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## US monetary policy is more powerful in low economic growth regimes

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## Working Paper Series

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US monetary policy is more powerful  
in low economic growth regimes

No 2919

## Abstract

We use nonlinear empirical methods to uncover non-linearities in the propagation of monetary policy shocks. We find that the transmission on output, goods prices and asset prices is stronger in a low growth regime, contrary to the findings of Tenreyro and Thwaites (2016). The impact is stronger on private investment and durables and milder on the consumption of nondurable goods and services. In periods of low growth, a contractionary monetary policy implies lower expected Treasury rates and higher premia along the entire Treasury yield curve. Similarly, the corporate excess bond premium rises and the stock market drops substantially during recessions. We use the monetary policy surprises and their predictors provided by Bauer and Swanson (2023a), and identify an additional predictor, the National Financial Condition Index (NFCI), which is relevant in the nonlinear setting. A Threshold VAR, a Smooth-Transition VAR and nonlinear local projection methods all corroborate the findings.

**Keywords:** Asset prices, business cycles, monetary policy, non-linearities, TVAR, STVAR, local projections **JEL Classification:** C32, E32

## Non-technical summary

Is the transmission mechanism of the monetary policy on asset prices, output and goods prices state dependent? Specifically, is there a differential effect of monetary policy depending upon the position of the state in the business cycle? Should the policymaker force the economy to expand below a certain threshold to be effective? We investigate whether the efficacy of monetary policy depends upon the state of the real economy using a new available set of monetary policy surprises computed by Bauer and Swanson (2023a) through high-frequency identification; that is, focusing on interest rate surprises in a narrow window of time around Federal Reserve's Federal Open Market Committee (FOMC) announcements.

First, we remove the components of the monetary policy surprises that is predictable with publicly available macroeconomic and financial market information in the nonlinear setting. Then, we estimate a Bayesian Threshold Vector Autoregression (TVAR) model and a Bayesian Smooth Transition Vector Autoregression (STVAR) model to uncover nonlinearities in the propagation of monetary policy shocks. Specifically, we investigate the state dependence of monetary policy impulse response functions on a range of real, nominal and asset price variables.

We show that monetary policy shocks are more powerful in the low growth regime, with the negative response on output accompanied by a negative response on goods prices in a textbook fashion; in the high growth regime, instead, the response of goods prices is muted. Essentially, the policy makers cannot steer goods prices using traditional tools when the economy is growing above the underlying real trend, but can still affect output growth. Monetary policy is effective on goods prices when the real economy is weak and growing below the underlying trend. This suggests that the FOMC should reduce output growth below a certain threshold, before an additional monetary policy shock is effective on goods prices. Across sub-components of real GDP, the impact is larger on private investment and durable goods than on nondurables and services consumption, consistent with the fact that spending on durables can be postponed in response to adverse shocks.

These findings are relevant for the design of stabilization policies. If changes in the monetary policy rates have limited impact on inflation in periods of expansions, central banks may be required to cause a recession to be able to steer the inflation rate towards the

desired target.

With regard to asset prices, the response of expected Treasuries is positive in expansions but negative in recessions along the entire yield curve. In periods of low economic growth, a contractionary monetary policy is judged by markets as further conducive to lower nominal economic growth and therefore a decline in the short term interest rates is expected by investors in the medium- to the long-term. The response of the term premium is instead countercyclical along the entire yield curve. It is strongly positive in the low growth regime and mildly negative in the high growth regime. The impact on the term premia is largest for Treasuries at longer maturity, after 16 month horizon and in the low growth environment. Similarly, the component of corporate credit spreads not associated to credit risk (e.g. the excess bond premium of Gilchrist and Zakrajšek (2012)) rises immediately in response to monetary policy shocks in the low growth regime, while it is muted in the high growth regime. All in all, bond holders demand higher premia for holding sovereign and corporate bonds in recessions. As for stock prices, they decline in response to a monetary policy shock and the drop is twice as deeper during recessions.

These results are broadly consistent with theoretical models assuming information asymmetries in financial markets. During recessions, asset prices drop, business contracts, the net worth of agents declines, the external finance premium rises and finance constraints are more likely to bind. The financial accelerator amplifies the effects of a tightening in monetary policy.

# I Introduction

Is the transmission mechanism of the monetary policy on asset prices, output and goods prices state dependent? Specifically, is there a differential effect of monetary policy depending upon the position of the state in the business cycle? Should the policymaker force the economy to expand below a certain threshold to be effective? We investigate whether the efficacy of monetary policy depends upon the state of the real economy using a new available set of monetary policy surprises computed by Bauer and Swanson (2023a) through high-frequency identification; that is, focusing on interest rate surprises in a narrow window of time around Federal Reserve’s Federal Open Market Committee (FOMC) announcements.

As suggested by recent studies (Cieslak, 2018; Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023a,b), monetary policy surprises are predictable with publicly available macroeconomic and financial market information.<sup>1</sup> Therefore, we remove the components of the monetary policy surprises that is predictable following the recommendations in Bauer and Swanson (2023a), but in a nonlinear setting. In particular, we regress the monetary policy surprises on the economic and financial variables that predate the FOMC announcements, allowing for possible nonlinearities in the relations, and take the residuals. In addition to the six variables suggested by Bauer and Swanson (2023a), we also propose a seventh variable, the National Financial Condition Index (NFCI), available in real time (Amburgey and McCracken, 2023) and relevant in the nonlinear setting. The NFCI was identified by (Adrian et al., 2019) as a strong explanatory variable of the level of risk in economic activity.

Equipped with the unpredictable monetary policy surprises derived from the nonlinear regression, we investigate the state dependence of monetary policy impulse response functions on a range of real, nominal and asset price variables. We show that monetary policy shocks are more powerful in the low growth regime, with the negative response on output accompanied by a negative response on goods prices in a textbook fashion; in the high growth regime, instead, the response of goods prices is muted. Essentially, the policy makers cannot

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<sup>1</sup>In Bauer and Swanson (2023b), the Fed’s responsiveness to the economy is both time-varying and unobserved by the private sector, and the private sector must form beliefs about this parameter. Therefore, monetary policy surprises are due not only to exogenous monetary policy shocks, but also to imperfect information about the Fed’s response parameter. As a consequence, monetary policy surprises can be correlated with economic variables observed prior to the policy announcements (i.e. “Fed response to news” channel). A precondition for this effect is that the public systematically underestimated how strongly the Fed would respond to economic news. Recent evidence supports this view (e.g. Schmeling et al., 2022).

steer goods prices using traditional tools when the economy is growing above the underlying real trend, but can still affect output. Monetary policy is effective on goods prices when the real economy is weak and growing below the underlying trend. This suggests that the FOMC should reduce output growth below a certain threshold, before an additional monetary policy shock is effective on goods prices. Across sub-components of real GDP, the impact is larger on private investment and durable goods than on nondurables and services consumption, consistent with the fact that spending on durables can be postponed in response to adverse shocks (Browning and Crossley, 2009).

With regard to asset prices, the response of the expected component of Treasuries is positive in expansions but negative in recessions along the entire yield curve. In periods of low economic growth, a contractionary monetary policy is judged by markets as further conducive to lower nominal economic growth and therefore a decline in the short term interest rates is expected by investors in the medium- to the long-term. The response of the term premium component is instead countercyclical along the entire yield curve. It is strongly positive in the low growth regime and mildly negative in the high growth regime. The impact on the term premia is largest for Treasuries at longer maturity, after 16 month horizon and in the low growth environment. Similarly, the component of corporate credit spreads not associated to credit risk (i.e. the excess bond premium of Gilchrist and Zakrajšek (2012)) rises immediately in response to monetary policy shocks in the low growth regime, while it is muted in the high growth regime. All in all, bond holders demand higher premia for holding sovereign and corporate bonds in low growth regimes. As for stock prices, they decline in response to a monetary policy shock and the drop is twice as deep during recessions.

The empirical literature on how monetary policy transmits in different phases of the business cycle is rather small. To our knowledge, the most recent published paper for the US is the study by Tenreyro and Thwaites (2016), who estimate the responses using the local projection model of Jordà (2005), combined with the smooth-transition regression method of Granger and Teräsvirta (1993) and with shocks identified in the manner of Romer and Romer (2004).<sup>2</sup> Contrary to our findings, their results indicate that the effects of US monetary policy

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<sup>2</sup>Angrist et al. (2018) evaluate the effect of monetary policy contractions and expansions on macroeconomic outcomes proposing a propensity score method. They find that target rate increases reduce employment and industrial output, and somewhat less successfully, inflation. At the same time, target decreases appear to have little stimulative effect on output or inflation. We instead condition the effects of the monetary



shocks are less powerful in recession. In this state of the economy, surprisingly, they found that a contractionary monetary policy shock even increases aggregate output, real business investment and real consumption in the first year. Moreover, the negative permanent effect on goods price inflation in the linear model, during expansions and during recessions, is striking. That would imply that goods prices unceasingly decline. Instead, theory suggests that after a drop in goods prices, they stabilise to a lower level, which implies that the response of inflation should revert to zero in the long run. Our findings are more in line with textbook economic principles.

There is one main issue that undermines the reliability of Tenreyro and Thwaites (2016)'s evidence. The method that Romer and Romer (2004) employs to identify the monetary policy shocks used by Tenreyro and Thwaites (2016) consists of running a regression of the change in the policy rate on central bank's forecasts, motivated by an empirical Taylor rule. Importantly, Miranda-Agrippino and Ricco (2021) show that such residuals are not exogenous. They are autocorrelated and can be predicted using past information.<sup>3</sup> We show that the difference between Tenreyro and Thwaites (2016)'s results and the findings in our paper is primarily driven by the approach used to identify the monetary policy shocks. A key feature of a VAR is that it includes the lagged value of all regressors, which allows to control for autocorrelation in monetary policy surprises and past information intrinsic in all regressors. Interestingly, we detected that by employing the database of Tenreyro and Thwaites (2016), including the Romer and Romer (2004)'s monetary policy shocks, a Smooth Transition Vector Autoregressive model (STVAR), which closely resembles their model in a VAR context, cannot replicate their results. The same conclusions can be drawn by using a Threshold Vector Autoregression (TVAR), after having consistently re-estimated the Romer and Romer (2004)'s monetary policy surprises. Instead, we can replicate qualitatively our findings with our estimated nonlinear monetary policy shocks using both TVAR and STVAR models as well as smooth-transition and threshold local projection methods.<sup>4</sup>

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policy on the state of the business cycle.

<sup>3</sup>Using quantile regressions, Mumtaz and Surico (2015) study the relationship between real consumption growth and real interest rates, instrumented with Romer and Romer (2004)'s monetary policy shocks. They also find that the effects are stronger when consumption growth is above its conditional average.

