (HAC) standard errors. * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

specification in Column (3) computes the correlation within each calendar month and regresses the resulting time series on monthly averages of the natural rate of interest. This approach avoids any sample overlap in the time series for the trailing-window correlation entering the regression. The estimated coefficient on r^* in this specification is positive, with a magnitude about the same as for the regression at the daily frequency, and is also significant at the 1% level.

Columns (4), (5), and (6), running the analogous specifications using the five-year zero-coupon TIPS rates as a proxy for r^* , yields the same conclusions. The estimated coefficients on the TIPS r^* in these regressions are slightly lower, due to the greater range in this measure over the sample period, but are still positive in magnitude and significantly different from zero at the 1% level, so they have the same interpretation as the other estimates.

These results establish a significantly positive link between the natural real rate of interest and the impact of uncertainty shocks on inflation expectations. The positive sign of the coefficient indicates that this correlation between changes in uncertainty and changes in inflation expectations tends to decline when r^* falls, as we observe broadly in the data over our sample period. This positive sign is consistent with the target equilibrium of our theoretical model and at odds with the liquidity trap equilibrium.

In the following section, we show that these results carry over to nominal interest rates.

IV.D The Effects of Uncertainty Shocks on Interest Rates

We now turn to test the empirical prediction of our model that the correlation between uncertainty shocks and changes in interest rates also moves with a decline in the natural real rate of interest.

Analogously to Table 2 above, Table 3 presents time series regression results to formally test this hypothesis. Now the correlation is between changes in the VIX and changes in interest rates as measured by five-year ahead five-year forward Treasury constant maturity yields (TFR). Again, the left panel – Columns (1), (2), and (3) – uses the CR measure of r^* as a regressor, while the right panel – Columns (4), (5), and (6) – uses the TIPS proxy for r^* .

Table 3: Regression of correlation between changes in VIX and TFR on the natural rate of interest

	<i>Dependent variable:</i>					
	corr($\Delta VIX_t, \Delta TFR_t$)					
	r^*_{CR}			r^*_{TIPS}		
	(1)	(2)	(3)	(4)	(5)	(6)
r^*	0.158*** (0.042)	0.129 (0.107)	0.167*** (0.051)	0.086*** (0.020)	0.035 (0.043)	0.090*** (0.024)
constant	-0.396*** (0.051)	-0.277 (0.194)	-0.417*** (0.058)	-0.395*** (0.045)	-0.218 (0.175)	-0.415*** (0.050)
Frequency	daily	monthly	within	daily	monthly	within
Sample Start	Jan 1998	Jan 1998	Jan 1998	Jan 1999	Jan 1999	Jan 1999
Sample End	Jul 2020	Jul 2020	Jul 2020	Jul 2020	Jul 2020	Jul 2020
Observations	5,625	271	271	5,381	259	259
R ²	0.049	0.119	0.087	0.066	0.043	0.117

Note: Newey-West (HAC) standard errors. * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

As shown in Column (1) of the table, the regression of the two-week trailing-window correlation between changes in the VIX and changes in long-term expected interest rates on the CR estimates of r^* and an intercept at the daily frequency, we estimate a positive and statistically significant relationship between the responsiveness of interest rate expectations to uncertainty shocks and the natural rate of interest. The same result holds for the regression on the TIPS proxy for r^* at the daily frequency, shown in Column (4) of the table, with an estimated coefficient on the natural rate lower in magnitude but nonetheless positive and statistically significant. Moving from daily to monthly frequency for these time series regressions, with monthly results shown in Columns (2) and (5), leads to an estimated coefficient on r^* that is still positive, but that is no longer

significant. However, computing the correlation within a month (as described above), shown in Columns (3) and (6), yields positive estimated coefficients on r^* that are significant at the 1% level.

In sum, all the empirical results in this paper are consistent with the theoretical prediction of the target equilibrium described in the theoretical section of this paper, and inconsistent with being in a liquidity trap equilibrium. These results supporting the U.S. economy being in a target equilibrium are consistent with the findings of Mertens and Williams (2021) using data from derivatives markets.

V Robustness of Empirical Results

We assess the robustness of our main empirical results along two dimensions, using alternative measures of uncertainty and of inflation expectations. Overall, we find that our qualitative results are robust to these alternative measures, although the level of statistical significance is weaker in some instances.

Our first robustness exercise examines alternative measures of inflation expectations. The breakeven inflation measure analyzed earlier is potentially contaminated by time-varying liquidity and risk premiums. To assess the importance of these premiums on our results, we consider two measures of inflation expectations that are arguably less prone to this type of mismeasurement.

The first alternative measure of inflation expectations is the five-year, five-year forward expected inflation rates, which are constructed to remove estimated risk and liquidity premiums by D'Amico et al. (2018) (DKW). The third column of Table 4 reports the subsample mean correlations between changes in the VIX and changes in this measure of inflation expectations at the daily frequency over one-year trailing windows. These results are similar to those using the breakeven inflation rates, and the difference in the mean correlations across subsamples is statistically significant.

