FOMC Responses to Credit Spreads

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July 2025

Abstract

Achieving the divine coincidence requires setting the *private* borrowing rate to its value under price flexibility, so the government borrowing rate should fall one-for-one with the private-government spread. Even when Phillips-curve shocks place the divine coincidence out of reach, tracking the natural rate on private borrowing reduces Keynesian fluctuations by offsetting IS-curve shocks. This paper uses an SVAR with the Gilchrist and Zakrajšek (2012) excess bond premium to measure the potential effects of such policy. We estimate that FOMC responses to credit spreads could reduce consumption fluctuations by 20 percent, inflation fluctuations by 10 percent, and unemployment fluctuations by 30 percent.

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We are grateful to participants at an early presentation of this paper at Erasmus University Rotterdam for their constructive comments. The views expressed are those of the authors. They do not necessarily represent the views of the Federal Reserve Bank of Chicago, the Federal Reserve System, or its Board of Governors.

JEL Codes: E44, E52, E58.
Keywords: Monetary Policy, Dual Mandate, Divine Coincidence, Natural Rate of Interest, Vector Autoregression

1 Introduction

This paper evaluates the consequences for the FOMC – as measured by deviations from its dual mandate – of incorporating credit spreads into its policy rule.

Shocks to demand for government bonds – as in Fisher (2015) – create a stochastic spread between the private borrowing rate and the monetary policy rate and move the NK IS curve's intercept. These shocks have emerged as one of the two main sources of business cycles in empirical DSGE economies (Smets and Wouters, 2007; Fisher, 2015; Campbell, Ferroni, Fisher, and Melosi, 2023). In those economies, achieving the divine coincidence outcome of Blanchard and Galí (2007) requires setting the rate on *private* borrowing to its value in an analogous counterfactual economy with flexible prices. We call this the *private natural rate*. Monetary policy can do this by reducing the government borrowing rate one-to-one to the private-government credit spread. This kills fluctuations originating with these shocks by inducing the private borrowing rate to track its natural value. The FOMC can respond to supply-side disturbances in the Phillips curve by adjusting policy around this baseline (Giannoni and Woodford, 2004).

The Lucas (1976) critique cautions us from evaluating a monetary policy change without a fully specified model economy. Nevertheless, the economics of the divine coincidence allow us to calculate counterfactual outcomes using only a partially-identified structural VAR to measure shocks to demand for government bonds. Tracking the private natural rate removes the effects of such IS curve shocks on a new Keynesian economy's outcomes while leaving fluctuations from other sources unchanged, *regardless of the economy's production and price-adjustment technologies.* Therefore, we may estimate the effects of tracking the private natural rate on inflation and unemployment gaps by turning off these identified bond-demand shocks. We get the counterfactual monetary policy rate by replacing the estimated effects of the shocks with the one-to-one response to the credit spread required for the divine coincidence. This quantifies changes to the policy rate's variance and the frequency and severity of its spells at the zero lower bound.

Our baseline SVAR consists of four monthly variables, Gilchrist and Zakrajšek (2012) excess bond premium (EBP), Wu and Xia's (2016) monetary policy rate, the consumption gap (constructed from the Congressional Budget Office's measure of potential output), and core PCE inflation. Because Gilchrist and Zakrajšek removed the effects of time-varying default risks from private bond rates, the resulting spread changes should embody changes in demand for liquidity rather than changes in borrower quality. The SVAR's only required structural parameter equals the elasticity of

government-bond demand with respect to current consumption. Empirically, identified spread shocks account for about 20 percent of the consumption gap's variance and about 10 percent of inflation's variance. In our sample, most of these variance reductions occurred in the current millennium. These percentages depend little on either the exact SVAR specification or the elasticity of government-bond demand.

In-sample, the one-to-one response to private-government spreads after maturity correction only slightly affects the magnitude of the fluctuations in the monetary policy rate (with its standard deviation decreasing by 5 basis points), but it significantly alters the timing of those responses. This is because the monetary policy rate already responds to the spread, but too late. Initially, there is no immediate response of the interest rate to shocks in credit spreads. However, the policy rate responds aggressively as the decline in other macroeconomic aggregates unfolds. The counterfactual version of Wu and Xia's (2016) monetary policy rate also spends approximately the same time below zero, so tracking the private natural rate also does not require expanding the use of unconventional policy tools. We conclude that tracking the private natural rate does not require the FOMC to explain additional policy changes to the public.

Gilchrist and Zakrajšek (2012) interpret credit spread increases as mostly coming from changes in financial institutions' lending capacity.[1] Earlier work by Curdia and Woodford (2010) examined the potential of targeting credit spreads within a full-scale New Keynesian framework. In their framework increase in the credit spreads can be attributed to both default risk and or an increase in the resource cost of loan origination. They have built a model with heterogeneous savers and borrowers, and segmentation of participation in different financial markets. They have found that contemporaneous response to credit spreads can reduce the distortions caused by a financial disturbances. More recently, Caballero et al. (2024) have proposed a broader approach involving the FOMC targeting a financial conditions index (FCI) with the objective to ensure the soundness of the financial system. In their framework, promise to keep financial condition at the target triggers an endogeneous return volatility-reducing feedback loop that stabilizes the output gap.

Our approach takes a more monetary view – we primarily aim to satisfy liquidity demand of the private sector, rather than aim to improve broad financial conditions. Under this view, the FOMC can improve private-sector liquidity access by offsetting changes in the demand for the government bonds by tracking credit spreads in the stochastic intercept of an interest rate rule. We have taken a first step towards quan-

[1] They state "The confluence of our results is thus consistent with the notion that an increase in the excess bond premium reflects a reduction in the effective risk-bearing capacity of the financial sector and, as a result, a contraction in the supply of credit; ... ." (page 1693)

tifying the financial health story contributions to credit spread fluctuations by placing equity's share of financial firms' assets in a quarterly version of our SVAR's first position. Although the measured shocks to this share account for about a quarter of total fluctuations in the consumption gap, they have no discernible impact on the excess bond premium. Furthermore, the identified credit demand shocks' estimated effects remain unchanged.

The remainder of this paper proceeds as follows. The next section presents and analyzes an extension of Fisher's (2015) IS curve with bonds in the utility function, and Section 3 shows how monetary policy responses to credit spreads can offset shocks to the demand for government bonds. Section 4 presents our SVAR estimates of the costs and benefits of the policy option under consideration. This includes a sensitivity analysis using several alternative measures of the credit spread. Section 5 examines the results in the context of other financial shocks, and Section 6 offers concluding remarks about practical policy implementation issues.

2 Credit Spreads in the Natural Rate of Interest

Because Gilchrist and Zakrajšek (2012) removed time-varying default premiums from their corporate bond prices, the payoffs from the private and government bonds in their spread resemble each other. Nevertheless, their spread's standard deviation from 1973 through 2019 equals 55 basis points. Because government bonds can be traded for cash more easily than can corporate bonds, we assert that this variance reflects a time-varying price of liquidity. Here, we develop this assertion for New Keynesian economies by adding inelastically-supplied government bonds to the representative household's utility function, as in Fisher (2015).

Consider a representative household trading in markets for nominal government bonds and nominal private bonds. The fiscal authority issues government bonds, whereas private private firms issue private bonds. Both bonds are in zero net supply, and their associated nominal interest rates equal i_t^G and i_t^P. Government bonds provide a utility service. Specifically, the representative household's utility flow depends on consumption C_t and the current real value of government bonds:

$$U(C_t, B_{t+1}^G, \text{other goods}) = \frac{1}{1-\sigma}\left(\left(C_t^{1-\iota} + (1-\iota)e^{\zeta_t}\frac{B_{t+1}^G}{e^{i_t^G}P_t}\right)^{\frac{1-\sigma}{1-\iota}} - 1\right) + \text{other utility}.$$

Here, σ equals the inverse elasticity of intertemporal substitution, and $\iota \in [0, \sigma]$ equals the consumption elasticity of bond demand. The nominal par value of one-period gov-

ernment bonds owned by the household equals B^G_{t+1}, and the current price level equals P_t. This specification nests several well-known preferences. When $\iota = \sigma$, utility is additively separable in consumption and government bonds, as in Fisher (2015). When $\iota = 0$, consumption does not impact the household's demand for government bonds.[2] The random variable ζ_t changes the government bonds' utility services over time. The terms in "other goods" and "other utility" can represent the benefits from leisure and home production.

The representative household chooses nominal government bonds B^G_{t+1}, nominal private bonds B^P_{t+1}, and consumption C_t to maximize its utility subject to the budget constraint

$$P_t C_t + \frac{B^G_{t+1}}{e^{i^G_t}} + \frac{B^P_{t+1}}{e^{i^P_t}} + \text{other expenses} = B^G_t + B^P_t + \text{other income}.$$

The term "other income" may include labor income, transfers, or dividend payments, while "other expenses" could capture taxes or consumption of other goods. The first-order optimality conditions for bond purchases are

$$\frac{1 - e^{\zeta_t} C_t^{\iota}}{e^{i^G_t}} = \beta \mathbb{E}\left[\left(\frac{C_t}{C_{t+1}}\right)^{\sigma} \frac{P_t}{P_{t+1}}\right], \tag{1}$$

$$\frac{1}{e^{i^P_t}} = \beta \mathbb{E}\left[\left(\frac{C_t}{C_{t+1}}\right)^{\sigma} \frac{P_t}{P_{t+1}}\right]. \tag{2}$$

We call these the Intertemporal Substitution with Government Bond (IS-G) and Intertemporal Substitution with Private Bond (IS-P) curves. Their right-hand sides both equal the marginal rate of intertemporal substitution. Together, they require the spread between the interest rates on government and private bonds to equal

$$i^g_t - i^p_t = \ln\left(1 - e^{\zeta_t} C_t^{\iota}\right). \tag{3}$$

The bond spread combines an exogenous term (ζ_t) with an endogenous component (C_t^{ι}). When $\iota = 0$, the spread evolves exogenously. If instead $\iota > 0$, the credit spread increases with consumption. Henceforth, we refer to ζ_t as the *credit spread shock*.

To study the joint behavior of monetary policy and deviations from efficient outcomes, define variables' potential values as those that would arise under flexible prices. We denote these with tildes, e.g. $\tilde{C}_t$ for potential consumption, and we denote the variables' deviations from their potential values (their *gaps*) with hats, e.g. $\hat{C}_t$. Any variable equals the product of its potential value with its gap. Substituting these definitions

[2] This specification is isomorphic to the zero-income-effect on labor-supply preferences of Greenwood, Hercowitz, and Huffman (1988)

in the IS-G curve yields

$$\frac{1 - e^{\zeta_t}\tilde{C}_t^{\iota}\hat{C}_t^{\iota}}{e^{i_t^G}} = \beta\mathbb{E}\left[\left(\frac{\tilde{C}_t\hat{C}_t}{\tilde{C}_{t+1}\hat{C}_{t+1}}\right)^{\sigma}\frac{P_t}{P_{t+1}}\right].$$

The FOMC's dual mandate calls for price stability and maximum sustainable employment. The FOMC itself defines price stability as bringing inflation as close as possible to its target, $\pi_\star$ (currently 2 percent). We and many others interpret achieving "maximum sustainable employment" as equivalent to closing gaps in real activity. Accordingly, the ideal outcome (from the perspective of the dual mandate) sets $\hat{C}_t = \hat{C}_{t+1} = 1$ and $P_{t+1}/P_t = e^{\pi_\star}$. Blanchard and Galí (2007) note that a monetary authority can achieve this outcome if producers' optimal markups do not change over time, and they call this *the divine coincidence.* Here, we wish only to characterize the government bond rate that is consistent with the divine coincidence outcome. This is the *government natural rate.*

$$\tilde{i}_t^G = \ln\left(\frac{1 - e^{\zeta_t}\tilde{C}_t^{\iota}}{\beta\mathbb{E}\left[\left(\frac{\tilde{C}_t}{\tilde{C}_{t+1}}\right)^{\sigma}e^{-\pi_\star}\right].}\right)$$

We can use the IS-P curve to derive the corresponding *private natural rate* as

$$\tilde{i}_t^P = -\ln\left(\beta\mathbb{E}\left[\left(\frac{\tilde{C}_t}{\tilde{C}_{t+1}}\right)^{\sigma}e^{-\pi_\star}\right]\right).$$

When the divine coincidence is feasible, achieving it requires setting $i_t^G = \tilde{i}_t^G$ and $i_t^P = \tilde{i}_t^P$. Equation (3) guarantees that the monetary authority can do so even though it only controls i_t^G. That is, the government natural rate equals the private natural rate (which itself equals the log marginal rate of intertemporal substitution plus a steady-state inflation compensation) plus a spread arising from the utility services of government bonds.