<sup>4</sup>It is preferable to extract the response functions from nonlinear VARs rather nonlinear local projection methods. Gonçalves et al. (2024) show that, when the state of the economy is endogenous, the local projections' estimator of the response functions, applied by Tenreyro and Thwaites (2016), tends to be asymptotically biased except for the impact response. A similar critique applies to the study of Alpanda



A recent working paper by Bruns and Piffer (2021), which makes use of a STVAR, finds similar results to ours. However, their identification of the monetary policy shocks is obtained by using as an external instrument the high frequency monetary policy surprises of Gertler and Karadi (2015) and Jarociński and Karadi (2020), which are autocorrelated (see Miranda-Agrippino and Ricco, 2021). Moreover, their identification is obtained including sign restrictions at impact on all financial variables and sign restrictions within six months on real GDP and goods prices. Instead, the response of all variables in our model is always left unrestricted. This allows us to be completely agnostic about the impact of the monetary policy shocks on asset prices and the macroeconomy.<sup>5</sup>

Several other earlier studies make use of old-fashioned approaches to identify monetary policy shocks, often based on assumptions that are difficult to maintain in light of more recent advancements in the empirical and theoretical literature. Weise (1999) identifies money supply shocks using the monetary aggregate M1 through a Choleski orthogonalization ordering money last. The regime is indicated by the first lag of quarterly GDP growth and therefore subject to high-frequency shifts. The results are difficult to interpret because expansionary monetary shocks are contractionary in a high-growth regime. In other words, monetary policy decisions would achieve exactly the opposite of the desired results. Thoma (1994) also estimates a nonlinear VAR in output and monetary variables. Focusing on the VAR coefficients, he finds that the relationship between M1 and output becomes stronger, when real activity experience a cyclical decline, while it becomes weaker, when an upswing in real activity occurs. In general, monetary aggregates have lost their appeal in signaling monetary policy surprises.

Garcia and Schaller (2002) studies the response of industrial production growth to monetary policy in a VAR for the United States from 1955:II to 1993:I with a two-state Markov switching regime. They find that interest rate changes or monetary policy shocks identified using a Choleski orthogonalization have a larger negative impact on output in recessions

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et al. (2021), who use local projections in a panel setting with 18 advanced economies and monetary policy shocks identified using sign restrictions on interest rates, output and prices. Francis et al. (2023) corroborate the results that nonlinear local projections are biased and, in addition, find that state-dependent IRFs from STVARs are typically more biased than those obtained from a TVAR, given the larger number of parameters to estimate.

<sup>5</sup>A recent BIS paper provides a view of the inflation characterised by a low and a high inflation regime (Borio et al., 2023). The study documents the stylised facts describing the two regimes and the transitions between them based on disaggregated price dynamics and the joint behaviour of wages and prices.

than expansions. However, it is assumed that the persistence of GDP is the same in booms and recessions, which can be an overly restrictive assumption (Acemoglu and Scott, 1997).

Peersman and Smets (2001) and Lo and Piger (2005) also study the asymmetric response of industrial production growth to monetary policy respectively in seven Euro-area countries and US, using a traditional recursive VAR where the policy variable is ordered after output, and find that the impact is more negative during recessions than boom.

Chen (2007) studies the impact on stock returns with a two-state Markov switching regime and finds that monetary policy has larger effects on stock returns in bear markets. However, monetary policy is measured by interest rate instruments, which are endogenous, and innovations to the Fed fund rate estimated from a linear recursive VAR model, thereby generating an inconsistency given the nonlinear framework.<sup>6</sup>

Burgard et al. (2019) estimate a logit mixture VAR model to assess the effects of monetary policy shocks in the euro area with monetary policy shocks identified using policy rates and Choleski identification. They show that monetary policy transmission in the euro area can be described as a mixture of two states. In both states, output and prices decrease after monetary policy shocks. During crisis times, the contraction is much stronger, as the peak effect of both variables is roughly one-and-a-half times as large when compared to normal times. The key issue is that this type of models requires the estimation of time-varying state weights and, therefore, the model should be very parsimonious. Burgard et al. (2019) estimate their four variable nonlinear VAR with one lag only. In addition, the states depend on the included variables and requires an interpretation.

All in all, Weise (1999), Garcia and Schaller (2002), Lo and Piger (2005), Burgard et al. (2019) and Bruns and Piffer (2021) find that US monetary policy is more effective during recessions than during expansions, while Thoma (1994) and Tenreyro and Thwaites (2016) find the opposite results as in their papers monetary policy in the United States is less powerful during recessions. However, all these studies can be challenged because of the method used to identify the monetary policy shocks.

We revisit the analysis (i) by employing a Bayesian TVAR and a Bayesian STVAR, which

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<sup>6</sup>Perez-Quiros and Timmermann (2000) study the asymmetric effects of short-term interest rates and monetary growth on returns of size-sorted decile portfolios across recession and expansion states. However, they do not identify monetary policy shocks.

define the states of interests and allow to compute state-dependent impulse response functions, as well as nonlinear local projection methods, and (ii) by using the recently computed monetary policy surprises by Bauer and Swanson (2023a) through the high-frequency identification, orthogonalized exploiting macroeconomic and financial data that are publicly available prior to the monetary policy announcement. Over the past two decades, high-frequency identification of monetary policy surprises has become an important tool for identifying the effects of monetary policy in a linear framework on asset prices (Kuttner, 2001; Bernanke and Kuttner, 2005; Gürkaynak et al., 2005; Hanson and Stein, 2015; Altavilla et al., 2019; Jarociński and Karadi, 2020; Swanson, 2021; Bauer and Swanson, 2023a,b) and the macroeconomy (Cochrane and Piazzesi, 2002; Faust et al., 2003, 2004; Gertler and Karadi, 2015; Ramey, 2016; Stock and Watson, 2018; Jarociński and Karadi, 2020; Bauer and Swanson, 2023b).

The paper is structured as follows. Section II presents the model. Section III describes the method and the variables used to identify the monetary policy shocks. Section IV discusses the key results. Section V explains the reasons behind the differences with Tenreiro and Thwaites (2016). Section VI provides a number of robustness checks. Section VII concludes.

## II Framework

### II.A Model specification

Our baseline model is a TVAR. The results obtained with the STVAR and the nonlinear local projections are provided in Section VI. The reduced form TVAR takes the following form

$$\mathbf{X}_t = (\mathbf{c}_l + \mathbf{\Pi}_l(\mathbf{L})\mathbf{X}_{t-1})F_t\{z_{t-1} < z^*\} + (\mathbf{c}_h + \mathbf{\Pi}_h(\mathbf{L})\mathbf{X}_{t-1})(1 - F_t)\{z_{t-1} \geq z^*\} + \mathbf{u}_t^T, \quad (1)$$

$$\mathbf{u}_t^T \sim N(0, \mathbf{\Omega}_t), \quad (2)$$

$$\mathbf{\Omega}_t = \mathbf{\Omega}_l(F_t)\{z_{t-1} < z^*\} + \mathbf{\Omega}_h(1 - F_t)\{z_{t-1} \geq z^*\}, \quad (3)$$

where  $\mathbf{u}_t^T$  denotes the  $n \times 1$  vector of reduced form residuals,  $\mathbf{\Omega}_t$  the state-contingent covariance matrix of the residuals,  $z_t$  the state variable,  $z^* = P_k(z_t)$  the  $k$ th percentile of  $z_t$ ,  $\mathbf{c}_l$  and  $\mathbf{c}_h$  the vector of intercepts in the two regimes (low ( $l$ ) and high ( $h$ ) growth,  $S \in \{l, h\}$ ) and  $\mathbf{\Pi}_l$  and  $\mathbf{\Pi}_h$  the lag polynomials. The regime switches are governed by the indicator function  $F_t \in \{0, 1\}$  and are indexed by  $t - 1$  to avoid endogeneity problems. A regime shift occurs only if  $z_t$  remains in the new regime for at least two periods. By doing so, we avoid high frequency shifts when  $z_t$  is close to  $z^*$ .

The vector  $\mathbf{X}_t = [u_t^{T-mp}, r_t, i_t^{en}, i_t^{tpn}, b_t, e_t, y_t, p_t]'$  defines the endogenous variables of the baseline model, where  $u_t^{T-mp}$  denotes the orthogonalized monetary policy surprises,  $r_t$  the policy shadow rate,  $i_t^{en}$  and  $i_t^{tpn}$  are the expected Treasury rate and the correspondent term premium at maturity  $n$ , respectively,  $b_t$  the corporate excess bond premium,  $e_t$  the stock market price,  $y_t$  output and  $p_t$  goods prices. The variables  $e_t$ ,  $y_t$  and  $p_t$  are defined in log levels. By construction, the 10-year Treasury rate  $i_t^n = i_t^{en} + i_t^{tpn}$ . Therefore, we can recover the impulse response functions (IRFs) for the nominal Treasury yield.

We set the lag order  $p$  to 12. This helps estimating a trend-stationary process. As for linear VARs, the stability condition of a TVAR requires that all the roots,  $r$ , of  $\mathbf{\Pi}_l$  and  $\mathbf{\Pi}_h$

$$|\mathbf{\Pi}_S(r)| = |\mathbf{I}_n - \mathbf{\Pi}_{S,1}r - \mathbf{\Pi}_{S,2}r^2 - \dots - \mathbf{\Pi}_{S,p}r^p| = 0, \quad S \in (l, h),$$

lie outside the unit circle. This guarantees that the system of equations is stationary.

The state variable,  $z_t$ , is assumed to depend on current and past month-on-month output growth using an exponentially weighted moving average (EWMA), which gives larger weights,  $\alpha$ , to the most recent observations and geometrically declining weights to past growth rates,  $z_t = \sum_{i=0}^{\infty} \alpha(1 - \alpha)^i (y_{t-i} - y_{t-1-i})$ . Hence,  $z_t$  is a function of the entire history of  $y_t$  and can be written as:

$$z_t = \alpha(y_t - y_{t-1}) + (1 - \alpha)z_{t-1}, \quad \alpha \in (0, 1). \quad (4)$$

To construct the structural impulse response functions, the feedback from future changes of  $z_{t-1}$  into the dynamics of macroeconomic system ought to be taken into account.

The reduced form TVAR is estimated in a Bayesian framework using a multivariate version of the sampler developed in Chen and Lee (1995), which allows to draw from the

posterior of the model parameters. For the parameters of both regimes, we assume natural conjugate Normal-Inverse-Wishart (N-IW) priors. The IW priors for  $\Omega_l$  and  $\Omega_h$  have  $n + 2$  degrees of freedom and diagonal scale matrix with the  $i$ -th diagonal elements equal to the mean squared error from estimating an AR(1) for the  $i$ -th variable. Conditional on  $\Omega_l$  and  $\Omega_h$ , the priors for  $\Pi_l$  and  $\Pi_h$  are Normal with Minnesota-type mean and variance (Doan et al., 1984), and complemented with a dummy-initial observation prior (Sims, 1993) that is consistent with the assumption of cointegration.<sup>7</sup> The sample spans over the monthly period from February 1988 to December 2019 owing to the availability of the monetary policy surprises.