The time series for the DKW measure is available at the daily frequency going back in time further than the BEI series. This longer time series allows us see how the correlation between changes in long-run inflation expectations and uncertainty shocks behaved in the 1990s. Figure 5 shows that this correlation using the DKW estimates of inflation expectations was positive in the late 1990s, before trending downward over the subsequent two decades.

Our second alternative measure of inflation expectations is the median ten-year CPI inflation expectation from the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters (SPF).¹¹ Unlike the financial

¹¹The SPF also collects five-year ahead forecasts on CPI inflation, but this measure is only available for a shorter sample. We use ten-year inflation expectations for consistency with the sample used throughout this paper.

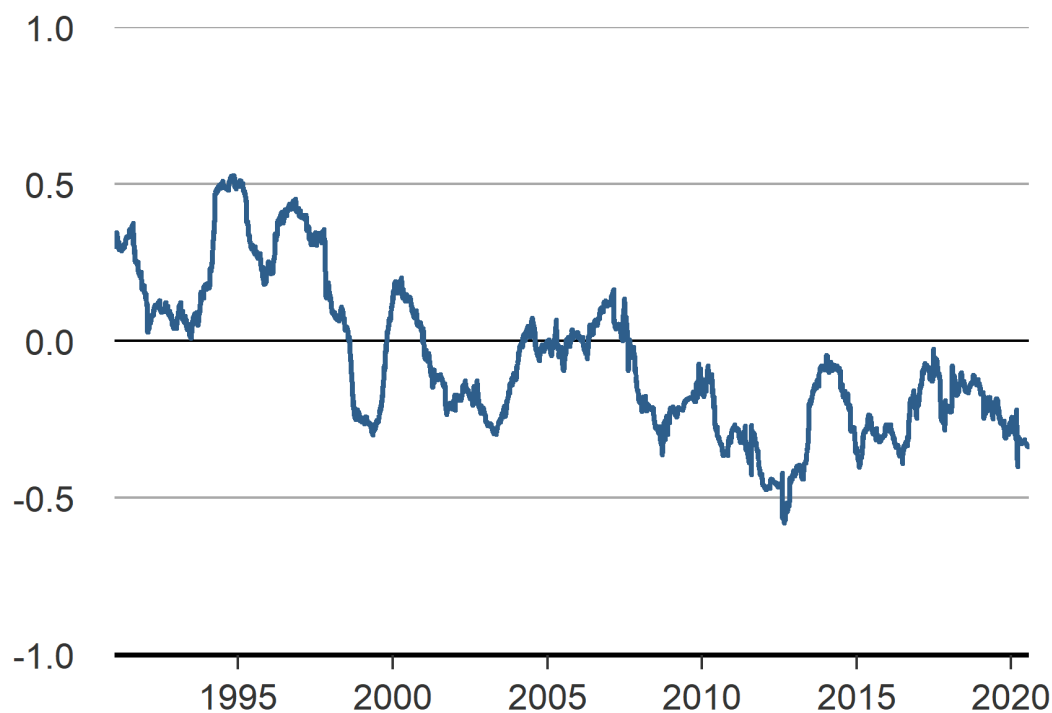


Figure 5: Correlation between changes in the VIX and changes in the five-year, five-year forward expected inflation rate, adjusted for risk premia and the liquidity premium using the estimates of D’Amico et al. (2018): daily frequency, one-year trailing-window correlation.

market data that can be analyzed at higher frequencies, the SPF data are collected quarterly. Therefore, for this measure, we average the VIX over each quarter up through the first month of the quarter to be broadly consistent with the information set available to survey participants. We compute five-year trailing-window correlations with changes in the VIX for this survey-based measure of expected inflation.

The fourth column of Table 4 shows that the results using the SPF measure of inflation expectations are broadly consistent with those using market-based measures. The difference in the mean correlations across subsamples using the SPF is both of similar size to those using market-based measures and is statistically significant.

Second, we consider two alternative measures of uncertainty: the one-year-ahead estimate of economic financial uncertainty (EFU) from Ludvigson et al. (2021) and the economic policy uncertainty (EPU) index of Baker et al. (2016). Unlike the VIX, which is priced in financial markets, the EFU index is a weighted average of conditional standard deviations of one-year-ahead forecast errors across a variety of financial indicators, while the EPU index is a news-based measure. The final two columns of Table 4 report subsample comparisons using these measures of uncertainty. These statistics are constructed as in Table 1.