3 Monetary Policy with Credit Spreads

In empirically-relevant New Keynesian economies, producers' changing markups put the divine coincidence out of reach.[3] Nevertheless, incorporating $\tilde{i}_t^G$ into monetary policy as a time-varying baseline can improve outcomes by removing the first-order effects of ζ_t on $\tilde{C}_t$. To demonstrate this, we first create log-linear certainty-equivalent approximations of the IS-G and IS-P curves, Equations (1) and (2) around the economy's steady state. Using c_t and c_{t+1} to denote the logarithmic deviations of C_t and C_{t+1} from their steady-state value, these are

$$c_t = \frac{\sigma}{\sigma + \iota s_\star}\mathbb{E}\left[c_{t+1}\right] - \frac{s_\star}{\sigma + \iota s_\star}\left(\zeta_t - \zeta_\star\right) - \frac{1}{\sigma + \iota s_\star}\left(i_t^G - i_\star^G - \mathbb{E}\left[\pi_{t+1} - \pi_\star\right]\right), \quad (4)$$

$$c_t = \mathbb{E}\left[c_{t+1}\right] - \frac{1}{\sigma}\left(i_t^P - i_\star^P - \mathbb{E}\left[\pi_{t+1} - \pi_\star\right]\right). \quad (5)$$

Here, $i_\star^P$ and $i_\star^G$ are the steady-state interest rates; and $s_\star \equiv e^{i_\star^P - i_\star^G} - 1$ approximately equals the steady-state credit spread in log points. Either of these equations can be obtained from the other using the log-linear approximation of Equation (3). We choose to focus henceforth on Equation (4), because it contains the interest rate under the monetary authority's control.

Both i_t^G and ζ_t appear in Equation (4) multiplied by negative coefficients. Ceteris paribus, increasing either of them reduces c_t. To appreciate how monetary policy itself can offset movements in ζ_t, replace $i_t^G - i_\star^G$ in Equation (4) with $\tilde{i}_t^G - \tilde{i}_\star^G + \hat{i}_t^G$ and use the log linearization of $\tilde{i}_t^G$'s definition to write the (4) in terms of gaps.

$$\hat{c}_t = \frac{\sigma}{\sigma + \iota s_\star}\mathbb{E}\left[\hat{c}_{t+1}\right] - \frac{1}{\sigma + \iota s_\star}\left(\hat{i}_t^G - \mathbb{E}\left[\pi_{t+1} - \pi_\star\right]\right) \quad (6)$$

Equation (6) does not *explicitly* include ζ_t. Whether or not that shock can implicitly influence inflation and activity depends on the specification of monetary policy in $\hat{i}_t^G$. We illustrate this with three examples.

Example 1. Following Taylor (1993), the standard interest rate rule in the canonical New Keynesian model sets

$$i_t^G = i_\star^G + \phi\left(\pi_t - \pi_\star\right) + \omega_t.$$

[3] Schmitt-Grohé and Uribe (2025) provide a microeconomic foundation for markup variations with sticky prices.

Here, ω_t is a disturbance to monetary policy. This rule's implied specification for $\hat{i}_t^g$ is

$$\hat{i}_t^G = i_\star^G - \tilde{i}_t^G + \phi\left(\pi_t - \pi_\star\right) + \omega_t.$$

When the government natural rate falls, either because expected potential consumption growth falls or because the value of government bonds' utility services rises, the interest rate relevant for the economy's gaps rises. Through this channel, both expected technological slowdowns and increases in credit spreads can (inefficiently) reduce both inflation and economic activity.[4]

Example 2. The second example allows ζ_t to influence i_t^G directly.

$$i_t^G = i_\star^G + \phi\left(\pi_t - \pi_\star\right) - \lambda\left(\zeta_t - \zeta_\star\right) + \omega_t$$

This produces

$$\hat{i}_t^G = \left(s_\star - \lambda\right)\left(\zeta_t - \zeta_\star\right) + s_\star \iota \tilde{c}_t + \sigma\left(\tilde{c}_t - \mathbb{E}\left[\tilde{c}_{t+1}\right]\right).$$

If $\lambda = s_\star$, then $\zeta_t - \zeta_\star$ disappears from $\hat{i}_t^G$ and therefore from Equation 6. In this case, the exogenous driver of the credit spread influences neither inflation nor activity gaps. However, shocks to potential consumption and to its growth rate can still impact these Keynesian outcomes.

Example 3. The third monetary policy example uses $\tilde{i}_t^G$ as the interest-rate rule's time-varying intercept.

$$\begin{aligned} i_t^G &= \tilde{i}_t^G + \phi\left(\pi_t - \pi_\star\right) + \omega_t \\ &= \sigma\left(\mathbb{E}\left[\tilde{c}_{t+1}\right] - \tilde{c}_t\right) - s_\star\left(\zeta_t - \zeta_\star + \iota \tilde{c}_t\right) + \phi\left(\pi_t - \pi_\star\right) + \omega_t \end{aligned} \tag{7}$$

With this, the gap-relevant monetary policy setting is

$$\hat{i}_t^G = \phi\left(\pi_t - \pi_\star\right) + \omega_t.$$

If $\omega_t = 0$ always, then this monetary policy specification is consistent with achieving the divine coincidence outcome. If we insist on *exactly* achieving the divine coin-

[4] Much theoretical work characterizes when this and similar interest-rate rules guarantee uniqueness of non-explosive equilibria (Benhabib et al., 2001; Davig and Leeper, 2007; Coibion and Gorodnichenko, 2011; Cochrane, 2011). There is a subtle relationship between equilibrium determinacy and the pass through of spread shocks into economic outcomes. We can make the responses of inflation and activity gaps to ζ_t as small as we like by choosing ϕ to be large enough. However, large values of ϕ amplify the forward guidance puzzle and thereby magnify responses to *expected* disturbances (Cochrane, 2017).

cidence, then setting ϕ large enough makes the divine coincidence the *unique and bounded* equilibrium

These examples show that the monetary authority can remove $\zeta_t - \zeta_\star$ from the gap-relevant IS curve (Equation (6)) with an appropriate interest rate rule. This has the policy rate decline one-for-one with every increase in the credit spread's exogenous component, $s_\star (\zeta_t - \zeta_\star)$. Usual specifications of the Phillips curve include only shocks that directly influence real marginal cost and producers' optimal markups, so *monetary policy that lowers the policy rate one-for-one with the credit spread's exogenous component makes ζ_t irrelevant for Keynesian fluctuations.*

4 Measurement of Costs and Benefits

In theory, the benefits from monetary policy systematically responding to credit spreads come without cost, like the proverbial five dollar bill on the sidewalk. In practice, two obvious costs present themselves. First, this policy change could require large unexpected changes in the policy rate; while the FOMC has revealed a strong preference for gradual and small adjustments to the policy rate. Second, responding to credit spreads could bring the policy rate to its zero lower bound more frequently and thereby require greater use of unconventional policy tools.

This section measures these two costs and the policy benefits of monetary policy responding to credit spreads for the United States economy using a partially-identified structural vector autoregression with a private-government credit spread, economic activity, inflation, and the monetary policy rate. The system's single structural parameter is ι, which we require to recover ζ_t (up to a scale parameter) from the credit spread and economic activity. Because ζ_t evolves exogenously in our theory, we choose ι to maximize the share of ζ_t's long-run variance accounted for by its own one-step-ahead forecast error. We measure the variance reduction from the policy change just as theory suggests, by removing those fluctuations driven by ζ_t from the data. We measure the costs similarly, by calculating the counterfactual monetary policy rate as if the FOMC responded one-for-one to ζ_t and responded to all other shocks as it actually did in the sample period. This requires the scale parameter, which embodies the maturity structure of the bonds used to calculate the credit spread. We report results for several empirically-plausible values of this parameter below.

4.1 Data

The private-government credit spread theoretically examined above compares the returns to two bonds nominal risk-free bonds. We believe that the best available empirical counterpart to this is Gilchrist and Zakrajšek's (2012) Excess Bond Premium (hereafter, the EBP). To construct this, Gilchrist and Zakrajšek matched month-end secondary market corporate bond prices with prices of government securities with the same payoffs synthesized from the zero-coupon Treasury prices of Gürkaynak, Sack, and Wright (2007). They then removed time-varying default risk, as measured by the corporate issuer's distance to default. The Federal Reserve System's Board of Governors posts updates of this monthly time series at a permanent URL, from which we obtained them.[5] Figure 1 plots this series from its first measured value (in January 1973) through January 2020. By construction, this series has a zero mean, and its volatility increased at the turn of the century.



Figure 1: The Gilchrist and Zakrajšek (2012) Excess Bond Premium.

Textbook NK models often equate consumption, output, and labor supply. In DSGE models, capital accumulation breaks this equivalence. Nevertheless, in this empirical analysis, we measure economic activity with consumption gap to align closely with the theory. The consumption gap is defined as consumption's deviation from its potential value. We measure consumption using real personal consumption expenditures (real PCE), and we approximate potential consumption using the Congressional Budget Office's (CBO) estimate of real potential GDP.[6] Figure 2 plots the consumption gap alongside an alternative measure of economics activity, the unemployment gap, similarly constructed from the CBO's non-cyclical rate of unemployment. (The figure inverts the unemployment gap, so that it and the consumption gap comove pos-

[5] This page contains a description of the updating process.
[6] For this, we interpolate the quarterly real potential GDP series to a monthly frequency.

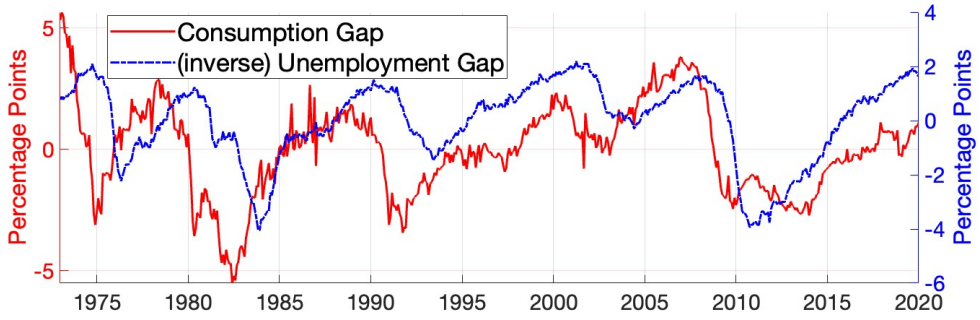


Figure 2: Consumption gap and (inverse) Unemployment gap.

itively.) As expected, movements in the unemployment gap have a significant lag to the movements in the consumption gap.

The FOMC assesses its achievement of price stability by comparing the price index for personal consumption expenditures (PCE) to its target, 2 percentage points on an annual basis. Because supply-driven changes in food and energy prices dominate its very short-run fluctuations, the FOMC pays more attention to core PCE inflation (which excludes food and energy price changes) when forecasting future inflation. We consider the supply-driven changes in food and energy prices to be "noise" when assessing the interactions of monetary policy with the choices of private firms setting (possibly) sticky prices, so we measure inflation with core PCE.