As for the data, the interpolation of GDP to a monthly frequency using the Chow and Lin (1971)’s method employs industrial production and real retail sales, while the GDP deflator is interpolated using the consumer price index and the producer price index; thereby, both variables are coincident indicators including supply and demand considerations. We show the results with the consumer price index replacing the GDP deflator in the robustness section. The policy shadow rate is equal to the Federal Fund Rate until May 2009 and to the shadow rate provided by Wu and Xia (2016) for the rest of the sample. The excess bond premium is provided by Gilchrist and Zakrajšek (2012). The expected yields and the term premia are obtained from Adrian et al. (2013). All the other data underlying the vector  $\mathbf{X}_t$  are obtained from the Federal Reserve Economic Data (FRED) online database and Haver analytics.

## II.B The state variable

In several studies, the state variable is computed using a moving average of the last months of the variable of interest (e.g. Tenreyro and Thwaites, 2016; Ramey and Zubairy, 2018; Knotek and Zaman, 2021). This approach tends by construction to postpone the potential change in regime, if the shock is not relatively large. The solution proposed by others is to take a centered moving average, between  $t - k$  and  $t + k$  (e.g. Auerbach and Gorodnichenko, 2012; Ascari and Haber, 2022). However, this provides inconsistent estimates, because the state variable ought to be predetermined, so that it is uncorrelated with the shock happening

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<sup>7</sup>The hyperparameters take standard values from the literature. The hyperparameter, which determines the tightness of the Minnesota prior, is set equal to 0.2. The parameter, which governs the variance decay with increasing lag order, is set equal to 2. The hyperparameter, which determines the tightness of the “dummy-initial-observation” prior is set equal to 1, a value recommended by Sims and Zha (1998).

at time  $t$  or in future periods.

Therefore, we construct the state variable,  $z_t$ , using Equation 4 and setting  $\alpha = 0.125$ . This attributes a relatively larger weight to the most recent observations and allows to better capture the timing of the starts and the end of a recession as defined by the NBER (see Figure 1). Specifically,  $z_t$  is calculated using month-on-month output growth starting from February 1959, as a number of observations are required to compute the underlying output growth. The state variable is highly correlated with year-on-year real GDP growth (94%) and its median which defines the threshold for the regime switch is 2.53% in annualised terms over the sample period 1988-2019. To put this figure in perspective, the median of year-on-year real GDP growth using quarterly data is 2.56%.

As an alternative benchmark, we estimate the threshold by maximizing the marginal likelihood through a grid search. The percentile is set at 0.25, which suggests setting the threshold for underlying real GDP growth at 1.83% annualised. We show in the robustness section that the results are the same using either of the thresholds.

## II.C Nonlinear Structural Impulse Responses

Structural shocks,  $\epsilon_t$ , may have nonlinear effects on  $\mathbf{X}_t$ . Responses then depend on the history of the data and on the sign and magnitude of the structural shocks with effects from  $t$  to  $t + k$ .  $z_{t-1}$  is a function of  $y_{t-1}$  and, therefore,  $z_t, z_{t+1}, \dots, z_{t+k-1}$  are endogenously determined in the TVAR. To construct the structural response functions, the feedback from future changes of the state variable into the dynamics of the macroeconomic system is taken into account.

Following Balke (2000) and Koop et al. (1996),<sup>8</sup> who proposed the construction of the response functions through the conditional expectations, we compute the nonlinear structural IRFs,  $IRF_{\mathbf{S}}^{\mathbf{X}}(\epsilon_{S,t}, \Gamma_{t-1}(z_{t-1}))$ , as the difference between the expectations of the realizations  $\mathbf{X}_{S,t+k}$  at horizon  $k$ , conditional on  $\epsilon_t$  and the information set at time  $t - 1$ ,  $\Gamma_{t-1}$ , and the

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<sup>8</sup>Koop et al. (1996) were not concerned about structural identification, they used the reduced form residuals. Given that we focus on structural identification, the algorithm differs from Koop et al.'s approach. See also Kilian and Lütkepohl (2017, Chapter 18) for a discussion of state dependent IRFs.

expectations of the realizations  $\mathbf{X}_{S,t+k}$  conditioned only on  $\Gamma_{t-1}$ :

$$IRF_{\mathbf{S}}^{\mathbf{X}}(\epsilon_{S,t}, \Gamma_{t-1}(z_{t-1})) \equiv \mathbb{E}(\mathbf{X}_{S,t+k} | (\Gamma_{t-1}(z_{t-1}), \epsilon_{S,t})) - \mathbb{E}(\mathbf{X}_{S,t+k} | \Gamma_{t-1}(z_{t-1})), \quad (5)$$

where  $S \in \{l, h\}$  indicates whether the economy is in the low- or high-growth regime at time  $t + k - 1$ . The conditional expectations are calculated by simulating forward the model.

It is worth emphasizing that the switch among regimes is treated as endogenous, as the economy can shift from low to high growth regimes or *viceversa* over the simulation horizon, depending on the sign, the size of the shock, the estimated parameters and the specific history of the system prior to the shock. The starting points are assumed to be the mean of all the in-sample observations in each regime, in order to obtain the most representative picture of the dynamics associated to each regime.

### III Identification of the Monetary Policy Shocks

We use the high-frequency monetary policy surprises provided by Bauer and Swanson (2023a) based on the work of Swanson and Jayawickrema (2021). The surprises are estimated following the approach of Gürkaynak et al. (2005) and Nakamura and Steinsson (2018) computing the changes in quarterly Eurodollar future contracts over a 30-minute window starting 10 minutes before each FOMC announcement and ending 20 minutes afterwards. Following Nakamura and Steinsson (2018), Bauer and Swanson (2023a,b) take the first principal component of the changes in the first four quarterly Eurodollar future contracts. As pointed out by the authors, taking the first principal component is reminiscent of taking a weighted average of the target and path factors, thereby parsimoniously capturing the main features of both conventional and forward guidance monetary policy surprises. The dataset covers the period from 1988 to 2019 and includes 322 FOMC announcements.

Event study regressions are useful to assess the impact of monetary policy shocks on asset prices, with the underlying surprises measured over tight windows around the policy announcement. With intradaily data and the usual 30-minute announcement windows, plus the fact that FOMC decisions are taken few hours before the actual announcement, the likelihood that monetary policy surprises are exogenous to contemporaneous movements



in asset prices is arguably high. Yet, such regressions cannot provide information about the subsequent asset price dynamics and persistence. Furthermore, event study regressions cannot be used to assess the macroeconomic implications.

To study the transmission mechanism of monetary policy shocks on asset prices, output and goods prices, we rely on a Bayesian VAR. At the same time, by making use of Bauer and Swanson (2023a) surprises, we avoid the difficulties associated with structural identification and narrative-based efforts. However, we have to make sure that these surprises are in fact monetary policy shocks; that is, there are no omitted variables that are correlated with the monetary policy surprises and independently affect the endogenous variables under investigation. Therefore, we remove the components of the monetary policy surprises that is predictable following the recommendations in Bauer and Swanson (2023a), using data publicly available in real time, but in a nonlinear setting.

### III.A Monetary Policy Surprises: Nonlinear Regression

We regress the monetary policy surprises on the economic and financial variables that predate the FOMC announcements and have predictive power for them as well as on the same set of variables interacted with a dummy defining the low and high economic growth regimes, and then take the residuals.<sup>9</sup>

The regressions take the following form:

$$mps_t = \alpha_0 + \alpha_1 \times F_t + \alpha'_2 \mathbf{X}_{t-} + \alpha'_3 \mathbf{X}_{t-} \times F_t + u_t^{T-mp}, \quad (6)$$

where  $mps_t$  are the monetary policy surprises in the 30-minute announcement window,  $\mathbf{X}_{t-}$  the vector of economic news observed prior to the FOMC announcement at time  $t$ ,  $F_t$  is the indicator function equal to unity if the underlying GDP growth is above the threshold,

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<sup>9</sup>Miranda-Agrippino and Ricco (2021) argue that high frequency surprises are likely to combine the monetary policy shocks with information about the state of the economy disclosed through the policy action and, therefore, they orthogonalize these surprises with respects to the FED's internal "Greenbook" forecast. Bauer and Swanson (2023a,b) show that Blue Chip forecasts have predictive power for monetary policy surprises that is just as strong as the predictive power of the Fed's "Greenbook" forecasts. This implies that the Fed is unlikely to have significantly private information, and that Fed information effects may not be an important source of that predictability. The approach suggested by Bauer and Swanson (2023a,b) is to employ data for the orthogonalization publicly available in real time, such as the employment report, which is a strong predictor of the Blue Chip forecast revision. Instead, the Fed's "Greenbook" forecasts is publicly available only five years after the FOMC meeting.

$F_t\{z_{t-1} \geq z^*\} = 1$ , and zero otherwise,  $F_t\{z_{t-1} < z^*\} = 0$ , and  $u_t^{T-mp}$  are the residuals of the monetary policy surprises. Predictors, which are observed prior to the FOMC announcement, are those suggested by Bauer and Swanson (2023a), which have an intuitive relationship to the Fed' monetary policy rule: (i) the surprise component of the most recent nonfarm payrolls release, (ii) employment growth over the last year, (iii) the log change in the Standard & Poor's 500 index (S&P 500) from 3 months (65 trading days) before the FOMC announcement to the day before the FOMC announcement, (iv) the change in the yield curve slope over the same period, (v) the log change in a commodity price index over the same period and (iv) the option-implied skewness of the 10-year Treasury yield, averaged over the proceeding month, from Bauer and Chernov (2021).

Column 1 of Table 1 replicates exactly the results of Bauer and Swanson (2023a) in a linear setting. News about nonfarm payrolls, employment growth, commodity prices and skewness of the 10-year Treasury yield are statistically significant at the 5% level predicting a hawkish monetary policy surprise and explain 16.2% of the variation of the monetary policy surprises. Column 3 adds the interaction terms. The prediction of the monetary policy surprises is independent of the state of the economy, as the interaction terms are not statistically significant based on the Wald test performed jointly on all  $\hat{\alpha}'_3 = 0$ .