Table 4: Differences in moments across subsamples.

	mean correlation with ΔVIX				mean correlation with ΔBEI	
	BEI daily	BEI monthly	DKW daily	SPF quarterly	EFU monthly	EPU monthly
frequency						
Jul 2011 - Jul 2020	-0.29	-0.35	-0.26	-0.12	-0.24	-0.10
Jan 1999 - Jun 2011	-0.09	-0.08	-0.11	0.01	-0.14	-0.06
difference	-0.20***	-0.27***	-0.15***	-0.13*	-0.10	-0.04

Note: ***, **, and * signify that the difference in means across subsamples is statistically significant at the 1% level, 5% level, and 10% level, respectively, accounting for serial correlation. Correlation is the trailing-window correlation of the change in the variable with the change in the breakeven inflation rate (here denoted as ΔBEI) or the change in the VIX (here denoted as ΔVIX), as labeled. Correlations are computed over one-year windows at the daily frequency and over five-year windows at the monthly and quarterly frequencies. EFU is the one-year ahead estimate of economic financial uncertainty from Ludvigson et al. (2021), and EPU is the economic policy uncertainty index of Baker et al. (2016). BEI is the five-year ahead five-year forward breakeven inflation rate, and DKW is the five-year ahead five-year forward breakeven inflation rate from D'Amico et al. (2018) which removes risk premia and liquidity premia from the Treasury-implied rate. SPF is the median ten-year ahead CPI inflation rate from the Survey of Professional Forecasters conducted by the Federal Reserve Bank of Philadelphia.

As seen in the table, the subsample changes in mean correlations using these alternative measures of uncertainty are qualitatively consistent with those using the VIX as the measure of uncertainty (shown in the first two columns of the table). That said, the magnitudes of the declines in the mean correlations are smaller than in the case of the VIX.

These robustness exercises demonstrate that our key findings are not sensitive to the presence of liquidity or risk premiums in the breakeven measures of inflation expectations. Although the results using these alternative measures of uncertainty are not as quantitatively or statistically strong as for the market-based measure, qualitatively they are in the same direction and consistent with the model prediction for the target equilibrium.

VI Conclusion

This paper documents a marked downward trend in the correlation between changes in market-based measures of uncertainty and changes in expected inflation and expected interest rates over the past quarter century. These changes took place as estimates of the natural rate of interest declined significantly and the

incidence of monetary policy being near or at the lower bound increased. In this paper, we show that this constellation of observations is consistent with the predictions of a standard New Keynesian model with a lower bound on interest rates when the economy is in the target equilibrium. The lower bound on interest rates is crucial because, absent the lower bound, the correlations studied here would be zero. We also show that average-inflation targeting policies can mitigate the effects of uncertainty shocks on the unconditional means of inflation and output.

Throughout this paper, we focus on the theoretical predictions for unconditional moments and empirical counterparts of far-forward expectations. A future extension of this analysis is to examine business-cycle frequency comovements between uncertainty and inflation and interest rate expectations in the presence of persistent shocks to the economy.

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Appendix

A Model Derivations: The New Keynesian Model with a Lower Bound

Our log-linearized New Keynesian model is described by equations (1) and (2) with i.i.d. demand shocks ϵ_t .

We can express these two equations in terms of only inflation by solving equation (1) for the output gap as $x_t = \frac{1}{\kappa}(\pi_t - \beta\mathbb{E}_t\pi_{t+1})$ and iterating forward and taking conditional expectations to give $\mathbb{E}x_{t+1} = \frac{1}{\kappa}(\mathbb{E}_t\pi_{t+1} - \beta\mathbb{E}_t\pi_{t+2})$, using the law of iterated expectations to simplify $\mathbb{E}_t\mathbb{E}_{t+1}\pi_{t+2} = \mathbb{E}_t\pi_{t+2}$. Substituting these expressions for x_t and $\mathbb{E}_t x_{t+1}$ into equation (2) yields $\pi_t - \mathbb{E}_t\pi_{t+1} = \kappa[\epsilon_t - \alpha(i_t - \mathbb{E}_t\pi_{t+1} - r^*)] + \beta(\mathbb{E}_t\pi_{t+1} - \mathbb{E}_t\pi_{t+2})$ after rearranging. We take unconditional expectations and apply the law of total expectation on expected inflation, using the fact that with i.i.d. shocks optimal monetary policy under discretion requires that unconditional expectations of inflation at any current or future horizon are equal to the same constant value, i.e. $\mathbb{E}\pi_t = \mathbb{E}\pi_{t+j} = \mathbb{E}\pi$. The Fisher relation follows: $\mathbb{E}\pi_t = \mathbb{E}i_t - r^*$.