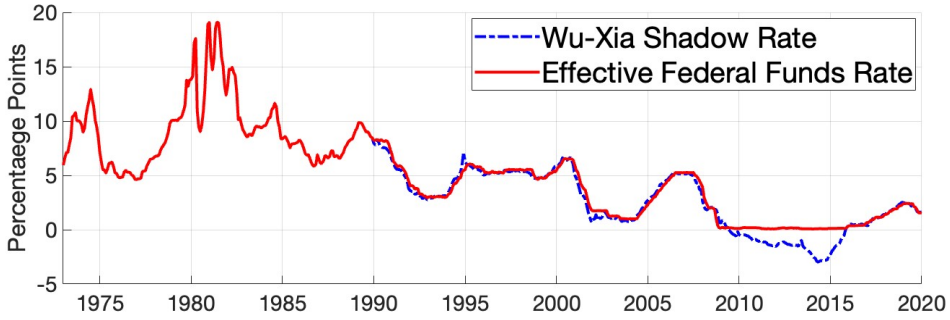


Figure 3: Fed Funds Rate and Wu-Xia Shadow Rate.

The FOMC communicates the stance of monetary policy with the federal funds rate, but this hit its effective lower bound in December 2008 and stayed there until December 2015. To account for the effects of unconventional monetary policy during this period, we measure monetary policy with the Shadow Rate of Wu and Xia (2016). This equals the intercept of a term structure model, which may (and indeed does) go

below zero. Figure 3 plots this shadow rate along with the effective federal funds rate (EFFR). The difference between the two series outside of the zero-lower-bound period reflects the term structure model's approximation error. For our work, we splice the EFFR for the period before 1990 with the Wu and Xia Shadow Rate thereafter.

4.2 The Baseline SVAR

The theory discussed above tells us to measure ζ_t (the shifter of demand for government bonds) with $i_t^P - i_t^G - \iota \hat{c}_t$. Given a particular value of ι, we do so using the EBP as $i_t^P - i_t^G$ and the consumption gap $\hat{c}_t$. For our SVAR's (partial) identification, we also assume that it is uncorrelated with other economic disturbances. This allows us to measure other variables' responses to innovations in ζ_t by placing it in the first place of a Cholesky identification scheme. The assumption that ζ_t evolves independently of other shocks implies that no other variables' lags should appear on the right-hand side of the SVAR's first equation. In our baseline specification, we do *not* impose those exclusion restrictions.[7] Our baseline specification also sets $\iota = 0$, so that the EBP directly reveals ζ up to scale. Some of our alternative specifications set $\iota > 0$.

In principle, our SVAR's four variables should be covariance stationary. However, it is well known that the average rate of inflation declined substantially during Alan Greenspan's tenure as the Federal Reserve's chair. To account for this, we removed linear trends from all four variables before estimation. We also report results below from alternative specifications which either drop this detrending step or add a quadratic term to the trend.

Our sample period begins in January 1985 and runs through January 2020. We begin in 1985 rather than at the start of the available data series (January 1973) both to avoid the heteroskedasticity associated with the great moderation and because Gilchrist and Zakrajšek (2012) show that the EBP insubstantially contributes to macroeconomic forecasts in the 1970s and early 1980s. We exclude the COVID-19 period to avoid introducing structural breaks or extraordinary shocks that could unduly influence the estimated relationships between the credit spread and macroeconomic variables. For our baseline specification, we include 12 lags of all four variables in the right-hand sides of all four equations.[8]

Figure 4 presents the responses of all four variables to a one percentage point shock to ζ_t. (Because $\iota = 0$, this shock is one-step-ahead error in forecasting the EBP.) The shaded areas give point-by-point 95 percent confidence intervals constructed using the

[7] Estimates of Specification 7, reported below, show that imposing this stricter form of exogeneity changes our results little.

[8] Some of the alternative specifications examined below expand this to 24 months.

recursive moving-block bootstrap of Brüggemann, Jentsch, and Trenkler (2016). The shock's effects on the EBP itself dissipate little during the five months after its impact. Thereafter, the EBP falls gradually, reaching its pre-shock level after about 18 months. On impact, the shock reduces the monetary policy rate very little, and it falls gradually about one percentage point over the next 18 months. It then takes about 18 to 30 months to return to its pre-shock level. Relative to the theory described above, this response to the shock is unambiguously "too late". The falls of inflation and the consumption gap on the shock's impact are consistent with this lack of timely accommodation and basic NK theory. Inflation returns relatively quickly to its pre-shock level, while the consumption gap continues falling in spite of the reduced monetary policy rate.

Figure 5 plots the share of each variable's k-step-ahead forecast error accounted for by the shocks to ζ_t along with 95 percent bootstrap confidence intervals analogous to those in Figure 4. As noted above, our SVAR includes lags of the consumption gap, inflation, or the monetary policy on the right-hand side of the EBP's forecasting equation. This allows shocks to these other variables to contribute to the EBP's long-run variance. Our point estimates indicate that these other shocks account for about 32 percent of the EBP's long-run variance.[9] The shocks to ζ_t account for little of the other variables' forecast errors over short horizons, but they gain influence over the consumption gap and monetary policy at longer horizons. In the long run, shocks to ζ_t account for approximately 20 percent of the consumption gap's variance and nearly 25 percent of the interest rate's variance. The long-run variance share of shocks to ζ_t on inflation is much more modest, 9 percent.

Taken together, these estimates suggest that having monetary policy respond to ζ_t as it changes instead of when those changes effects manifest themselves could significantly reduce the variance of economic activity and modestly reduce that of inflation. To assess this change's consequences in a historical context, we have used our SVAR estimates to recreate the EBP, the consumption gap, and inflation as if there had been no shocks to the EBP. The northeast, northwest, and southwest panels of Figure 6 plot these counterfactual series along with the original data. The EBP shocks have very little effects on any of these series until the turn of the century. Thereafter, the counterfactual EBP and consumption gap move much less than their historical counterparts. Although this is particularly the case immediately before and during the great recession, the stabilization of the consumption gap from 2000 to 2009 is substantial. As suggested by the forecast error variance decompositions, the counterfactual inflation series differs little from its historical value.

[9] A test of the null hypothesis that the first equation excludes lags of these three variables, discussed in more detail below, rejects at the 1 percent level.

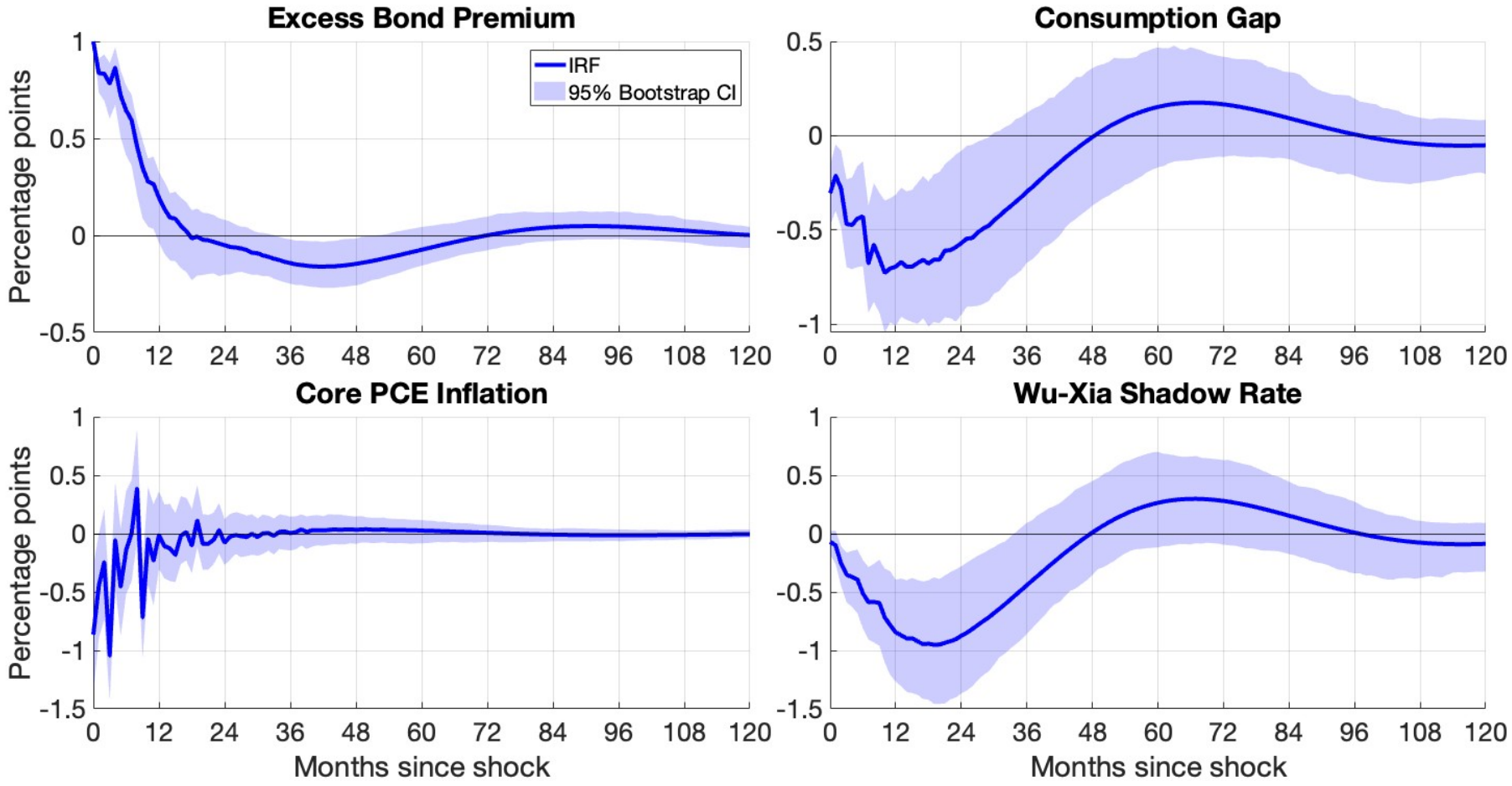


Figure 4: Responses to a shock to ζ_t.