In addition to the six variables proposed by Bauer and Swanson (2023a), we also find that the NFCI, proposed by Adrian et al. (2019) to assess the risk around the real GDP growth forecasts, and provided in real time by Amburgey and McCracken (2023), is a good predictor of the monetary policy surprises in the nonlinear model. Specifically, any change in the NFCI from two up to eleven weeks before the FOMC announcement to the day before the FOMC announcement is strongly statistically significant. We select for the analysis the change corresponding to the period with the largest t-statistics, that is seven weeks. The results for the regressions including the NFCI as a predictor are provided in column 2 for the linear model, where the coefficient is not statistically significant, and in column 4 for the nonlinear model, where the coefficients for both the low and high growth regime are statistically significant at the 1% level with both t-statistics above 3, but with opposite sign. Because the NFCI is a generated regressor, we report in parentheses bootstrapped t-stat using 50,000 bootstrap replications.

Increasing risk, tighter credit conditions and declining leverage are consistent with increases in the NFCI. The FED seems to respond to the dynamics in the NFCI more strongly than expected by market participants in the low growth regime, as tighter financial conditions predict a hawkish monetary policy surprise. In the high growth regime, the opposite occurs. Investors expect relatively high policy rates after a tightening in credit condition, whereas the FOMC delivers a lower increase in rates, as  $\hat{\alpha}_{2,NFCI} + \hat{\alpha}_{3,NFCI} < 0$  and the Wald test performed on the coefficients' sum set equal to zero,  $\hat{\alpha}_{2,NFCI} + \hat{\alpha}_{3,NFCI} = 0$ , has a  $\tilde{\chi}(1)^2 = 5.99$  with the correspondent P-value equal to 0.014.

Moreover, the results indicate that market based factors are better predictors within the nonlinear setting and by controlling for the NFCI, as the change in the stock market price and the change in slope of the yield curve, with the intuitive positive and negative signs respectively, become strongly statistically significant, and together with the change in commodity prices, have a larger  $\hat{\alpha}_2$ . More specifically, when the yield curve becomes more upward-sloping (i.e., when short-term interest rates fall relative to long-term rates, as they do during monetary easing cycles), the Fed is likely to follow with an easing surprise. Interestingly, once controlling for the NFCI, the nonfarm payrolls is a weaker predictor of the monetary policy surprises. Overall, the  $R^2$  of the nonlinear model including the NFCI (22.3%) increases by about five percentage points relative to the same model that does not control for the NFCI (17.1%).

The high-frequency residuals  $u_t^{mp}$  are converted to a monthly series by summing over all of the figures within the same month and set them to zero in absence of FOMC announcements (see Figure 2). Equivalent aggregation methods are adopted in the literature (e.g. Stock and Watson, 2012; Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023a).

### III.B Additional Issues

It could be argued that most high-frequency policy surprises occur around FOMC announcements during low growth regimes (as suggested by Cieslak (2018) and Schmeling et al. (2022)). In other words, it could be argued that the size of the shocks differs across regimes; partly affecting some of the results.  $mps_t$  has a standard deviation equal to 5.7 basis points in the low growth regime and 4.9 basis points in the high growth regime. However,  $u_t^{mp}$  has a stan-

dard deviation equal to 4.6 basis points in the low growth regime and 4.4 basis points in the high growth regime. We thus use a new measure of monetary policy surprises that is both more relevant and more exogenous than those used by previous researchers and investigate whether monetary policy is more effective in recessions or expansions.

Once a shock is available, the typical approach is to use it as an external instrument on the residuals of a VAR (Gertler and Karadi, 2015; Stock and Watson, 2012, 2018; Caldara and Herbst, 2019; Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023b). In our nonlinear setting, this would require estimating two different VARs with an external instrument depending upon the regime. A TVAR achieves this objective with the advantage of removing any remaining potential autocorrelation structures in the residuals of the monetary policy surprises (e.g. Miranda-Agrippino and Ricco, 2021). By so doing, we account for the slow absorption of information by the agents characterising models of imperfect information (e.g. Coibion and Gorodnichenko, 2015).

An alternative approach consists of studying the transmission of the monetary policy shocks using local projections (e.g. Ramey, 2016; Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023a). In our context, the nonlinear IRFs depend on the future state of the economy, which could be steered by the policymaker, which we can trace with a TVAR. Gonçalves et al. (2024) show that, when the state of the economy is endogenous, the local projections' estimator of the response function tends to be asymptotically biased except for the impact response. Francis et al. (2023) draw a very similar conclusion. Moreover, in a linear setting, Plagborg-Møller and Wolf (2021) prove that local projections and VARs estimate the same population impulse responses. The authors also recommend including the instrument in the VAR, ordering it first, and using a recursive ordering to estimate its effects, which makes results robust to non-invertibility of the structural shock of interest. Following this suggestion, we estimate a TVAR including the monthly announcement-frequency residual series aggregated from  $u_t^{T-mp}$ , order it first and investigate the potential heterogeneous transmission mechanism of the monetary policy shocks. Yet, we also show the results using nonlinear local projections.

## IV The Impact of Monetary Policy Shocks

### IV.A Nonlinear Transmission Mechanism of Monetary Policy

We start the analysis with the 8-variable monthly VAR,  $\mathbf{X}_t = [u_t^{T-mp}, r_t, i_t^{en}, i_t^{tpn}, b_t, e_t, y_t, p_t]'$ , and by considering the effects of the monetary policy shocks in the linear model using the recursive identification with the residuals of the monetary policy surprises ordered first.<sup>10</sup> This allows to further clean the variable of interest from autocorrelation structures and makes results robust to non-invertibility of the monetary policy shock. The resulting IRFs of the linear model are displayed in Panel A of Figure 3. The blue line provides the median IRFs and the shaded bands around the median IRFs provide the corresponding posterior 68% credible sets. A monetary policy shock amounting to 25 basis points increases at impact the effective federal fund rate by 9 basis points and the excess bond premium by 14 basis points, and causes a drop in the stock market by 4%. Conversely, the impact on the expected component of the 10-year Treasury rate is rather volatile, while the credible set of the term premium is almost all in the positive territory after about two years. The monetary policy shocks reduce output reaching the trough at -0.4% after four years and goods prices at -0.2% after few months.

We now turn to the key research question of the paper: does the transmission of the monetary policy shocks to financial and macroeconomic variables differ across the low and high growth regimes? Panel B of Figure 3 provides an answer to this question. The solid blue line provides the median IRFs in the low growth regime and the shaded bands around the median IRFs provide the corresponding posterior 68% credible sets. The solid red line provides the median IRFs in the high growth regime and the dashed red lines around the median IRFs provide the corresponding posterior 68% credible sets.

The first point to note in Figure 3 is that, despite a similar response at impact amounting to about 5-10 basis points, the persistence of the effective federal fund rate is much lower in the low growth regime, returning back to steady state in few months, rather than two years and half in the high growth regime. The response of the interest rates is intuitive, because a contractionary monetary policy in a low growth environment could push the economy into

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<sup>10</sup>In the case of the linear model, the dummy in Equation (6) is set equal to zero.

a recession and, in such circumstances, interest rates are expected to decline.

The second key point is that a contractionary monetary policy exercised in the low growth regime causes an immediate increase in the excess bond premium of the corporate sector by 22 basis points and a large decline in the the stock market price by 5.2%. In the high growth regime, instead, the excess bond premium is not really affected and the stock market price declines by only 2.7%. In an environment characterised by low economic growth, the corporate sector is already relatively weak. Therefore, companies have to offer higher premia for investors to hold their assets in their portfolio in response to a contractionary monetary policy shock.

The third point is that output and goods prices are strongly affected in a low growth regime. A 25 basis point contractionary monetary policy implies a drop in GDP by 0.5% at through after about two years and in goods prices by 0.2% already after four months. In the high growth regime, about 85% of the credible set is below zero at two year horizon, while a large fraction of the credible sets of goods prices include zero. Contrary to the most recent conclusions reached by the literature provided by Tenreyro and Thwaites (2016), we find that monetary policy is more powerful in states of low economic growth. When the economy is already weak, the balance sheets of companies and households are rather fragile. In such an environment, the tightening of the financing conditions by the monetary authorities has stronger macroeconomic implications. We will show that these results are corroborated when using the quarterly frequency without interpolation or when substituting the GDP deflator with the Personal Consumption Expenditure (PCE) goods inflation. In the quarterly model, the response of GDP to monetary policy shocks is also more negative in the high growth regime relative to the monthly model.

The fourth point is the different response of the long-term expected yields and term premia. The expected yields decline after a contractionary monetary policy shock in a low growth regime, while they rise in a high growth regime (see Figure 3). The distribution of their IRFs in the high growth regime shifts downwards with the upper bound of the credible set declining very close to zero at 2-year horizon. Conversely, the term premia increase after a contractionary monetary policy shock in a low growth regime, while they decrease in a high growth regime. The response of the long term expected interest rates is intuitive, because a

contractionary monetary policy in a low growth environment could push the economy into a recession and, in such circumstances, the long term interest rates are expected to decline. Also the response of the term premia is intuitive, because they are countercyclical; they rise in a low growth environment, but remain unaffected if the economy is booming. The overall impact on nominal yields, which is obtained by summing the two IRFs for each draw (see Panel C of Figure 4), shows that the long-term interest rates rise in the high growth regime, but mean-revert in the low growth regime declining in the first few months.

We replicate the above exercise changing the maturity of the Treasury yield (see Figure 4). The results are essentially the same in both regimes when using the 5-year Treasury yield and qualitatively similar with shorter rates, with the shape of the IRFs being very similar across the yield curve. As for the term premium, it rises countercyclically in the low growth regime along the entire yield curve after about 6 months, reaching the peak after about 16 months, and the impact is larger the higher is the maturity of the Treasury.

A summary of the results is also provided in Table 2 at impact and 6-month horizon. Focusing on the median value, there is a clear hump-shape response in the high growth regime along the expected yield curve, which is rising together with the horizon with the largest value recorded at the 3-year maturity. Conversely, the 3-year expected yield already tends to decline at impact in the low growth regime. As for the countercyclical term premium, the impact is largest for Treasuries at longer maturity, at longer horizon and in the low growth environment. Also the corporate excess bond premium rises most immediately and in the low growth environment. Similarly, the stock market prices decline immediately and mean-revert especially in the low growth environment.

Following the analysis carried out by Tenreyro and Thwaites (2016), we also study the differential impact of the monetary policy shocks on the volume of fixed private business investment,  $y_t^f$ , the volume of durable goods and housing investment,  $y_t^d$ , and the consumption of nondurable goods and services in real terms,  $y_t^n$ . The Bayesian VAR,  $\mathbf{X}_t = [u_t^{T-mp}, r_t, y_t, p_t, y_t^f, y_t^d, y_t^n]'$ , uses quarterly data and in addition it includes real GDP, DGP deflator, the shadow rate and the orthogonalized monetary policy surprises, which are ordered first. The high-frequency  $u_t^{t-mp}$  are converted to a quarterly series by summing over all of the figures within the same quarter and set them to zero in absence of FOMC



announcements.<sup>11</sup> Figure 5 plots the impulse response of the volumes of the three expenditure aggregates to the same monetary policy shock. In line with the response of aggregate output, all the volume indices decline in low and high growth regimes. The effect on private investment and durable goods is three times the size of the impact on non-durables and services. A 25 basis point standard deviation monetary policy shock implies at through after half year a 1.4% drop in both the volume of fixed private business investment and the volume of durable goods and housing investment. The consumption of nondurable goods and services in real terms records a 0.4% drop at through after one year, which resembles the impact on aggregate real GDP. The impact, however, is more uncertain for the consumption of durable and housing investment volume in the low growth regime. The results across aggregates are consistent with the fact that spending on investment and durables can be postponed in response to adverse shocks, as existing stocks of investment and durable goods can still provide utility given their longer lifespans.