A.A Derivation for Cutoff of Binding Lower Bound

Start from the following equations:

$$P^{LB} = G(\bar{\epsilon}^{LB}) \quad (13)$$

$$\bar{\epsilon}^{LB} = \alpha \left(i^{LB} - \theta - r^* + \frac{\psi}{\psi - 1} (\mathbb{E}i^\Delta + \theta) \right) \quad (14)$$

$$\mathbb{E}i^\Delta = \frac{1}{\alpha} \int_{-\infty}^{\bar{\epsilon}^{LB}} (\bar{\epsilon}^{LB} - \epsilon) g(\epsilon) d\epsilon. \quad (15)$$

A.A.1 Link between expectations, cutoff, and lower bound probability

Equation (14) implies that the link between expectations in equation 6 can be written as

$$-\frac{1}{\psi - 1} (\mathbb{E}i^\Delta + \theta) = \frac{1}{\psi} \left(i^{LB} - \theta - r^* - \frac{1}{\alpha} \bar{\epsilon}^{LB} \right) = \frac{1}{\psi} \left(i^{LB} - \theta - r^* - \frac{1}{\alpha} G^{-1}(P^{LB}) \right).$$

A.A.2 Critical value for P^{LB}

We rewrite equation (14) as

$$-\frac{\psi}{\psi - 1} \mathbb{E}i^\Delta = \left(i^{LB} - r^* - \frac{1}{\alpha} \bar{\epsilon}^{LB} + \left(\frac{\psi}{\psi - 1} - 1 \right) \theta \right)$$

This equation gives us an affine relationship between the expected wedge $\mathbb{E}i^\Delta$ and the cutoff $\bar{\epsilon}^{LB}$. The slope $\frac{\psi-1}{\alpha\psi}$ is positive.

Equation (15) is zero for $\bar{\epsilon}^{LB} \rightarrow -\infty$. It is monotonic since the integrand is nonnegative. Furthermore, it is concave as we shall see below. And, it tends to infinity for $\bar{\epsilon}^{LB} \rightarrow \infty$. As a result, it can intersect with the previous line twice, once, or not at all. Now note that

$$\frac{d}{d\bar{\epsilon}^{LB}} \mathbb{E}i^\Delta = \frac{d}{d\bar{\epsilon}^{LB}} \frac{1}{\alpha} \int_{-\infty}^{\bar{\epsilon}^{LB}} (\bar{\epsilon}^{LB} - \epsilon)g(\epsilon)d\epsilon = \frac{1}{\alpha} \int_{-\infty}^{\bar{\epsilon}^{LB}} g(\epsilon)d\epsilon = \frac{1}{\alpha} P^{LB}$$

by Leibniz's rule. This derivative is always positive and increasing in $\bar{\epsilon}$. The two lines thus have the same slope when

$$\bar{P}^{LB} = \frac{\psi - 1}{\psi}.$$

The two intersections between the two lines must thus be to the left and right of \bar{P}^{LB} . In the case with one intersection, it coincides with the critical lower bound probability.

A.A.3 Expression for expected interest rate wedge through integration by parts

We start from equation (15) and compute the integral over the first part of the integrand

$$\mathbb{E}i^\Delta = \frac{1}{\alpha} \left(\bar{\epsilon}^{LB} P^{LB} - \int_{-\infty}^{\bar{\epsilon}^{LB}} \epsilon g(\epsilon) d\epsilon \right). \quad (16)$$

For the second part, use the formula for integration by parts to simplify this expression further:

$$\int_{-\infty}^{\bar{\epsilon}^{LB}} \epsilon g(\epsilon) d\epsilon = [\epsilon G(\epsilon)]_{-\infty}^{\bar{\epsilon}^{LB}} - \int_{-\infty}^{\bar{\epsilon}^{LB}} G(\epsilon) d\epsilon = [\epsilon G(\epsilon)]_{-\infty}^{\bar{\epsilon}^{LB}} - [\mathcal{G}(\epsilon)]_{-\infty}^{\bar{\epsilon}^{LB}} = \bar{\epsilon}^{LB} G(\bar{\epsilon}^{LB}) - \mathcal{G}(\bar{\epsilon}^{LB}) \quad (17)$$

$G(\cdot)$ denotes the cumulative distribution function and the super-cumulative distribution function, i.e. the primitive of the c.d.f., $\mathcal{G}(\cdot)$. Plugging this into the previous expression results in

$$\mathbb{E}i^\Delta = \frac{1}{\alpha} \mathcal{G}(\bar{\epsilon}^{LB}). \quad (18)$$

B Data Appendix

	variable description	frequency	sample	source
<i>Measures of uncertainty</i>				
VIX	CBOE Volatility Index	daily	1990-	Bloomberg: VIX Index
EFU	economic financial uncertainty	monthly	1960-	Jurado et al. (2015) ¹²
EPU	news-based policy uncertainty index	daily	1985-	Baker et al. (2016) ¹³
<i>Measures of long-term inflation expectations</i>				
BEI	Treasury-implied breakeven inflation rate	daily	1999-	Bloomberg: USGG5Y5Y Index
DKW	dynamic factor model	daily	1983-	D'Amico et al. (2018) ¹⁴
SPF	Survey of Professional Forecasters	quarterly	1991-	Philadelphia Fed - SPF: CPI10 ¹⁵
<i>Measures of long-term interest rate expectations</i>				
TFR	five-year ahead five-year forward Treasury constant maturity rate	daily	1962-	Federal Reserve Board
<i>Measures of the real natural rate of interest (r^*)</i>				
CR	no-arbitrage model-implied	daily	1998-	Christensen and Rudebusch (2019) ¹¹
TIPS	five-year ahead five-year forward zero-coupon TIPS rate	daily	1999-	Gürkaynak et al. (2010) ¹¹