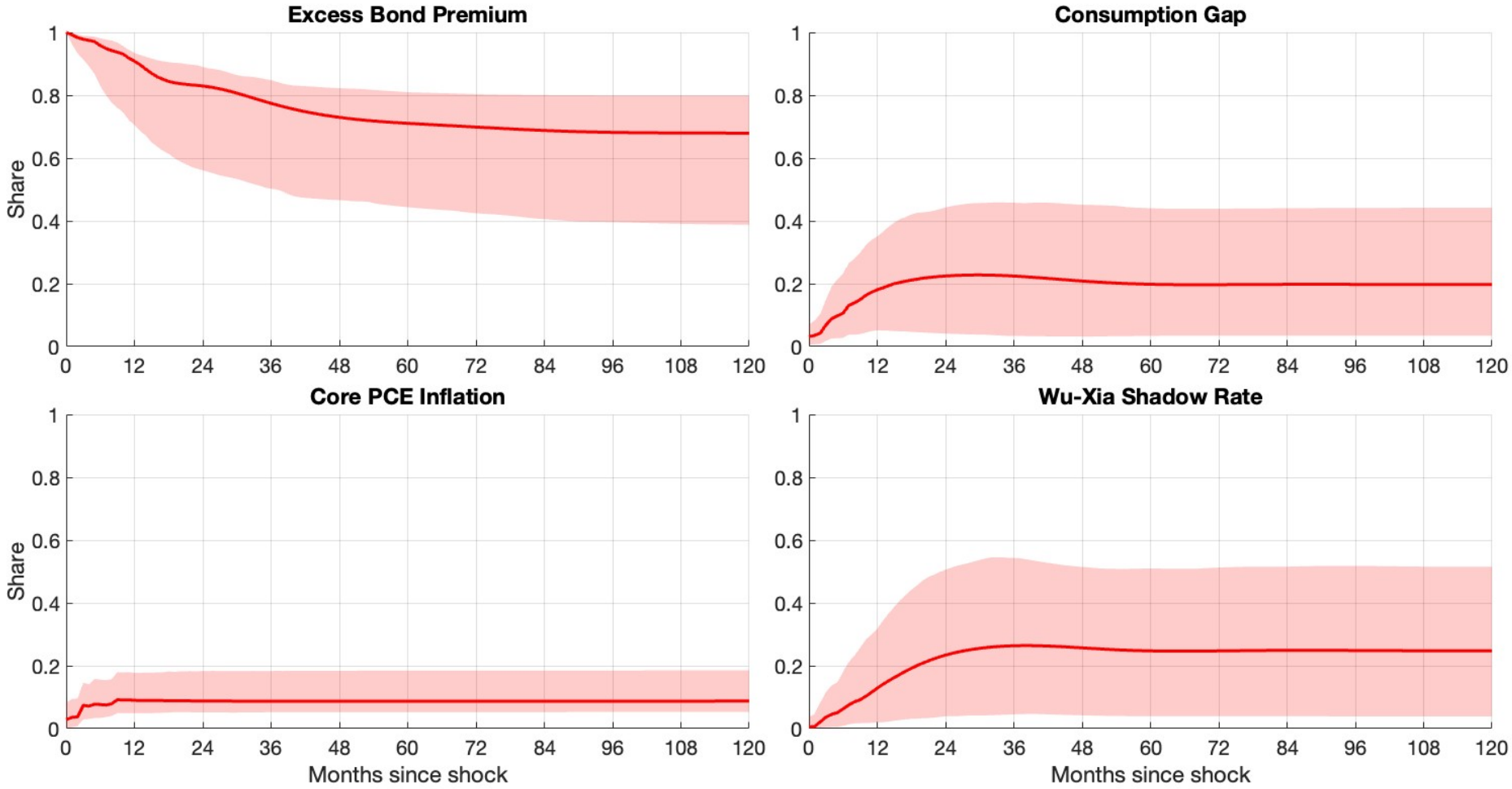


Figure 5: Proportions of forecast variance explained by shocks to ζ_t.

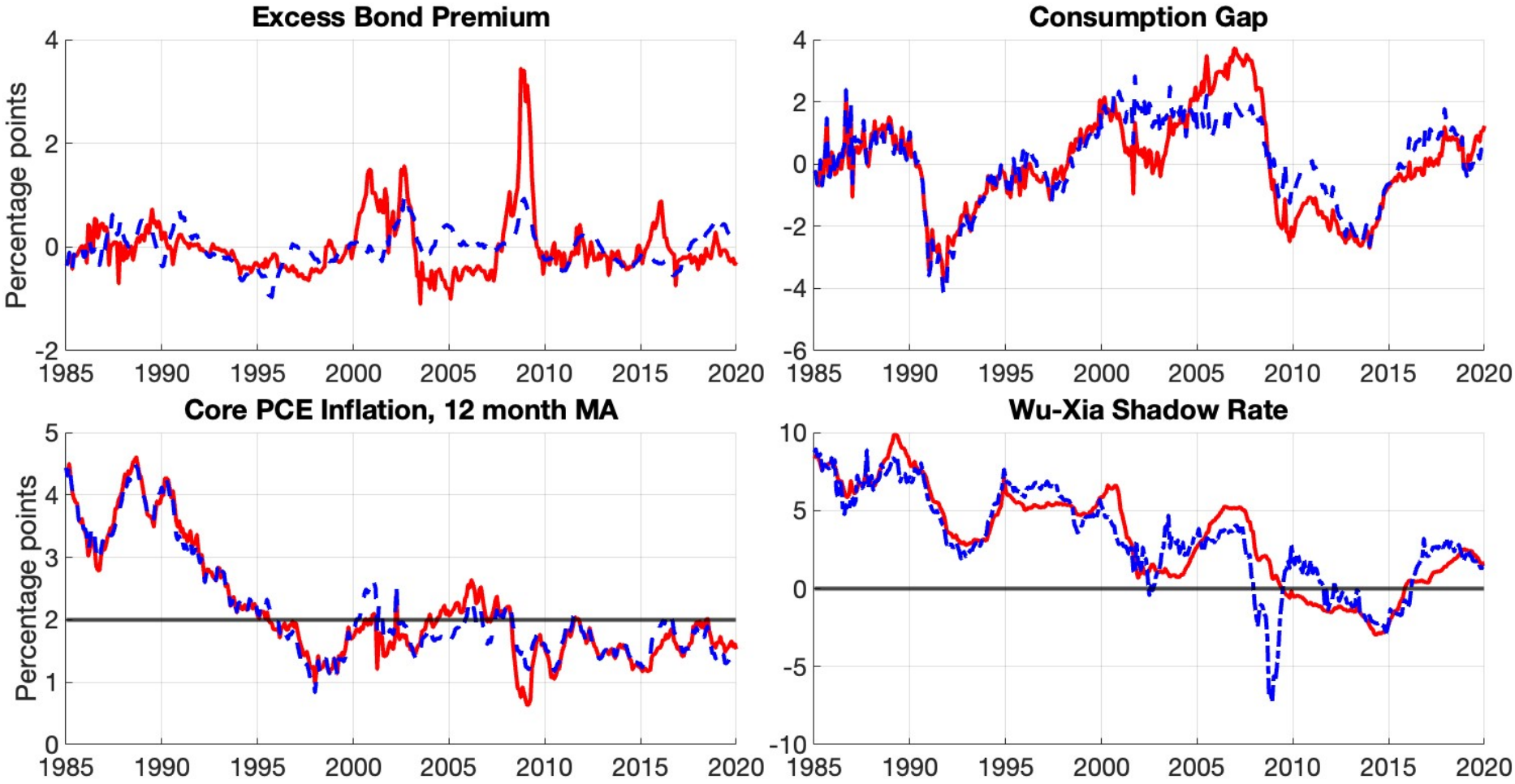


Figure 6: Counterfactual Policy Simulations[i]

(i) The northwest panel plots the evolution of the Excess Bond Premium after turning off its one-step-ahead forecast error. This illustrates the contribution of these shocks to the EBP's evolution, but it does *not* equal the EBP in our counterfactual policy experiment.

#	Specification	ι	Percentage of variance from ζ_t's shock				Months
			ζ_t	Activity	Inflation	Interest rate	below ZLB[ii]
1	**Baseline**	**0**	**68 (10.4)**	**20 (11.1)**	**9 (3.4)**	**25 (12.7)**	**74**
2	Omit trend adjustment	0	68 (10.0)	14 (10.1)	8 (3.3)	13 (10.5)	74
3	Add quadratic trend adjustment	0	69 (10.2)	20 (11.5)	9 (3.4)	24 (12.0)	73
4	Add 12 more lags (total 24)	0	65 (9.9)	27 (11.9)	11 (3.4)	33 (13.6)	71
5	Begin sample in January 1973	0	79 (7.8)	12 (7.4)	9 (5.1)	31 (11.4)	66
6	Place EBP last	0	65 (10.8)	16 (10.6)	5 (2.4)	16 (11.2)	78
7	Impose exogenous EBP	0	100 (0)	31 (17.9)	9 (5.0)	34 (17.5)	74
8	Replace consumption gap with unemployment gap	0	76 (9.4)	36 (14.5)	10 (3.9)	31 (13.8)	64
9[iii]	Add unemployment gap	0	70 (10.1)	26 (11.1)	9 (3.4)	28 (11.5)	71
	Specifications with $\iota > 0$						
10	Baseline	0.18	74 (9.9)	37 (12.9)	10 (3.5)	39 (13.9)	75
11	Omit trend adjustment	0.08	71 (9.6)	19 (11.3)	8 (3.3)	17 (11.6)	69
12	Add quadratic trend adjustment	0.18	75 (9.6)	38 (12.9)	10 (3.2)	36 (12.8)	79
13	Add 12 more lags (total 24)	0.20	71 (10.6)	45 (14.2)	13 (3.8)	48 (15.1)	76
14	Begin sample in January 1973	0.12	84 (7.8)	22 (9.4)	9 (5.2)	38 (12.1)	67
15	Place ζ_t last	0.04	66 (10.7)	16 (10.6)	5 (2.4)	16 (11.2)	80
16	Replace consumption gap with unemployment gap	0.08	77 (9.2)	39 (14.6)	10 (3.9)	34 (14.0)	77
17[iii]	Add unemployment gap	0.15	74 (9.9)	36 (12)	10 (3.5)	35 (12.1)	79

Table 1: SVAR Statistics[i]

(i) Numbers in parentheses are standard errors calculated with the recursive block bootstrap of Brüggemann et al. (2016).
(ii) Over the sample period, the federal funds rate was below its effective lower bound for 77 months.
(iii) Specifications 9 and 17 have five variables, with unemployment gap ordered after consumption gap. The reported percentage reduction in economic activity for these specifications is that for the consumption gap.

The southeast panel of Figure 6 plots our estimate of the monetary policy rate required to achieve the consumption and inflation outcomes in the other panels. For this, we first calculate the interest rate as if there had been no shocks to ζ_t. We then subtract our measure of ζ_t from this.[10] As the other variables' counterfactual values lead us to expect, the counterfactual interest rate series differs little from the historical series before the turn of the century. Thereafter, it leads the historical series. For example, the interest rate would have reached the zero lower bound in January 2008, approximately eighteen months earlier than observed in the actual data. the counterfactual Shadow Rate dips to nearly -7 percent. (Whether the FOMC had or has the unconventional policy tools to achieve such remarkable accommodation is an open question.) However, the counterfactual Shadow Rate rises above zero in June 2009, at almost the same time that the historical Shadow Rate fell below zero. It then crosses below zero again in June 2011, remaining below it until February 2016. During this second trip to the ZLB, the counterfactual Shadow Rate is at or above the actual rate. In total, the counterfactual Shadow Rate spent 74 months below the ZLB, three months *less* than did the actual Shadow Rate. Outside of the period immediately before and during the great recession, the counterfactual Shadow Rate fluctuates somewhat *less* than the historical rate. Once accounting for the counterfactual Shadow Rate's dramatic drop in 2008, the two series have almost identical standard deviations, 3.11 and 3.15 percentage points. We conclude from this exercise that responding contemporaneously to ζ_t would not dramatically increase either the policy rate's volatility or the time spent employing unconventional policy tools.[11]

4.3 Alternative SVAR Specifications

Our SVAR-based policy analysis required answering many questions which lack empirically definitive answers, such as "What trend adjustment should be applied to the data," "How many lags appear in the forecasting equations," and "What is the consumption elasticity of government bond demand?" We have assessed how our em-

[10] The text to this point has glossed over a practical difficulty with our exercise. The EBP averages credit spreads between bonds with different maturities, while the theoretically-relevant spread is that constructed from one-period bonds. If we add longer duration zero coupon bonds to the theoretical framework, it says that the log-linearized credit spread based on k period bonds equals $(\zeta_t + \iota\hat{c}_t)/k$. Therefore, the EBP is proportional to the one period spread, with a factor of proportionality equal to the average inverse duration of the securities used in its construction. The average duration of the securities used by Gilchrist and Zakrajšek (2012) equals 6.47 years. When the duration of the securities underlying the EBP is uniformly distributed from 1 to 156 (13 years), we must multiply the EBP by 2.3 to retrieve the one-period spread. We do so to construct the southeast panel of Figure 6. Integrating this issue and its consequences into our analysis is on our agenda for this paper's revision.