Our results are in line with the findings by Tenreyro and Thwaites (2016) in periods of expansions, while they are just the opposite in recession periods. Tenreyro and Thwaites (2016) find a positive response of all disaggregated volume components after a monetary policy contraction in recessions in the first year and in the case of fixed business investment over three years. Hence, their results are difficult to interpret in recession phases.

We also show the results on the other variables used in the quarterly model. The results are broadly similar to the monthly model for the shadow rate and real GDP and are the same for the GDP deflator. Focusing on the differences, the response of the shadow rate is less persistent and the response of real GDP is strong and clearly negative also in the high growth regime. Overall, the results of the quarterly model corroborate our findings that the monetary policymaker is able to bring down goods prices only when the real economy is already weak. Consistently with economic theory, the monetary multiplier on private investment and durable good is larger than the multiplier on the consumption of nondurable goods and services.

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<sup>11</sup>We construct the state variable,  $z_t$ , using Equation 4 and setting  $\alpha = 0.2$ . When using quarterly data, a relatively larger weight to the most recent observations allows to better capture the timing of the starts and the end of a recession as defined by the NBER. The median which defines the threshold for the regime switch is 2.56% in annualised terms in line with the monthly model.

## V On the Difference with Tenreyro and Thwaites (2016)

The aim of this section is to provide an explanation behind the opposite conclusions obtained by Tenreyro and Thwaites (2016). Two are the potential reasons. First, the differences could be due to the identification of the monetary policy surprises, high-frequency versus the Romer and Romer (2004)'s identification. To check if this is the case, we estimate the TVAR as well as a smooth-transition VAR (STVAR), the latter being even closer to the method of Tenreyro and Thwaites (2016), using the Romer-Romer monetary policy surprises and the overall database of Tenreyro and Thwaites (2016). If the results of Tenreyro and Thwaites (2016) do not hold, the findings will be suggestive that the use of Romer and Romer (2004)'s monetary policy surprises as exogenous monetary policy shocks is inappropriate. Second, the differences could be due to the different methodologies. Therefore, we also estimate the macroeconomic response of the monetary policy shocks following Tenreyro and Thwaites (2016) by using nonlinear local projection methods, but employing our high-frequency monetary policy surprises as exogenous shocks.

To make the results comparable, in this section only, we move away from our baseline model and adopt the strategy of Tenreyro and Thwaites (2016). The variables are (the log of) real GDP, PCE inflation, the Federal fund rate and the monetary policy shock. The model is estimated with two quarter lags and the transition to the other regime is not allowed. The latter is an important deficiency. In fact, Gonçalves et al. (2024) and Francis et al. (2023) show that, when the state of the economy is endogenous, the local projections' estimator of the response function tends to be asymptotically biased.

### V.A Tenreyro and Thwaites (2016) and the STVAR

Tenreyro and Thwaites (2016) argue that an important advantage of a smooth transition-local projection model is that it does not require to take a stand on how the economy shifts from one regime to another, except for the parameters of a logistic function. However, a local projection model is an appropriate method to estimate response functions, only if the surprises are exogenous (Jordà, 2005), which is not the case for Romer and Romer (2004) monetary policy surprises (Miranda-Agrippino and Ricco, 2021).

Therefore, we estimate the STVAR, employed in Auerbach and Gorodnichenko (2012) and Ramey and Zubairy (2018) to analyze fiscal policy, whose reduced-form has the same features of the smooth transition-local projection model used by Tenreyro and Thwaites (2016), but has the advantage to control for autocorrelation as well as past dynamics of other regressors:

$$\mathbf{X}_t = (1 - F(z_{t-1}))\mathbf{\Pi}_l^{ST}(\mathbf{L})\mathbf{X}_{t-1} + F(z_{t-1})\mathbf{\Pi}_h^{ST}(\mathbf{L})\mathbf{X}_{t-1} + \mathbf{u}_t^{ST}, \quad (7)$$

$$\mathbf{u}_t^{ST} \sim N(0, \mathbf{\Omega}_t^{ST}), \quad (8)$$

$$\mathbf{\Omega}_t^{ST} = \mathbf{\Omega}_l^{ST}(1 - F(z_{t-1})) + \mathbf{\Omega}_h^{ST}F(z_{t-1}), \quad (9)$$

$$F(z_t) = \frac{\exp(\gamma(z_t - c)/\sigma_z)}{1 + \exp(\gamma(z_t - c)/\sigma_z)}, \quad \gamma > 0, \quad (10)$$

where  $\mathbf{X}_t$  denotes the vector of endogenous variables and  $z_t$  the state variable. The logistic function  $F(z_t) = [0, 1]$  determines the smooth transition from one regime to the other with  $\gamma$  being the logistic growth rate,  $c$  the Sigmoid point and  $\sigma_z$  the standard deviation of  $z_t$ . If  $z_t = c$ ,  $F(z_t) = 0.5$ . If the difference between  $z_t$  and  $c$  is positive (i.e. high growth) and large,  $F(z_t) \approx 1$ . If, instead, the difference between  $z_t$  and  $c$  is negative (i.e. low growth) and (in absolute value) large,  $F(z_t) \approx 0$ . As in Tenreyro and Thwaites (2016),  $\gamma$  is set equal to 3,  $c$  is set at the worst 20 percent of the periods in the sample,  $Pr(F(z_t) < 0.5) = 0.2$ , and  $z_t$  is defined as a seven quarter lagging moving average of real quarterly GDP growth.

The differences in the propagation of shocks across regimes is due to differences in covariance matrices for disturbances  $\mathbf{\Omega}_l^{ST}$  and  $\mathbf{\Omega}_h^{ST}$  and differences in lag polynomial  $\mathbf{\Pi}_l^{ST}$  and  $\mathbf{\Pi}_h^{ST}$ .  $\mathbf{\Omega}_l^{ST}$  and  $\mathbf{\Pi}_l^{ST}$  describe the behaviour of the system in a sufficiently low-growth regime (i.e.  $1 - F(z_t) \approx 1$ ) and  $\mathbf{\Omega}_h^{ST}$  and  $\mathbf{\Pi}_h^{ST}$  describe the behaviour of the system in a sufficiently high-growth regime (i.e.  $F(z_t) \approx 1$ ).  $\mathbf{u}_t^{ST}$  is the vector of reduced form residuals and  $\mathbf{\Omega}_t^{ST}$  the time-varying, state-contingent variance-covariance matrix. The first vector of  $\mathbf{u}_t^{ST}$  provides the orthogonalized monetary policy surprises.

Given that we use the database of Tenreyro and Thwaites (2016) over the same sample period 1969Q1 - 2002Q4, we can use the nonlinear monetary policy shocks that they have estimated, as they depend upon the probability of expansion,  $F(z_t)$ , which we can replicate.

The vector  $\mathbf{X}_t = [u_t^{RR}, r_t, y_t, \pi_t]'$  defines the endogenous variables used by Tenreyro and Thwaites (2016), where  $u_t^{RR}$ , are the nonlinear Romer-Romer monetary policy surprises estimated by Tenreyro and Thwaites (2016),  $r_t$  the Federal Fund Rate,  $y_t$  the log of real GDP and  $\pi_t$  the quarterly PCE inflation rate. The STVAR is estimated in a Bayesian framework using the Monte Carlo Markov Chain sampler developed in Galvão and Owyang (2018). Shocks are identified using the recursive method with the Romer-Romer monetary policy surprises ordered first. This allows to clean the variable of interest from the autocorrelation and from the lagged structure of the VAR. Following Tenreyro and Thwaites (2016), the model uses two quarter lags.

The key results are reported in Panel A of Figure 6. The impulse responses conditional on remaining in the same regime in all horizons can be compared with those in Figure 2 of Tenreyro and Thwaites (2016). A monetary policy shock causes a temporary increase in real GDP during recessions and a permanent decline in real GDP during expansions. Both results are in contradiction with a temporary decline in real GDP expected in theory. As for prices, monetary policy shocks have no impact on inflation during recessions and expansions. All in all, the results obtained by Tenreyro and Thwaites (2016) cannot be replicated using a STVAR. This can be explained by the fact that Romer-Romer surprises are autocorrelated (Miranda-Agrippino and Ricco, 2021) and depend upon the past variables of the VAR. A STVAR removes the autocorrelation structure in the Romer-Romer residuals and their predictability.

We carry out the same exercise using the same STVAR structure with four variables and two quarter lags to corroborate the previous findings obtained with the Bauer-Swanson monetary policy surprises. First, we remove the components of the Bauer and Swanson (2023b)'s monetary policy surprises that is predictable using a logistic function:

$$mps_t = (1 - F(z_{t-1}))\alpha'_l \mathbf{X}_{t-} + F(z_{t-1})\alpha'_h \mathbf{X}_{t-} + u_t^{ST-mp}, \quad (11)$$

where  $\mathbf{X}_{t-}$  is the vector of economic news observed prior the FOMC announcement at time  $t$  including the NFCI. To make the results comparable, as in Tenreyro and Thwaites (2016), the Sigmond point of  $F(z_{t-1})$ ,  $c$ , is set at 20% and  $\gamma = 3$ . The results, shown in Panel A of Figure 6, confirm the baseline findings that monetary policy is powerful in recessions,

as after a 25 basis point monetary policy shock output decline by about 0.8% with rather persistent effects and inflation declines temporarily by about 0.3 percentage points. However, in contrast with our findings obtained with the baseline model, in expansion phases, there is no impact on GDP and inflation rises marginally and temporarily. As we show in the robustness section, where the baseline model is used, including forward looking variables in the systems of equations, as recommended by the literature, is important.

## V.B Tenreyro and Thwaites (2016) and the TVAR

To re-evaluate the aforementioned conclusions, we also estimate the TVAR with the same set of variables proposed by Tenreyro and Thwaites (2016). In this case, we first need to estimate the nonlinear Romer-Romer monetary policy surprises consistently with the TVAR.

Therefore, we regress the FFR against the Fed's internal "Greenbook" forecast, provided by Wieland and Yang (2020), as well as on the same set of variables interacted with a dummy defining the low and high economic growth regimes. We follow Tenreyro and Thwaites (2016) in defining the low growth regime (i.e a recession), as the worst 20% of the periods in the sample.