¹²Updated data downloaded from the authors' website: www.sydneyludvigson.com/macro-and-financial-uncertainty-indexes

¹³Updated data downloaded from the authors' website: www.policyuncertainty.com/us_monthly.html - Daily News Index

¹⁴Data from: www.federalreserve.gov/econres/notes/feds-notes/DKW_updates.csv

¹⁵Data from: www.philadelphiafed.org/surveys-and-data/data-files

C Proofs and Derivations

C.A Proof of Proposition 1

Proof: For the comparative statics with respect to the mean-preserving spread, start with the derivative with respect to s .

$$\begin{aligned}
 \frac{d\mathbb{E}i^\Delta}{ds} &= \frac{d}{ds} \frac{1}{\alpha} \mathcal{G}^\varphi(s, \bar{\epsilon}^{LB}) \\
 &= \frac{1}{\alpha} \mathcal{G}_1^\varphi(s, \bar{\epsilon}^{LB}) + \mathcal{G}_2^\varphi(s, \bar{\epsilon}^{LB}) \frac{\psi}{\psi-1} \frac{d\mathbb{E}i^\Delta}{ds} \\
 &= \frac{1}{\alpha} \frac{\mathcal{G}_1^\varphi(s, \bar{\epsilon}^{LB})}{1 - \frac{\psi}{\psi-1} \mathcal{G}_2^\varphi(s, \bar{\epsilon}^{LB})} \\
 &= \frac{1}{\alpha} \frac{\psi-1}{\psi} \frac{\mathcal{G}_1^\varphi(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - \mathcal{G}_2^\varphi(s, \bar{\epsilon}^{LB})}
 \end{aligned}$$

□

C.B Derivation of $\frac{dP^{LB}}{ds}$

Differentiating the equilibrium correspondence in equation (9) on both sides with respect to s yields

$$\mathcal{G}_1(s, \bar{\epsilon}^{LB}) + G(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} = \frac{\psi-1}{\psi} \frac{d\bar{\epsilon}^{LB}}{ds}$$

Consequently, we get

$$\frac{d\bar{\epsilon}^{LB}}{ds} = \frac{\mathcal{G}_1(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - G(s, \bar{\epsilon}^{LB})}$$

To sign the derivative dP^{LB}/ds , we plug the expression for the derivative of the cutoff into the derivative of P^{LB} to get

$$\begin{aligned}
 \frac{dP^{LB}}{ds} &= G_1(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} \\
 &= G_1(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{\mathcal{G}_1(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - G(s, \bar{\epsilon}^{LB})}.
 \end{aligned} \tag{19}$$

C.C Proof of Lemma 3

Proof: We start by differentiating the equilibrium correspondence in (9) to get

$$\frac{\psi - 1}{\psi} \frac{d\bar{\epsilon}^{LB}}{ds} = \mathcal{G}_2(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} + \mathcal{G}_1(s, \bar{\epsilon}^{LB}).$$

Rewriting this expression, we get

$$\left(\frac{\psi - 1}{\psi} - P^{LB} \right) \frac{d\bar{\epsilon}^{LB}}{ds} = \mathcal{G}_1(s, \bar{\epsilon}^{LB}) > 0.$$

Consequently, $(\frac{\psi-1}{\psi} - P^{LB})$ and $\frac{d\bar{\epsilon}^{LB}}{ds}$ are of the same sign. The cutoff is thus increasing in the target equilibrium and decreasing in the liquidity trap equilibrium. \square

Intuitively, by Lemma 2, the function $\mathcal{G}(\cdot)$ shifts up when uncertainty increases. We furthermore know that $\lim_{\epsilon \rightarrow -\infty} \mathcal{G}(\epsilon) = 0$ since \mathcal{G} is obtained by integrating over the cumulative density function. Furthermore, for the same reason, the slope of \mathcal{G} converges to 1 for large values

$$\lim_{\epsilon \rightarrow \infty} \mathcal{G}'(\epsilon) = 1.$$

Lastly, since $\mathcal{G}''(\epsilon) = g(\epsilon)$ is the probability density function and thus non-negative, the function $\mathcal{G}(\cdot)$ is convex. At the same time, the right-hand side of the equilibrium correspondence in (9) is independent of uncertainty.

As Figure 2 illustrates, the \mathcal{G} -function intersects the right-hand side of the equilibrium correspondence from above to below in the target equilibrium and from below to above in the liquidity trap equilibrium.

As a result, increasing uncertainty will weakly shift the intersection to the right in the target equilibrium and weakly to the left in the liquidity trap equilibrium. That is, the equilibrium cutoff $\bar{\epsilon}^{LB}$ shifts to the right in the target equilibrium, thereby weakly increasing the part of the state space in which the lower bound binds.