[11] Subjecting this conclusion to further scrutiny by changing the maturity adjustment factor discussed in Footnote 10 is on our agenda.

pirical results change as we change these answers by estimating a set of alternative specifications. Table 1 reports statistics from the resulting estimated SVARs. For ease of reference, we have assigned each estimated specification a number, with our baseline getting "1". These numbers appear in the table's first column. The second column describes the change to the baseline specification, while the third column gives the value of ι used to identify ζ_t. The next four columns give the shares of long-run variance arising from shocks to ζ_t. A bootstrap standard error appears after each of these in parentheses. The last column reports the number of months that the associated counterfactual Shadow Rate was below the zero lower bound in the sample. Specifications 2 through 9 share the baseline's assumption that $\iota = 0$. For the remaining specifications, we estimate ι with a procedure described below.

Specifications 2 through 5 embody changes to basic econometric choices. Specifications 2 and 3 examine the impact of linear detrending by either estimating with the raw data (Specification 2) or by removing a quadratic trend (Specification 3). Removing the trend adjustment reduces the share of the consumption gap's variance explained by ζ_t's shock from 20 to 14 percent and reduces the analogous share of the interest rate's variance from 25 to 13 percent. Otherwise, this adjustment changes the results little. The estimates after performing a quadratic trend adjustment hardly differ from those of the baseline. Specification 4 adds 12 lags to the forecasting equations' right-hand sides (for a total of 24 lags). This change modestly increases the share of variance in economic activity explained by the EBP, suggesting that some of its effects unfold over longer horizons. Specification 5 begins the sample in January 1973 instead of in January 1985. Since Gilchrist and Zakrajšek (2012) report that their EBP has little predictive power for macroeconomic aggregates before 1985, we expect this to reduce the share of the consumption gap's variance explained, which it does.

Specifications 6 and 7 change how we treat the identifying assumption that ζ_t follows an exogenous process. Specification 6 identifies a shock to ζ_t after conditioning on the realizations of all other variables. This reduces the shares of the consumption gap's variance and the interest rate's variance explained both to 16 percent. Specification 7 moves the analysis in the opposite direction by excluding lags of the other three variables from the equation explaining the EBP. This raises the share of the consumption gap's variance explained by ζ_t to 31 percent and raises the analogous share for the interest rate to 31 percent.

The next two specifications investigate how our results change when we measure economic activity differently. For specification 8, we replaced the consumption gap with the unemployment gap plotted in Figure 2. The most notable change from this is an increase in the share of economic activity variance due to ζ_t, from 20 percent to

36 percent. This specification's impulse response functions (not reported) show that unemployment responds to a ζ_t shock with a considerable lag. Specification 9 expands the SVAR to five variables by adding the unemployment gap. This also increases the share of the consumption gap's variance explained to 26 percent. The analogous variance share for the unemployment gap (not listed in the table) equals 32 percent (with a standard error equal to 12.5 percent).

For the remaining specifications in Table 1, we estimated ι by choosing the value that maximizes the share of $\zeta_t \equiv i_t^P - i_t^G - \iota\hat{c}_t$'s long-run variance explained by its own one-step-ahead forecasting error. Figure 7 illustrates this for a specification otherwise equivalent to our baseline. The figure plots the share of the variance of ζ_t as a function of the assumed value of ι. Although we could require $\iota > 0$ ex ante, we in fact also allowed for $\iota < 0$. These negative values substantially reduce the ζ_t's long-run variance driven by its own forecasting error. The estimated value of ι equals 0.18. However, the figure shows that choosing other positive values of ι would change the criterion function modestly.

Table 1 reports the statistics from this version of our baseline specification with $\iota = 0.18$ as Specification 10. This change raises the fraction of the consumption gap's long-run variance explained by shocks to ζ_t to 37 percent, and the analogous share of the interest rate's variance increases to 39 percent. The remaining rows in Table 1 report results from analogues to Specifications 2 through 6, 8, and 9. There is no analogue for Specification 7 because it presumes that the criterion we use to select ι equals its maximum value at $\iota = 0$. All of the estimated values of ι are positive, and the range from 0.04 to 0.20. By construction, the shares of ζ_t's variance arising from its own shock increase. However, these changes are modest. As with the baseline specification, allowing for $\iota > 0$ generally increases the variance shares of the activity measure and the interest rate explained by the shock to ζ_t.

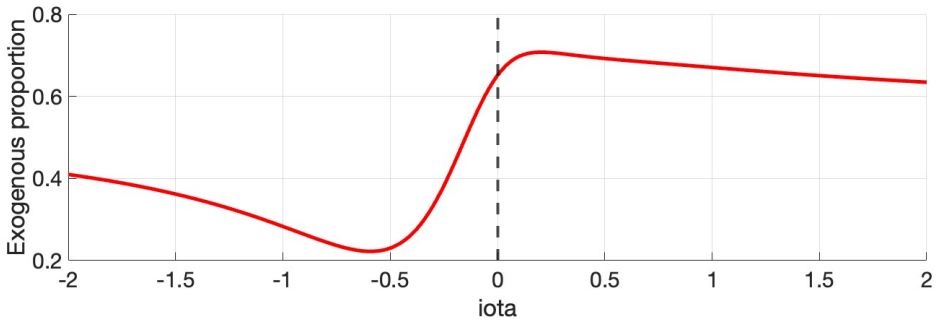


Figure 7: Identification of ι.

4.4 Alternative Spreads

For the reasons described above, we believe that the Gilchrist and Zakrajšek (2012) EBP measures the theoretically appropriate object for the question at hand. Nevertheless, we wish to know how our results change after either replacing the EBP with another possibly relevant credit spread or after simply adding that spread to our SVAR. Table 2 shows results analogous to those of Table 1 for specifications with alternative credit spreads.



Figure 8: GZ default risk compensation

Gilchrist and Zakrajšek's (2012) credit spread sums the EBP with compensation for default risk. Figure 8 plots this, and Specifications 18, 19, and 20 report the consequences of using this in our SVAR. For specification 18, we replaced the EBP with the default risk compensation series. Gilchrist and Zakrajšek report that default risk compensation has little forecasting power for macroeconomic outcomes, and these results confirm this. For Specifications 19 and 20, we added the GZ default risk compensation to our VAR system. Specification 19 keeps but the EBP in the first position, while Specification 20 places the EBP in the second position, behind the GZ default risk compensation. Both specifications' results are close to those from the baseline.

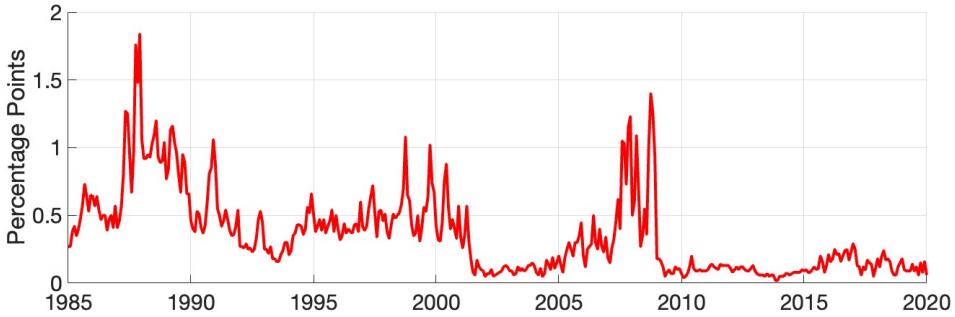


Figure 9: The 3-month non financial commercial paper - treasury bill spread

Specifications 21, 22, and 23 perform a similar exercise using the spread between 3-month AA non financial commercial paper and 3-month Treasury bills, which we plot in Figure 9.[12] Friedman and Kuttner's (1992) found that a similar spread had substantial forecasting power for economic activity. Specification 21 replaces the EBP with this, and the results indicate that its shocks account for 15 percent of the consumption gap's long-run variance and almost no long-run variance of inflation or the interest rate. Specifications 22 and 23 include both the EBP and the 3-month commercial paper-Treasury spread. The EBP shock's contributions to long-run variance depend little on whether it is ordered first or second, suggesting that the two spreads contain uncorrelated information. Adding the 3-month commercial paper-Treasury spread to the system reduces the share of the consumption gap's long-run variance explained by EBP shocks to about 12 percent. Although we share Gilchrist and Zakrajšek's (2012) opinion that the EBP has "high information content," the relevance of monetary policy responses to private-government credit spreads does not depend on this specific spread accounting for variance even in the presence of other possibly-relevant private-government spreads. Therefore, we have calculated the shares of variance accounted for by forecast errors to both spreads. These equal 30 percent for the consumption gap, 14 percent for inflation, and 33 percent for the interest rate.

[12]The series of non financial commercial paper rates (FRED mnemonic CPN3M) is available after 1997. We construct our spread before 1997 using rates on all commercial paper (FRED mnemonic CP3M). There are 4 data points missing in series SPN3M. We interpolate those data points.



Figure 10: Moody's AAA-BAA spread.

Gilchrist and Zakrajšek (2012) compare their spread to the AAA–BAA corporate bond spread, which we plot in Figure 10. They describe this and a commercial paper-treasury spread as " widely-used default risk indicators." Indeed, this spread's correlation with the GZ default compensation series over our sample period equals 0.55. Surprisingly to us, its correlation with the EBP is even higher, 0.71. Specifications 24, 25, and 26 incorporate this spread. In Specification 24, the AAA-BAA spread replaces the EBP. In spite of its high correlation with the EBP, shocks to this spread explain ten percent of the consumption gap's variance, five percent of inflation's variance, and six percent of the interest rate's variance. Adding the AAA-BAA spread to the SVAR while keeping the EBP in the first position changes the results little. Placing the AAA-BAA spread in the first position reduces the EBP's share of the consumption gap's variance to 12 percent and slightly increases its share of inflation's variance to 11 percent. We hypothesize that the high correlation between the EBP and the AAA-BAA spread arises from the ability to quickly turn AAA securities into cash in the repo market. Empirically scrutinizing that hypothesis is on our research agenda.

#	Specification	Ordering	Percentage of variance from the spread shock				Months
			Own shock	Activity	Inflation	Interest rate	below ZLB[ii]
1	**Baseline**	–	**68 (10.4)**	**20 (11.1)**	**9 (3.4)**	**25 (12.7)**	**74**
18	Replace EBP with GZ default risk compensation	–	44 (12.2)	4 (6.5)	6 (2.8)	22 (9.8)	97[iii]
19	Add GZ default risk compensation	EBP first	60 (10.1)	17 (10.3)	7 (2.8)	16 (10.5)	76
20	Add GZ default risk compensation	EBP second	67 (10.6)	19 (11.6)	9 (3.3)	28 (12.5)	67
21	Replace EBP with 3M CP–TB spread	–	61 (10.2)	15 (12.6)	3 (2.5)	2 (7.5)	75[iii]
22	Add 3M CP–TB spread	EBP first	60 (9.4)	13 (8.9)	11 (3.6)	31 (12.2)	76
23	Add 3M CP–TB spread	EBP second	57 (9.3)	12 (8.3)	11 (3.6)	31 (12.2)	76
24	Replace EBP with AAA–BAA spread	–	49 (11.5)	10 (7.8)	5 (2.8)	6 (7.1)	84[iii]
25	Add AAA–BAA spread	EBP first	65 (9.8)	17 (10.1)	8 (3.2)	24 (12.3)	73
26	Add AAA–BAA spread	EBP second	55 (9.9)	12 (9.6)	7 (2.7)	24 (13.3)	77

Table 2: SVAR Statistics[i]. Alternative Spreads. Monthly Data.

(i) Numbers in parentheses are standard errors calculated with the recursive block bootstrap of Brüggemann et al. (2016).
(ii) Over the sample period, the federal funds rate was below its effective lower bound for 77 months.
(iii) We take away the trend when calculating months at ZLB for GZ default risk compensation, AAA–BAA and 3M CP/TB spreads.