The original Romer and Romer (2004) regression is modified as follows:

$$\Delta FFR_t = \alpha_0 + \alpha_1 \times F_t + \alpha'_2 X_t + \alpha'_3 X_t \times F_t + u_t^{RR}, \quad (12)$$

where  $X_t$  are the control variables employed by Romer and Romer (2004),<sup>12</sup>  $F_t$  is the indicator function which is equal to unity if the underlying GDP growth is above the threshold that guarantees 80% of the observations and  $u_t^{RR}$  are the residuals of the Romer-Romer surprises. When estimated linearly over a common sample, we replicate the results exactly (see Figure 7). The state-dependent Romer-Romer surprises have a 0.890 correlation with the linear series over the 1969-2002 sample. The  $R^2$  of the FFR equation increases from 28.1% in the linear setting to 43.0% in the nonlinear model, suggesting that the response of the Federal Reserve Board to the business cycle is state-dependent. The surprises are then aggregated

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<sup>12</sup>Romer and Romer (2004) regress the change in the intended federal funds rate against the initial level of intended funds rate, the forecasted output growth, inflation and unemployment rate, and the change in forecasted output growth and inflation since the previous meeting over the subsequent two quarters.

to quarterly frequency.

The impulse responses resulting from the TVAR are depicted in Panel B of Figure 6. The results on real GDP resemble those of the STVAR. In addition, there is a price puzzle in expansions. This suggests again that the Romer and Romer (2004)’s monetary policy surprises cannot be treated as exogenous monetary policy shocks.

We carry out the same exercise using the same TVAR structure with four variables and two quarter lags and use the Bauer-Swanson monetary policy surprises. First, we remove the components of the Bauer and Swanson (2023b)’s monetary policy surprises that is predictable using the threshold function 12 and the above mentioned regressors. The results, shown in Panel B of Figure 6, confirm the findings obtained with the STVAR. Monetary policy shocks are more powerful in recession regimes.

## V.C High-frequency policy surprises and local projections

How important is the local projection methodology adopted by Tenreyro and Thwaites (2016) for the results? Local projections are semi-parametric in nature and do not assume a specific model. They are more flexible, but the estimation uncertainty is higher (Kilian and Kim, 2011). Moreover, local projections treat the orthogonalized monetary policy surprises as a shock. Instead, a VAR cleans the monetary policy surprises from autocorrelation and makes it orthogonal to the lags of other variables of the system. Despite these warnings, we test whether the results summarized in Section IV survive using local projections.

The consistently transformed and orthogonalized monetary policy surprises,  $u_t^{ST-mp}$ , are then used as shocks to obtain the response functions from the regime-switching local projection model à la Tenreyro and Thwaites (2016):

$$y_{t+k} = (1 - F(z_{t-1}))(c_l + \beta_{l,k}u_t^{ST-mp} + \eta'_{l,k}(L)\mathbf{X}_{t-1}) + F(z_{t-1})(c_h + \beta_{h,k}u_t^{ST-mp} + \eta'_{h,k}(L)\mathbf{X}_{t-1}) + u_{t+k}^{ST-LP}, \quad (13)$$

where the coefficients  $\beta_{S,k}$  measure the average effect of a shock at horizon  $k = 0, \dots, 48$  as a function of the state of the economy  $F(z_{t-1})$  at the time of the shock,  $c_S$  are state-dependent intercepts,  $\mathbf{X}_t$  are controls,  $F(z_t)$  is the aforedefined logistic function.<sup>13</sup> The regressand  $y_{t+k}$

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<sup>13</sup>We use Jordà (2005)’s code and extend it to the nonlinear case.

in this case empirically accounts for the possible regime changes between periods  $t$  and  $t+k$ .

Moreover, we run a threshold local projection, which is more comparable to our TVAR, running the following model

$$y_{t+k} = (1 - F_t)(c_l + \beta_{l,k}u_t^{T-mp} + \eta'_{l,k}(L)\mathbf{X}_{t-1}) + F_t(c_h + \beta_{h,k}u_t^{T-mp} + \eta'_{h,k}(L)\mathbf{X}_{t-1}) + u_{t+k}^{T-LP}, \quad (14)$$

where  $F_t$  is the indicator function equal to unity if the underlying GDP growth is above the threshold,  $F_t\{z_{t-1} \geq z^*\} = 1$ , and zero otherwise,  $F_t\{z_{t-1} < z^*\} = 0$  and  $u_t^{T-mp}$  are the residuals of the monetary policy surprises, as obtained from Equation 6.

Panels C and D of Figure 6 provide the results based on the smooth transition local projections and the threshold local projections, respectively. The results suggest that monetary policy shocks are more powerful in recessions, but with no clear impact on inflation, when using the Romer-Romer surprises. Conversely, the decline in output is accompanied with a decline in inflation in recessions when using the Bauer-Swanson surprises. All in all, the main difference we find with respect to the results reported in Tenreyro and Thwaites (2016) are not generated by the different response functions' estimators adopted.

## VI Robustness Checks

All the robustness checks are carried out using eight variables, including the forward looking variables of the baseline, and focus on the Bauer-Swanson monetary policy surprises.

As an alternative benchmark, the threshold is estimated through a grid search over possible values of the observations' percentiles. The marginal likelihood is maximised at percentile 0.25, which suggests setting the threshold for underlying real GDP growth at 1.83% annualised. The number of observations in the low growth regime declines to 90 at monthly frequency. The results are displayed in Panel A of Figure 8 and resemble those reported in the baseline Figure 3 for both regimes, corroborating the conclusions that monetary policy shocks are more powerful in recessions. It is useful to emphasise the difference with the results obtained with the four variables model (Panel B of Figure 6), which suggests that forward looking variables are important in extracting the shocks.

It could be argued that monetary policy is better illustrated using PCE inflation, as



employed by Tenreyro and Thwaites (2016). Results and conclusions remain broadly invariant, if substituting the GDP deflator with PCE inflation (see Panel B of 8). Following a contractionary monetary policy shock in the low growth regime, monthly PCE inflation declines and after few months it returns back to its long-run trend in line with economic theory. Conversely, monetary policy does not seem to succeed in bringing goods prices down in high growth regimes.

The literature recommends to use the Minnesota prior together with the dummy-initial-observation prior with VAR in levels, because sampling errors make results erratic in larger models under a flat prior (Sims, 1993; Sims and Zha, 1998). We provide a robustness check using a flat prior, but we reduce the number of lags to six, so that the number of parameters to estimate reduce substantially. The results, which are provided in Panel C of 8, corroborate the key findings of the paper.

The final sets of robustness checks refer to alternative nonlinear methods in estimating the baseline model. The STVAR and the local projection methods corroborate the key finding that monetary policy is more successful in reducing goods prices in low growth regimes (see Figure 9).

## VII Conclusions

We investigate whether the transmission mechanism of the monetary policy on asset prices, output and goods prices are state dependent. This question has been already addressed in the literature. However, we revisit the issue using monetary policy surprises obtained from high-frequency data (e.g. Bauer and Swanson, 2023a,b).

First, we remove the components of the monetary policy surprises that is predictable following the recommendations in Bauer and Swanson (2023a), but in a nonlinear setting. At the same time, we propose the National Financial Condition Index (NFCI) as an additional predictor of the monetary policy surprises, which is relevant in the nonlinear model.

We show that monetary policy shocks are more powerful in the low growth regime, with the negative response on output accompanied by a negative response on goods prices; in the high growth regime, instead the response of goods prices is muted. A Threshold VAR, a

Smooth-Transition VAR and nonlinear local projection methods all corroborate the findings. Across sub-components of real GDP, the impact is a larger on private investment and durable goods than on nondurables and services consumption, consistent with the fact that spending on durables can be postponed in response to adverse shocks.

Regarding asset prices, the response of expected Treasuries along the entire yield curve is positive in expansions but negative in recessions. In periods of low economic growth, a contractionary monetary policy is judged by markets as further conducive to lower nominal economic growth and therefore a decline in the short term interest rates is expected by investors. Bond holders demand higher premia for holding sovereign and corporate bonds in recessions, confirming that the primary driver of the term premia is the real economy and associated risks. As for stock prices, they decline in response to a monetary policy shock, as suggested by textbooks, and the drop is relatively somewhat deeper during recessions.

These results are broadly consistent with theoretical models assuming information asymmetries in financial markets (Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997). During recessions, asset prices drop, business contracts, the net worth of agents declines, the external finance premium rises and finance constraints are more likely to bind. The financial accelerator amplifies the effects of a tightening in monetary policy.

These findings are relevant for the design of stabilization policies. If changes in the monetary policy rates have limited impact on inflation in periods of expansions, central banks may be required to strongly slowdown economic activity with a risk of causing a recession to be able to steer the inflation rate towards the desired target.

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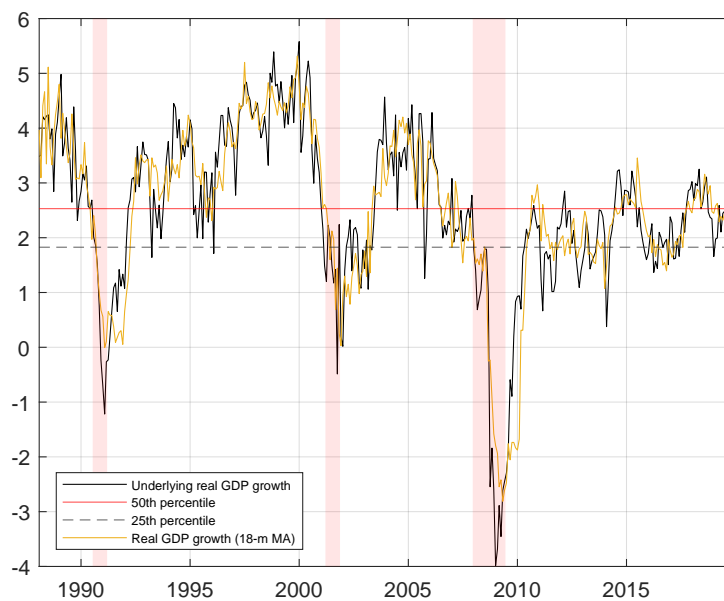
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# Figures and Tables

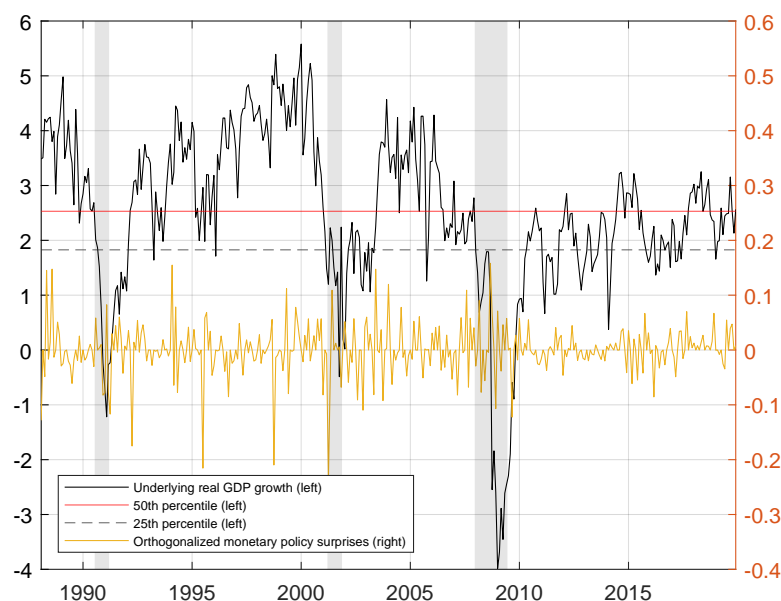
**Figure 1:** Underlying Real GDP Growth and 18-month Moving Average  
(*annualised and %*)



Note: The state variable is calculated as  $z_t = \alpha(y_t - y_{t-1}) + (1 - \alpha)z_{t-1}$ , where  $y_t$  is the log of real GDP and  $\alpha = 0.125$ . The exponential weighted moving average is shown annualised and in percent. Real GDP growth (18-m MA) is computed as  $100(y_t - y_{t-18})/3$ . Sample period: Feb. 1988 - Dec. 2019.