In the liquidity trap equilibrium, the predictions are reversed: The equilibrium cutoff $\bar{\epsilon}^{LB}$ shifts to the left, thereby weakly lowering the probability of a binding lower bound.

Under normality, these results hold globally. The term that determines the sign of the change in the lower bound probability in response to higher uncertainty is

$$\frac{dP^{LB}}{ds} = G_1(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{\mathcal{G}_1(s, \bar{\epsilon}^{LB})}{\frac{\psi-1}{\psi} - G(s, \bar{\epsilon}^{LB})}.$$

Assuming a normal distribution of the demand shock, i.e. $\varepsilon \sim N(0, \sigma^2)$, we write this term as

$$\frac{dP^{LB}}{ds} = \frac{1}{2\pi\sigma} \left(\frac{e^{-\frac{(\bar{\varepsilon}^{LB})^2}{\sigma^2}}}{\frac{\psi-1}{\psi} - \Phi\left(\frac{\bar{\varepsilon}^{LB}}{\sigma}\right)} - \sqrt{2\pi} \frac{\bar{\varepsilon}^{LB} e^{-\frac{(\bar{\varepsilon}^{LB})^2}{2\sigma^2}}}{\sigma} \right),$$

where $\Phi(\cdot)$ is the c.d.f. of the standard normal distribution. Substituting $\bar{\varepsilon}^{LB} = -\sqrt{2}\sigma\varepsilon$ and recognizing that the term outside the bracket is always positive, we analyze the term in the bracket

$$K(\varepsilon) = \frac{1}{2} \frac{1}{\frac{\psi-1}{\psi} - \Phi(-\sqrt{2}\varepsilon)} + e^{\varepsilon^2} \sqrt{\pi} \varepsilon$$

that is independent of the volatility parameter σ .

We now show that the sign of $K(\varepsilon)$ is negative for $\varepsilon < \varepsilon^*$ where $\varepsilon^* = \Phi^{-1}\left(\frac{\psi-1}{\psi}\right)$ is the point of the singularity and positive for $\varepsilon > \varepsilon^*$. To this end, we proceed in the following steps.

First, when ε approaches negative infinity, $K(\varepsilon) < 0$ approaches negative infinity, i.e. $\lim_{\varepsilon \rightarrow -\infty} K(\varepsilon) = -\infty$. Both terms in the expression are negative with one of them being bounded. On the other end, $K(\varepsilon)$ approaches infinity for large values of ε , i.e. $\lim_{\varepsilon \rightarrow \infty} K(\varepsilon) = \infty$.

Next, approaching the singularity from the left leads the function value to approach negative infinity, $\lim_{\varepsilon \rightarrow -\bar{\varepsilon}^*} K(\varepsilon) = -\infty$. Approaching the singularity from the right leads to positive unbounded function values $\lim_{\varepsilon \rightarrow +\bar{\varepsilon}^*} K(\varepsilon) = \infty$.

Third, the function has only one singularity where the denominator of the first term is zero. It is analytic otherwise. Therefore, the function starts from unbounded negative values, reaches a maximum, and then converges to negative infinity when approaching the singularity. On the right-hand side of the singularity, the function starts from a high value, falls to a minimum, and then converges to positive infinity. Next we compute the two extrema and show that the maximum to the left of the singularity sits below zero and the minimum to the right above zero.

To compute the extrema, we take first-order conditions and set them to zero, i.e., $K'(\varepsilon) = 0$, which results in

$$2\pi e^{2\varepsilon^2} (1 + 2\varepsilon^2) - \frac{1}{\left(\frac{\psi-1}{\psi} - \Phi(-\sqrt{2}\varepsilon)\right)^2} = 0.$$

Rearranging the equation and taking the square root results in two conditions that hold at the extrema

$$\frac{1}{\frac{\psi-1}{\psi} - \Phi(-\sqrt{2}\varepsilon_1)} = -e^{\varepsilon_1^2} \sqrt{2\pi(1 + 2\varepsilon_1^2)}$$

and

$$\frac{1}{\frac{\psi-1}{\psi} - \Phi(-\sqrt{2}\varepsilon_2)} = +e^{\varepsilon_2^2} \sqrt{2\pi(1 + 2\varepsilon_2^2)}.$$

Since the left-hand side is negative in the first and positive in the second equation, the two extrema are on opposite sides of the singularity at ε^* .

Evaluating the function value shows that the function value at the maximum to the left of the singularity is below zero. Therefore, replace the first expression in $K(\cdot)$ with the expression from the first-order condition to arrive at

$$K(\varepsilon_1) = \sqrt{\pi} e^{\varepsilon_1^2} \left(\varepsilon_1 - \sqrt{\frac{1}{2} + \varepsilon_1^2} \right) < 0.$$

As a result, $K(\varepsilon_1) < 0$ for all $\varepsilon < \varepsilon^*$. For $\varepsilon > \varepsilon^*$, we get

$$K(\varepsilon_2) = \sqrt{\pi} e^{\varepsilon_2^2} \left(\varepsilon_2 + \sqrt{\frac{1}{2} + \varepsilon_2^2} \right) > 0.$$

Since the maximum to the left of the singularity lies below zero, all function values for $\varepsilon < \varepsilon^*$ are negative. And since the minimum to the right is above zero, all function values for $\varepsilon > \varepsilon^*$ are positive.