#	Specification	Percentage of variance from ζ_t's shock				Quarters
		Credit Spread	Activity	Inflation	Interest rate	below ZLB[ii]
27	Baseline (Quarterly data)	67 (12.0)	22 (14.4)	11 (8.2)	25 (14.3)	23
28[iv]	Add Financial Sector's Equity Share	65 (11.1)	20 (11.7)	13 (8.5)	26 (12.8)	24
29[iv]	Add Share of current liabilities (Manufacturing)	67 (10.6)	30 (14.1)	10 (6.6)	29 (13.4)	20

Table 3: SVAR Statistics[i]. Other Financial Variables. Quarterly Data.

(i) Numbers in parentheses are standard errors calculated with the recursive block bootstrap of Brüggemann et al. (2016).
(ii) Over the sample period, the federal funds rate was below its effective lower bound for 26 quarters.
(iv) Ordering new variables first in the system.

5 Financial Imperfections

We have viewed credit spreads through the lens of Fisher's (2015) New Keynesian IS curve, which omits any discussion of credit market imperfections. Nevertheless, an alternative interpretation of credit spread fluctuations, favored by Gilchrist and Zakrajšek (2012), is that they embody changes in the relevance of financial frictions. To assess this alternative interpretation, we introduce two distinct indicators into the SVAR framework and examine whether the share of variance attributed to the EBP changes. Since both indicators are available only at the quarterly frequency, we estimate a quarterly SVAR model with four lags, covering the period from 1985Q1 to 2019Q4. The impulse responses for the baseline case for quarterly data mimic those from the monthly data, and are shown on Figure 15 in Appendix B. The first row of Table 3 reports the variance shares and time at the zero lower bound from this quarterly specification. These are very close to those estimated from monthly data.

The first variable we consider measures the U.S. financial sector's capitalization. To this end, we construct a series the financial sector's equity share using data from the Financial Accounts of the United States (Z.1). Specifically, we calculate the balance sheet ratio of total equity to total assets, excluding the monetary authority from both the numerator and the denominator.[13] Figure 11 illustrates the proportion of own capital held by institutions in the financial sector relative to their total assets. The correlation of this series with the EBP over our sample period is negative but modest, -0.26.

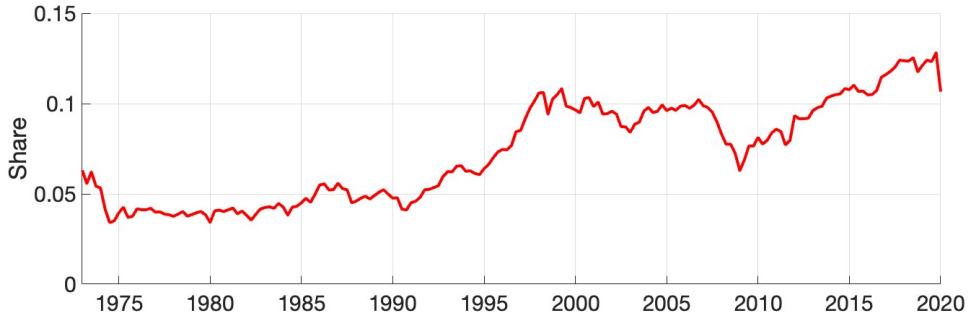


Figure 11: The Financial Sector's Equity Share (Quarterly data).

Specification 28 includes this financial balance sheet as the first variable in the SVAR. This change leaves the results from the baseline quarterly specification almost

[13]The numerator is computed as "Domestic Financial Sectors: Total Liabilities" minus "Monetary Authority: Total Liabilities". The denominator is calculated as "Domestic Financial Sectors: Total Liabilities and Equity" minus "Monetary Authority: Liabilities and Equity". The equity share is one minus the resulting ratio.

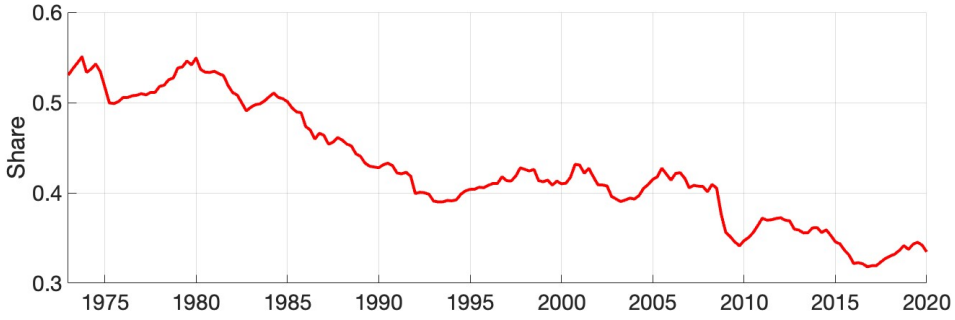


Figure 12: Manufacturing Sector's Current Liabilities Share (Quarterly data).

unchanged. However, we wish to note that shocks to this variable explain 28 percent of the consumption gap's variance. That is, both the EBP credit spread and the state of the financial sector's balance sheet matter for economic outcomes.

The considerable literature on economic fluctuations with financial constraints on producers' indebtedness motivates our choice of the second variable we examine, Second measure share of current liabilities to total liabilities of the Manufacturing sector in the United States. We construct this time series from the Quarterly Financial Report (QFR) historical data. This data provides statistics on the financial results and position of U.S. corporations. We focus on manufacturing, as it has any manufacturing corporation with assets of $\$250,000$ and more reported. Figure 12 shows this series.

Specification 29 includes this as the first variable in the SVAR this modification *increases* the share of the consumption gap's long-run variance explained by shocks to the EBP from 20 percent to 30 percent. Otherwise, adding this liability ratio changes our results little.

6 Conclusion

How to respond to credit spreads arose at a critical point of recent macroeconomic history during the June 2008 FOMC meeting. Presidents Jeffrey Lacker and Charles Plosser, Governors Randall Kroszner and Donald Kohn, and Chairman Ben Bernanke discussed the relevance of elevated private borrowing rates in the presence of a historically low monetary policy rate. Governor Kroszner stated " ... the way in which our actions are being translated into market prices is somewhat different from the way it was when the spreads were lower." President Lacker closed the discussion with "Spreads always rise in recessions. We always lower real rates in recessions. So to say that the fact that spreads have risen by itself doesn't make this an exceptional cir-

cumstance."[14] Our work provides a quantitative perspective on that discussion. In that month, the Wu-Xia Shadow Rate equalled 2 percent, while our estimate of the rate if the FOMC responded to spreads one-for-one was -1.15 percent. That is, combining the theory of the divine coincidence with measured credit spreads indicates that the FOMC's historically low policy rate exceeded the appropriate rate at that moment by 315 basis points.

Although the standard errors on our estimates are not small, they and the basic theory of the divine coincidence together indicate a possibility that the FOMC can get closer to achieving its dual mandate by responding to credit spreads *when they change* instead of waiting until their effects on the macroeconomy become evident. Doing so requires overcoming two practical implementation obstacles. First, the FOMC must measure the appropriate credit spread in real time. We have focused on the EBP, but combining this with other spreads seems useful to us. Second, the traders with the System Open Market Account (SOMA) at the New York Fed must learn how to impact private borrowing rates through their trades in the federal funds and Treasury repo markets. We believe that this will require more research into relevant financial market trading conventions, which all simple New Keynesian economies omit.

[14] Please see pages 117-119 in the meeting transcript. We thank John Fernald, who attended that meeting, for bringing this discussion to our attention.

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A Impulse Responses for the alternative spreads (Monthly Data)