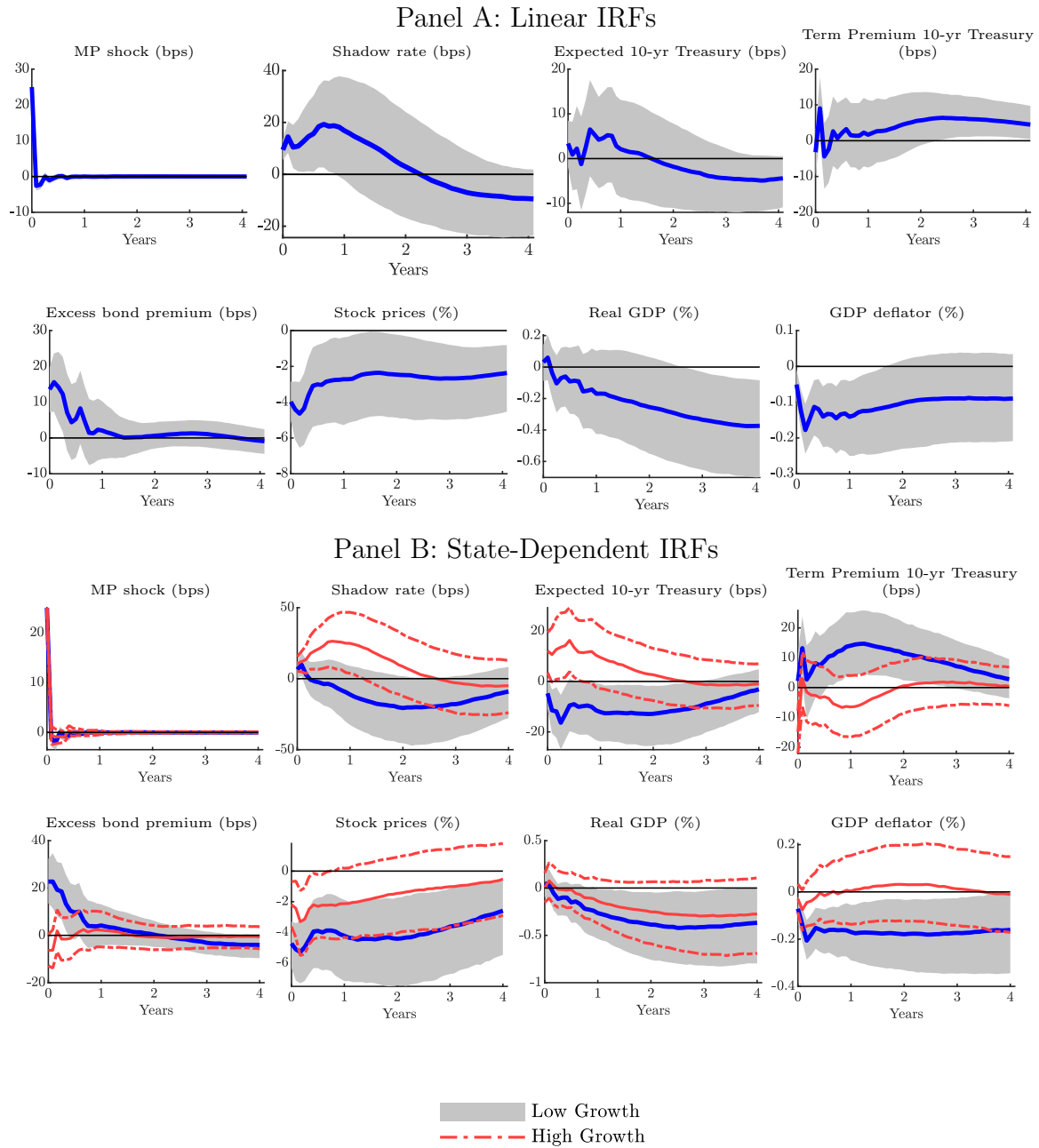
**Figure 2:** Underlying Real GDP Growth and Orthogonalised Monetary Policy Surprises

(left: annualised and %; right: percentage points)



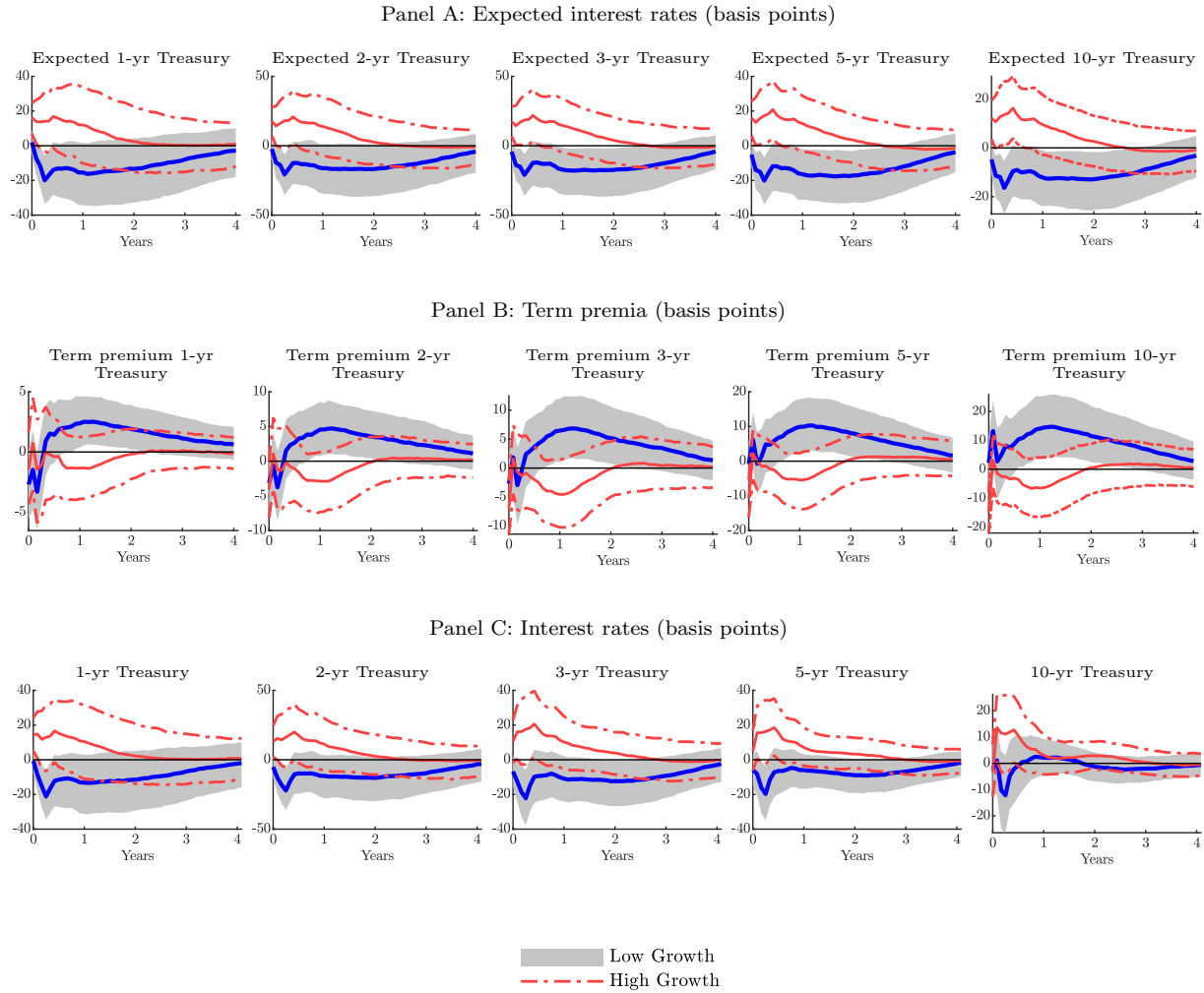
Note: The state variable is calculated as  $z_t = \alpha(y_t - y_{t-1}) + (1 - \alpha)z_{t-1}$ , where  $y_t$  is the log of real GDP and  $\alpha = 0.125$ . The exponential weighted moving average is shown annualised and in percent. The monetary policy shocks are defined as the residuals of an OLS equation (see column 4 of Table 1), where the monetary policy surprises are regressed on (i) nonfarm payrolls surprises; (ii) employment growth from one year earlier to the most recent release before the FOMC announcement; (iii) the log change on the S&P500 stock price index from three months before the FOMC announcement to the day before the FOMC announcement; (iv) the change in the slope of the yield curve from three months before the FOMC announcement to the day before the FOMC announcement; (v) the log change of commodity prices from three months before the FOMC announcement to the day before the FOMC announcement; (vi) the implied skewness of the ten-year Treasury yield; (vii) the change in the NFCI from seven weeks before the FOMC announcement to the day before the FOMC announcement as well as on the same set of variables interacted with a dummy defining the high economic growth regime. Monetary policy surprises and the first six factors are provided by Bauer and Swanson (2023a), the NFCI real time vintages are provided by Amburgey and McCracken (2023). The monetary policy shocks are summed over the month and are evaluated in percentage points. Sample period: Feb. 1988 - Dec. 2019.