C.D Proof of Proposition 2

Proof: We start from the derivative of the expected wedge with respect to the natural rate

$$\frac{d\mathbb{E}i^\Delta}{dr^*} = \frac{1}{\alpha} \mathcal{G}_{2(s, \bar{\varepsilon}^{LB})} \frac{d\bar{\varepsilon}^{LB}}{dr^*}. \quad (20)$$

Now, computing the second mixed derivative as

$$\frac{d^2\mathbb{E}i^\Delta}{dr^* ds} = \frac{1}{\alpha} \mathcal{G}_{1,2(s, \bar{\varepsilon}^{LB})} \frac{d\bar{\varepsilon}^{LB}}{dr^*} + \frac{1}{\alpha} \mathcal{G}_{2,2(s, \bar{\varepsilon}^{LB})} \frac{d\bar{\varepsilon}^{LB}}{dr^*} \frac{d\bar{\varepsilon}^{LB}}{ds} + \frac{1}{\alpha} \mathcal{G}_{2(s, \bar{\varepsilon}^{LB})} \frac{d^2\bar{\varepsilon}^{LB}}{dr^* ds},$$

where we get the last term by computing

$$\frac{d\bar{\epsilon}^{LB}}{dr^*} = -\alpha \left(1 - \frac{\psi}{\psi-1} \frac{d\mathbb{E}i^\Delta}{dr^*} \right) \quad (21)$$

and

$$\frac{d^2\bar{\epsilon}^{LB}}{dr^*ds} = \alpha \frac{\psi}{\psi-1} \frac{d^2\mathbb{E}i^\Delta}{dr^*ds}.$$

Plugging these expressions into the second mixed derivative yields

$$\frac{d^2\mathbb{E}i^\Delta}{dr^*ds} = \frac{1}{\alpha} \frac{\psi-1}{\psi} \frac{1}{\frac{\psi-1}{\psi} - P^{LB}} \left(\mathcal{G}_{1,2}(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} \right) \frac{d\bar{\epsilon}^{LB}}{dr^*}. \quad (22)$$

We combine equations (20) and (21) to get

$$\frac{\psi-1}{\psi} \left(1 + \frac{1}{\alpha} \frac{d\bar{\epsilon}^{LB}}{dr^*} \right) = \frac{1}{\alpha} \mathcal{G}_2(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{dr^*},$$

or, rewritten,

$$\frac{d\bar{\epsilon}^{LB}}{dr^*} = -\alpha \frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - P^{LB}}. \quad (23)$$

Taken together and using equation (19) results in

$$\begin{aligned} \frac{d^2\mathbb{E}i^\Delta}{dr^*ds} &= - \left(\frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - P^{LB}} \right)^2 \left(\mathcal{G}_{1,2}(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} \right) \\ &= - \left(\frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - P^{LB}} \right)^2 \frac{dP^{LB}}{ds}. \end{aligned}$$

□

C.E Details on Symmetric and Unimodal Distributions

For symmetric and unimodal distributions, we get following result.

Lemma 4 (Global result for symmetric unimodal distributions)

In the case of $\psi = 2$ and a unimodal symmetric distribution, a lower natural real rate of interest implies a more negative impact of increased uncertainty on inflation expectations across all target equilibria. Across all liquidity trap equilibria

in this special case, a lower natural rate renders the impact of higher uncertainty on inflation expectations less negative. The result also holds for all target equilibria when $\psi \in (1, 2)$ and all liquidity trap equilibria for $\psi > 2$.

Proof: We start by showing that the result of Proposition 2 holds globally for unimodal symmetric distributions when $\psi = 2$. Therefore, start with the observation that, in this case,

$$p^{LB} = \frac{\psi - 1}{\psi} = \frac{1}{2}.$$

Due to the symmetry of the distribution, this case corresponds to a cutoff $\bar{\epsilon}^{LB} = 0$. An increase in uncertainty for a symmetric unimodal distribution exhibits the single-crossing property for the distributions. As a result, the c.d.f. of the riskier distribution lies above the less risky distribution for all negative values and below for all positive values.

Due to single-crossing of the distributions, the probability of a binding lower bound increases for all negative values and decreases for all positive values with higher uncertainty. Formally, $\mathcal{G}_{1,2}(s, \epsilon) > 0$ for all $\epsilon < 0$ and $\mathcal{G}_{1,2}(s, \epsilon) < 0$ for all $\epsilon > 0$. Consequently the steps in the proof of Proposition 2 hold globally.