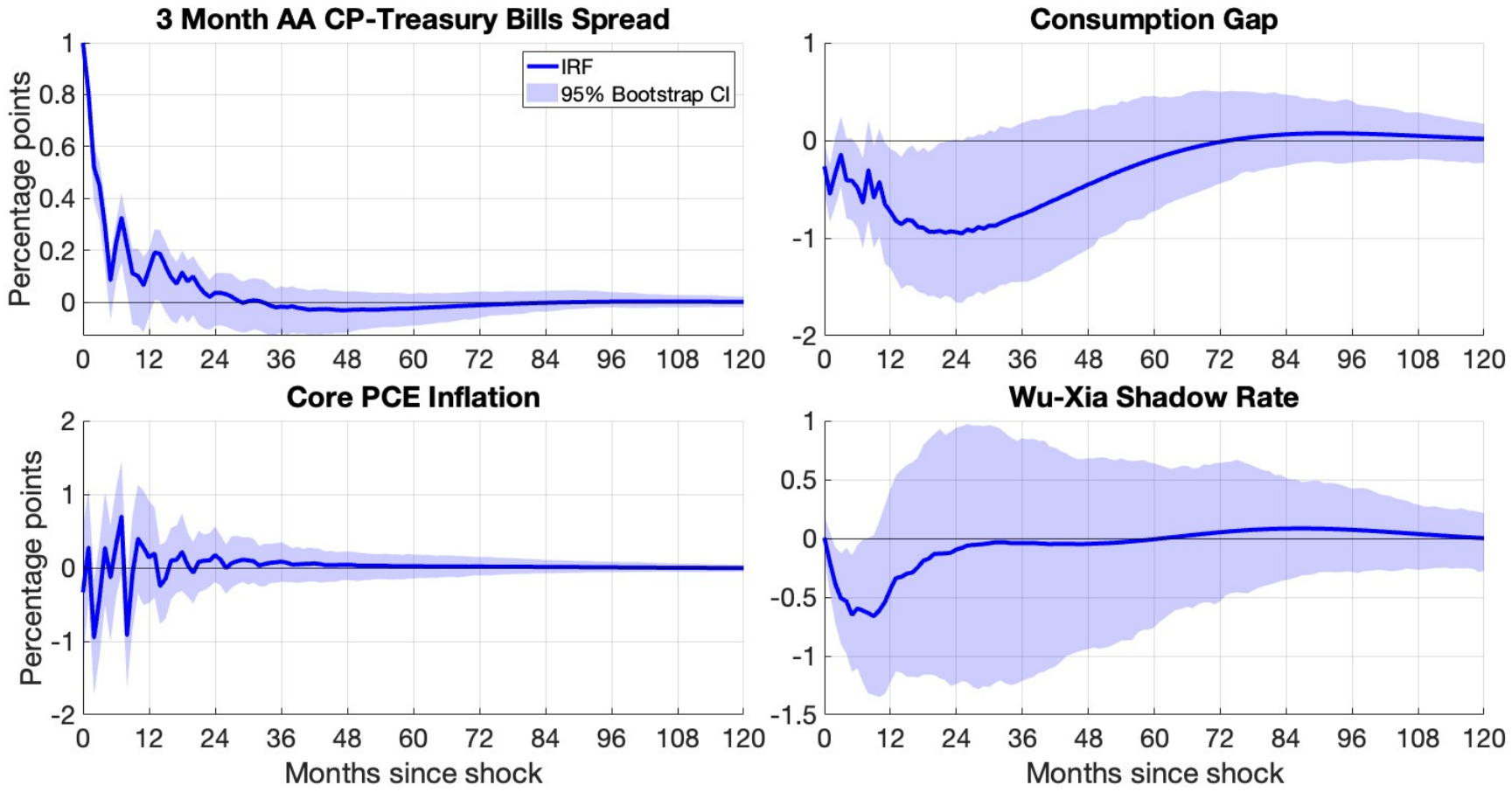


Figure 13: Impulse responses to the AAA-BAA spread shock, 95% Bootstrap CI

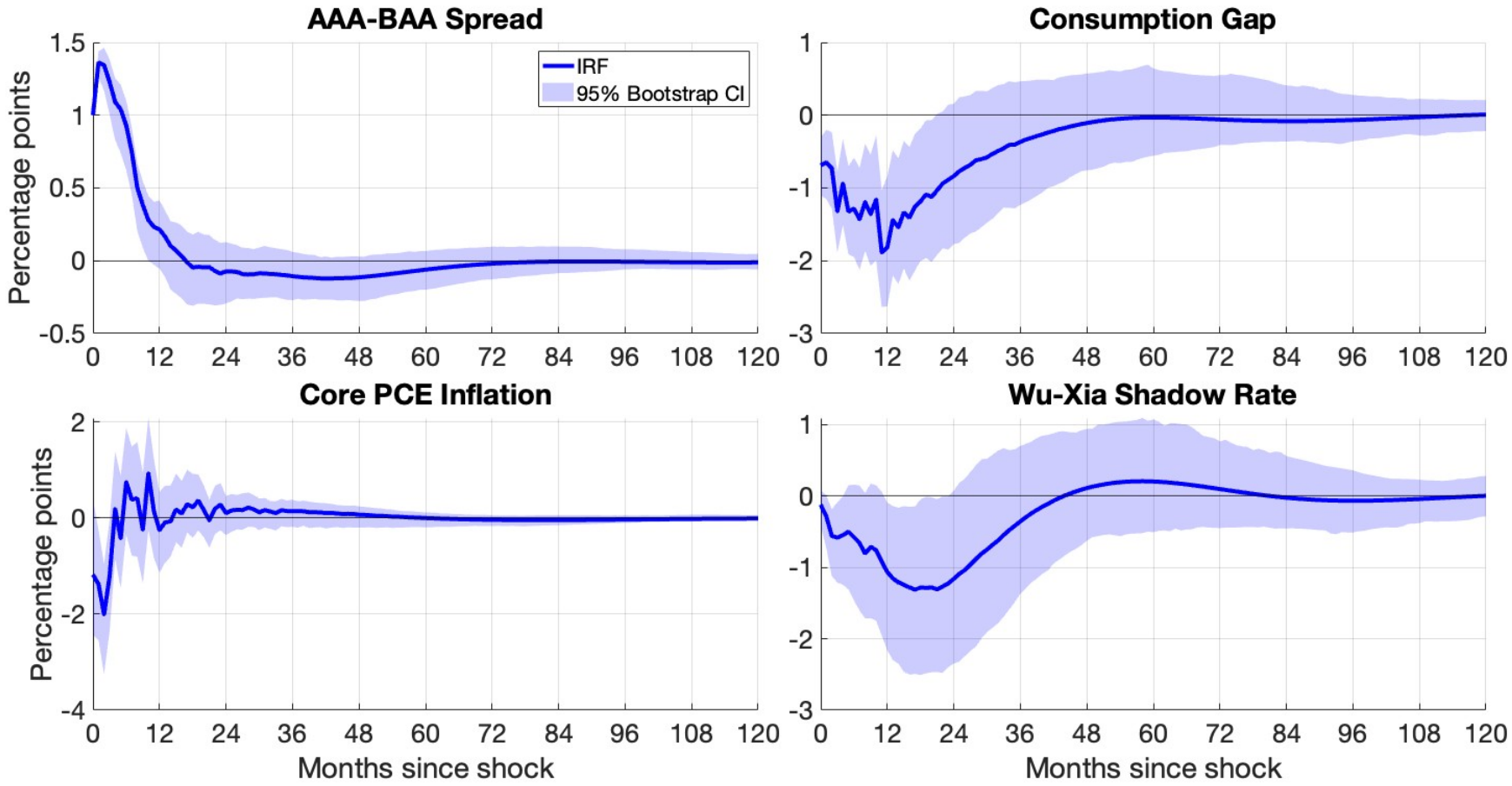


Figure 14: Impulse responses to the 3M Corporate/T-bills spread shock, 95% Bootstrap CI

B Impulse Responses for baseline and financial varaibles (Quarterly Data)

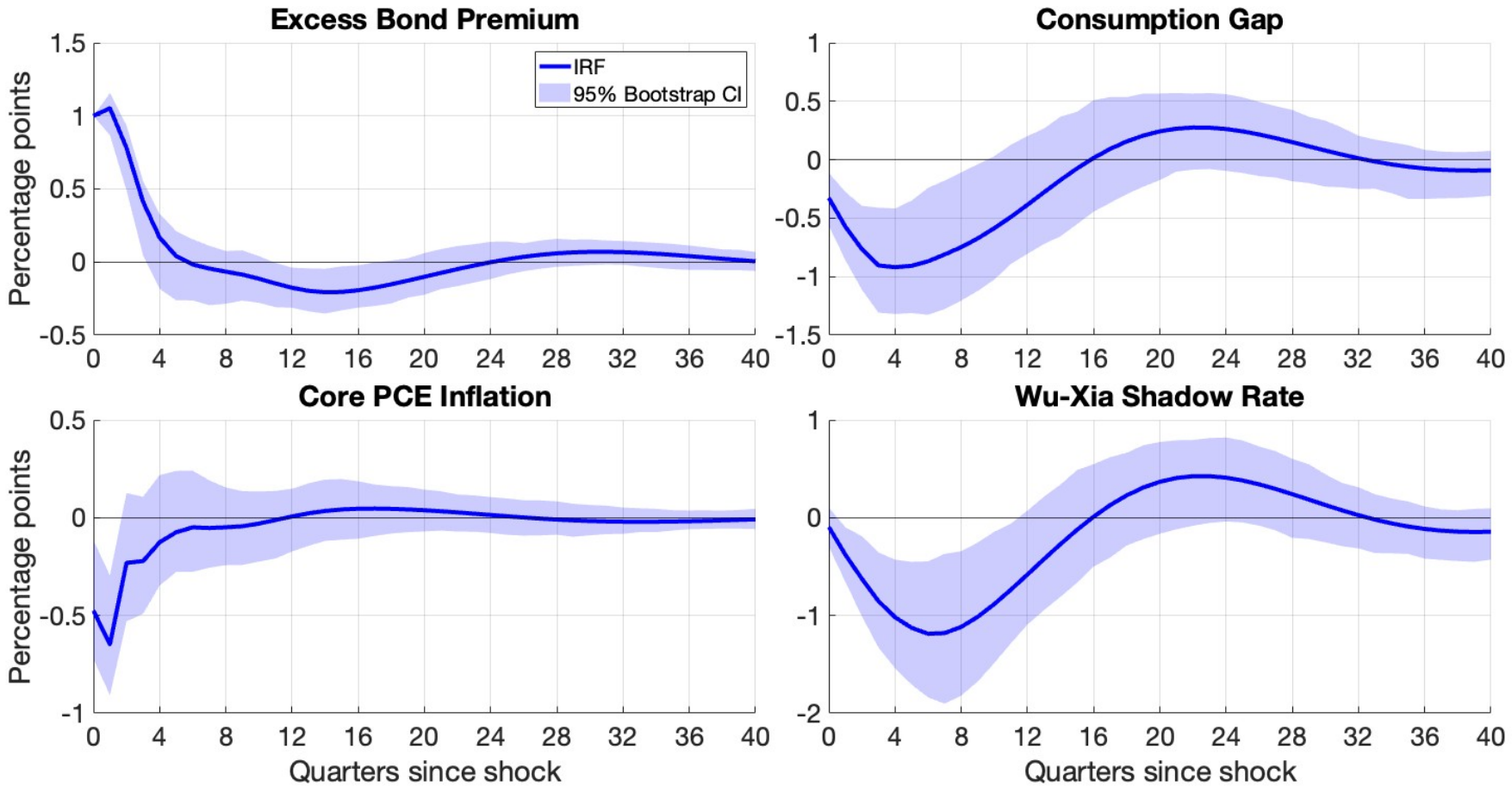


Figure 15: Impulse responses to the excess bond premium shock using quarterly data, 95% Bootstrap CI.

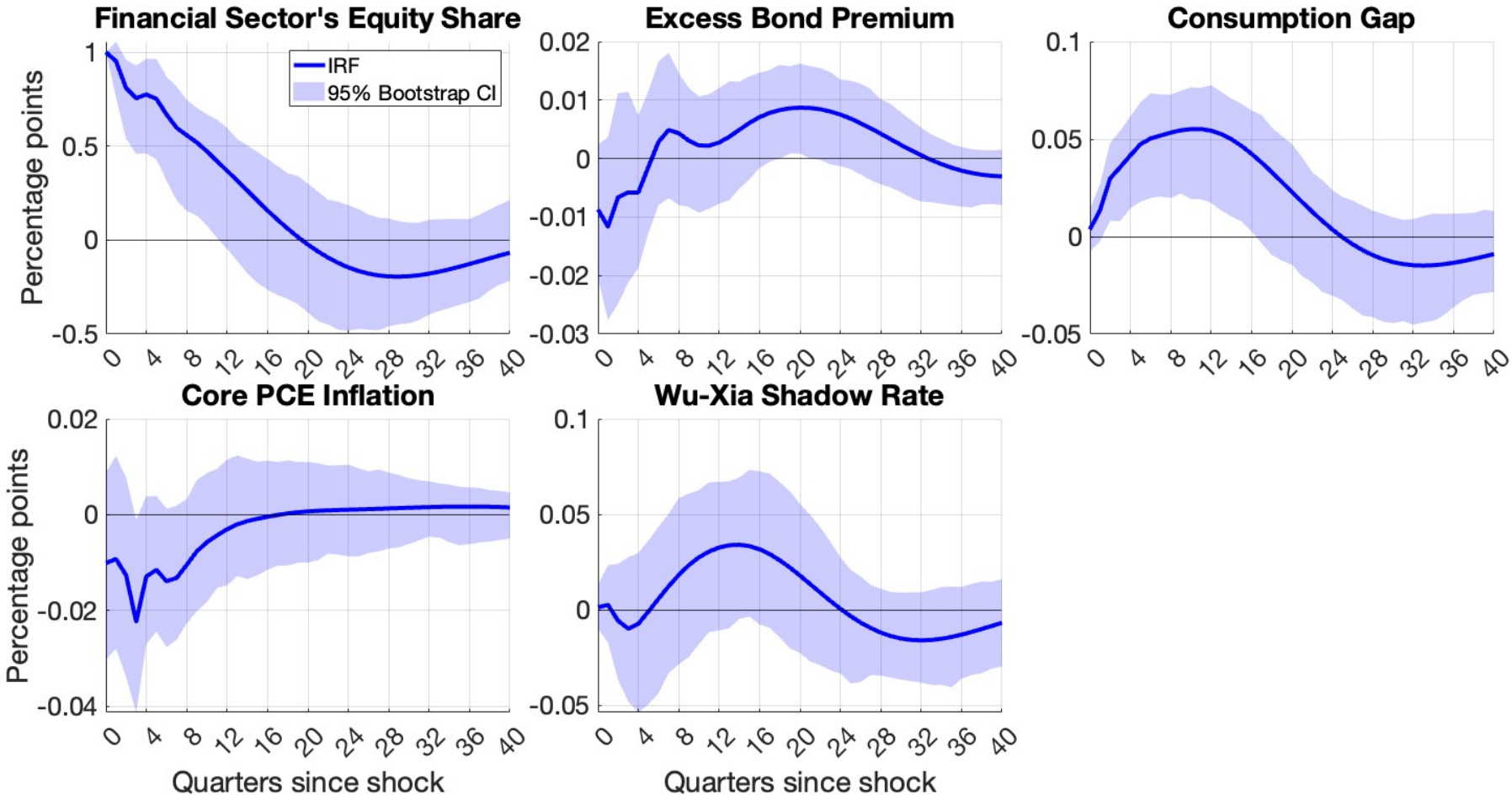


Figure 16: Impulse responses to the financial sector equity shock (as a share of total assets), 95% Bootstrap CI.

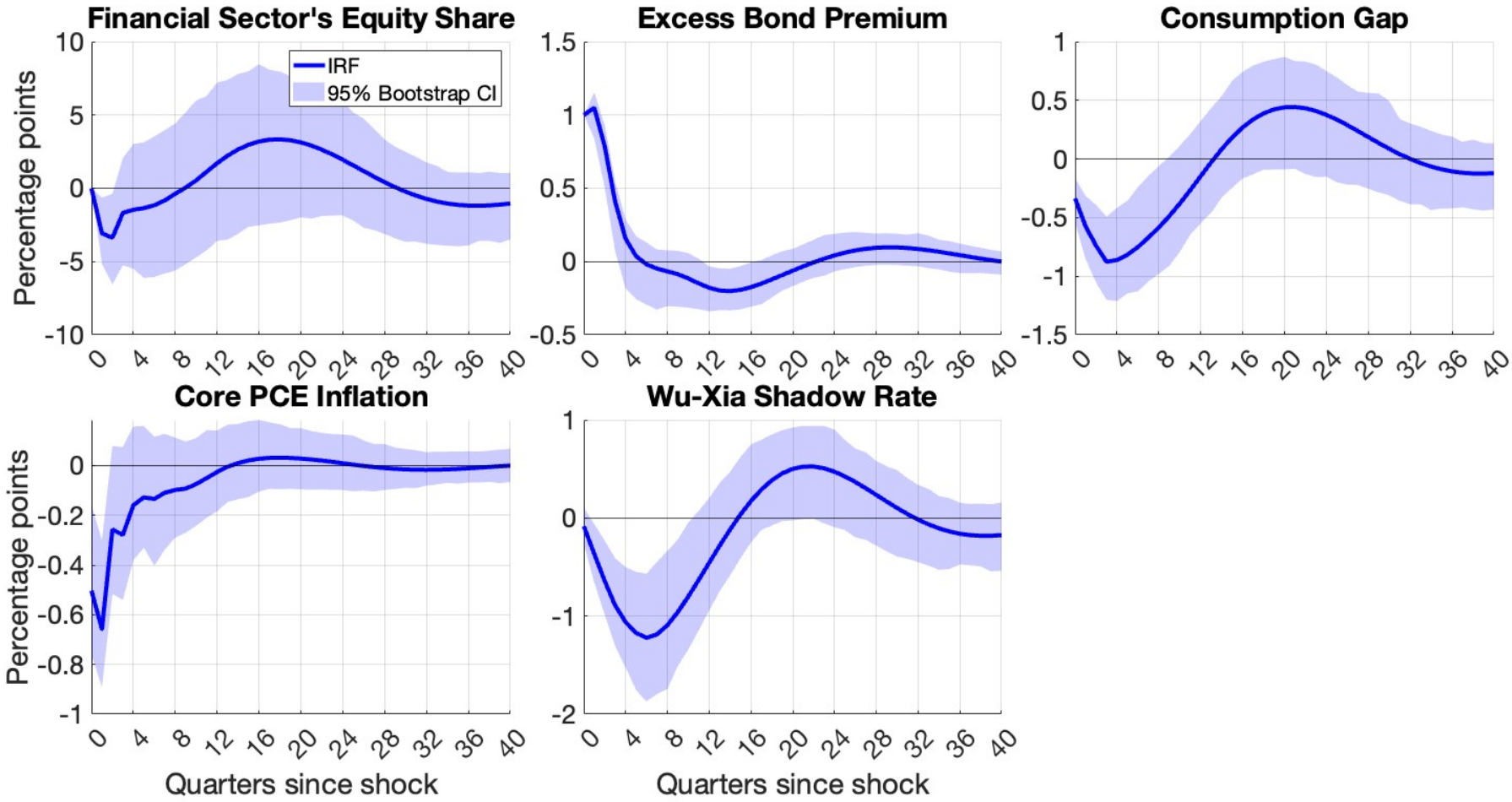


Figure 17: Impulse responses to the excess bond premium shock, 95% Bootstrap CI.

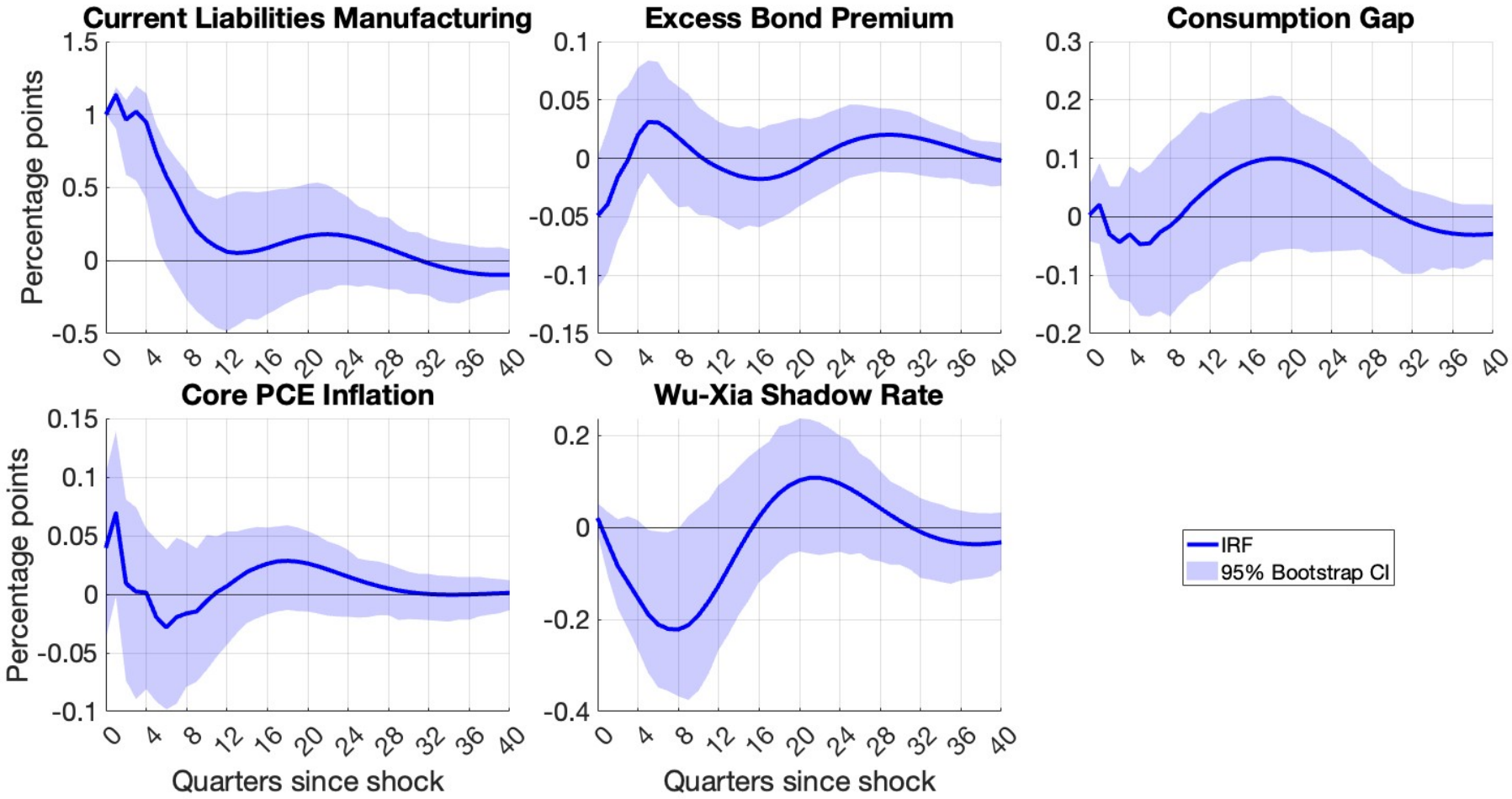


Figure 18: Impulse responses to the current liabilities of manufacturing sector shock (as a share of total liablities), 95% Bootstrap CI.

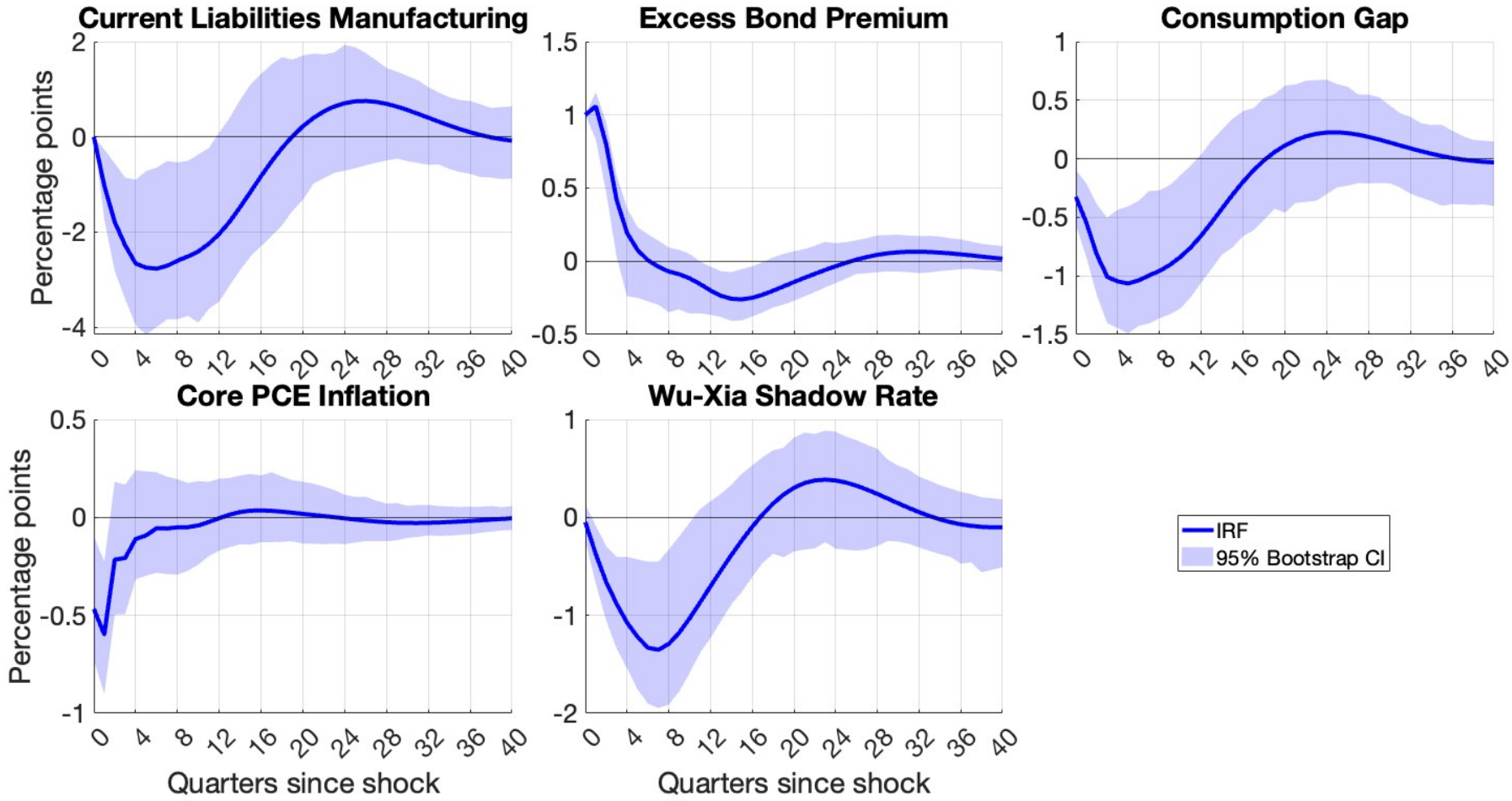


Figure 19: Impulse responses to the the excess bond premium shock, 95% Bootstrap CI.