By the same logic, when $\psi \in (1, 2)$, the critical value for the lower bound $\frac{\psi-1}{\psi} < \frac{1}{2}$. As a result, all the target equilibria have a lower bound probability of less than 50% and, with a symmetric distribution, the corresponding critical cutoff is in the negative territory. With single crossing, $\mathcal{G}_{1,2}(s, \epsilon) > 0$ for all target equilibria. When $\psi > 2$, the critical cutoff is in positive territory and the results for the liquidity trap equilibria hold globally. \square

C.F Proof of Proposition 3

Proof: We follow analogous steps as in Section C.D. We start from the derivative of the expected wedge with respect to the intercept θ in the policy rule

$$\frac{d\mathbb{E}i^\Delta}{d\theta} = \frac{1}{\alpha} \mathcal{G}_2(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{d\theta}. \quad (24)$$

Now, computing the second mixed derivative as

$$\frac{d^2\mathbb{E}i^\Delta}{d\theta ds} = \frac{1}{\alpha} \mathcal{G}_{1,2}(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{d\theta} + \frac{1}{\alpha} \mathcal{G}_{2,2}(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{d\theta} \frac{d\bar{\epsilon}^{LB}}{ds} + \frac{1}{\alpha} \mathcal{G}_2(s, \bar{\epsilon}^{LB}) \frac{d^2\bar{\epsilon}^{LB}}{d\theta ds},$$

where we get the last term by computing

$$\frac{d\bar{\epsilon}^{LB}}{d\theta} = \alpha \left(\frac{1}{\psi - 1} + \frac{\psi}{\psi - 1} \frac{d\mathbb{E}i^\Delta}{d\theta} \right) \quad (25)$$

and

$$\frac{d^2\bar{\epsilon}^{LB}}{d\theta ds} = \alpha \frac{\psi}{\psi - 1} \frac{d^2\mathbb{E}i^\Delta}{d\theta ds}.$$

Plugging these expressions into the second mixed derivative yields

$$\frac{d^2\mathbb{E}i^\Delta}{d\theta ds} = \frac{1}{\alpha} \frac{\psi - 1}{\psi} \frac{1}{\frac{\psi-1}{\psi} - p_{LB}} \left(\mathcal{G}_{1,2}(s, \bar{\epsilon}^{LB}) + g(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{ds} \right) \frac{d\bar{\epsilon}^{LB}}{d\theta}.$$

We combine equations (24) and (25) to get

$$\frac{d\bar{\epsilon}^{LB}}{d\theta} = \frac{\alpha}{\psi - 1} + \frac{\psi}{\psi - 1} G(s, \bar{\epsilon}^{LB}) \frac{d\bar{\epsilon}^{LB}}{d\theta},$$

or, rewritten,

$$\frac{d\bar{\epsilon}^{LB}}{d\theta} = \frac{\alpha}{\psi} \frac{1}{\frac{\psi-1}{\psi} - p_{LB}}.$$

Taken together and using equation (19) results in

$$\frac{d^2\mathbb{E}i^\Delta}{d\theta ds} = \frac{1}{\psi - 1} \left(\frac{\frac{\psi-1}{\psi}}{\frac{\psi-1}{\psi} - p_{LB}} \right)^2 \frac{dP^{LB}}{ds}.$$

□

D Technical Details on Time Series Analysis

D.A Controlling for Serial Correlation in Hypothesis Testing on Sample Means

Above we compute sample means of various time series in adjacent subsamples and test the null hypothesis that these estimated means are the same across subsamples. Because the time series for which we compute the mean – both the raw time series for r^* and the trailing-window correlations – are highly persistent, we control for serial correlation to conduct valid inference. To this end, we construct the following t -statistic for this null hypothesis:

$$t = \frac{\bar{y}_2 - \bar{y}_1}{\sqrt{\text{Var}(\bar{y}_2 - \bar{y}_1)}}$$

which is the ratio of the difference in subsample means to the standard error of this difference, where y_1 is the first subsample with n_1 observations and y_2 is the second subsample with n_2 observations, with \bar{y}_i denoting the sample mean of subsample i .

Following Wilks (1997), we use a scaling factor approach to control for serial correlation in our setting. Hence, we compute the variance of the difference in means as follows:

$$\text{Var}(\bar{y}_2 - \bar{y}_1) = \hat{\vartheta}_2 \frac{s_2^2}{n_2} + \hat{\vartheta}_1 \frac{s_1^2}{n_1}$$

with s_i^2 denoting the sample variance of subsample i . For each subsample i , the scaling factor ϑ_i given by $\vartheta_i = 1 + 2 \sum_{k=1}^{n_i-1} [1 - k/n_i] \rho_k$ accounts for serial correlation ρ_k at various lags k , and the modified scaling factor $\hat{\vartheta}_i = \vartheta_i \exp(2\vartheta_i/n_i)$ corrects for potential bias in estimating this scaling factor.