**Figure 3:** Nonlinear Impact of Monetary Policy Shocks  
*(response to a 25 basis points monetary policy shock)*



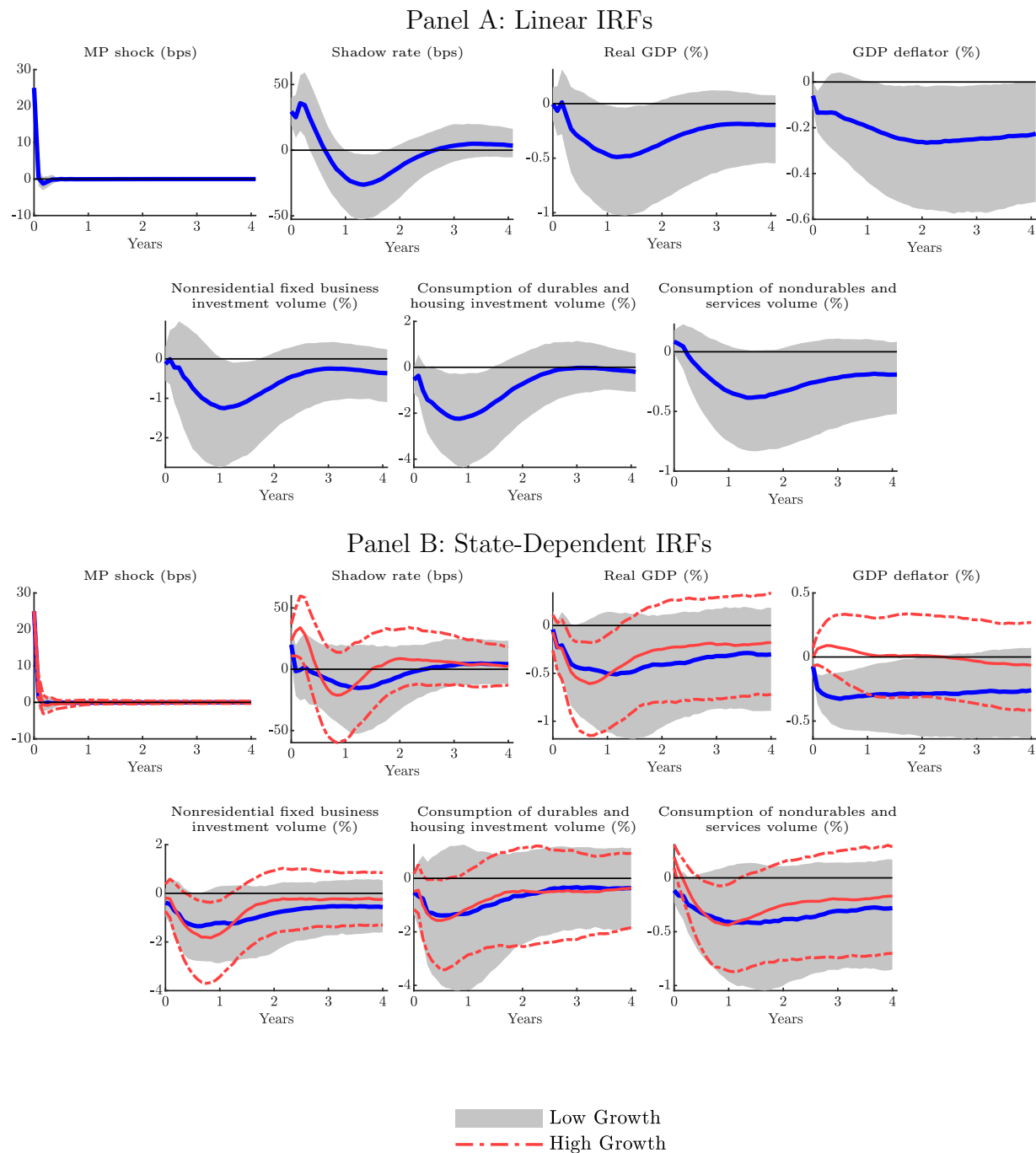
*Notes:* The vector  $\mathbf{X}_t = [u_t^{T-mp}, r_t, i_t^{en}, i_t^{tpn}, b_t, e_t, y_t, p_t]'$  defines the endogenous variables (see text for definitions). The monetary policy shocks are defined as the residuals of an OLS equation (see column 4 of Table 1 and notes in Figure 2). The threshold is set at the median. Each panel shows the median IRFs and the corresponding posterior 68% credible sets (dashed red lines and shaded bands). The red (blue) lines in Panel B are associated to the high (low) growth regime. Sample period: Feb. 1988 - Dec. 2019.

**Figure 4:** Nonlinear Impact of Monetary Policy Shocks on the Yield Curve  
*(response to a 25 basis points monetary policy shock)*



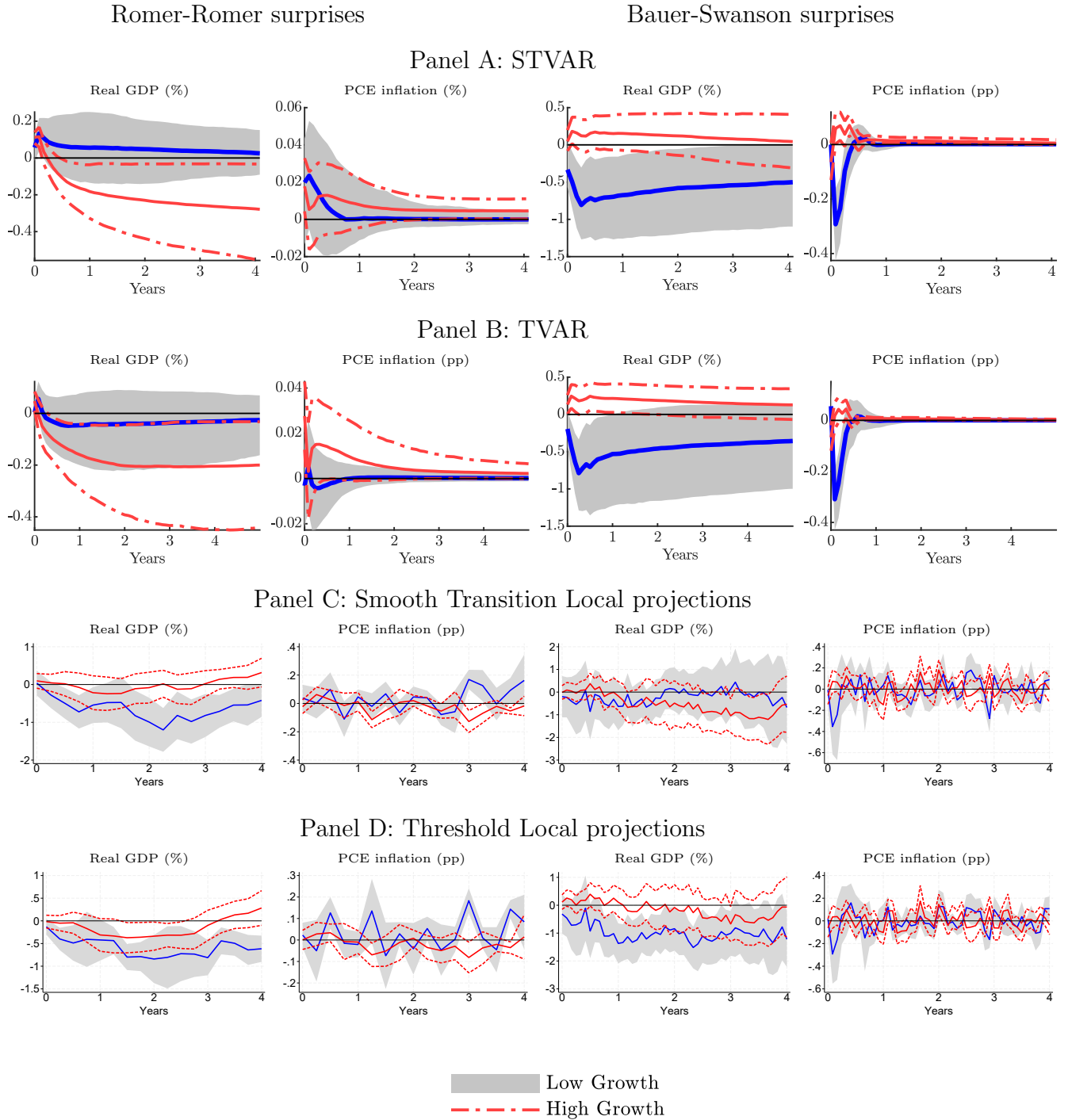
Notes: The vector  $\mathbf{X}_t = [u_t^{T-mp}, r_t, i_t^{en}, i_t^{tpn}, b_t, e_t, y_t, p_t]'$  defines the endogenous variables (see text for definitions).  $i_t^{en}$  and  $i_t^{tpn}$  can take 1-yr, 2-yr, 3-yr, 5-yr and 10-yr maturity. The monetary policy shocks are defined as the residuals of an OLS equation (see column 4 of Table 1 and notes in Figure 2). The threshold is set at the median. Each panel shows the median IRFs and the corresponding posterior 68% credible sets (dashed red lines and shaded bands). The red (blue) lines are associated to the high (low) growth regime. Sample period: Feb. 1988 - Dec. 2019.

**Figure 5:** Nonlinear Impact of Monetary Policy Shocks on Expenditure Aggregates  
(response to a 25 basis points monetary policy shock)



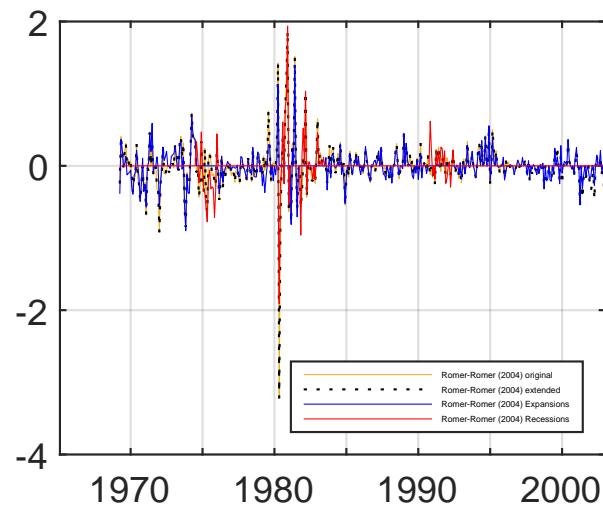
The vector  $\mathbf{X}_t = [u_t^{T-mp}, r_t, y_t, p_t, y_t^f, y_t^d, y_t^n]'$  defines the endogenous variables (see text for definitions). The monetary policy shocks are defined as the residuals of an OLS equation (see column 4 of Table 1 and notes in Figure 2). The threshold is set at the median. Each panel shows the median IRFs and the corresponding posterior 68% credible sets (dashed red lines and shaded bands). The red (blue) lines in Panel B are associated to the high (low) growth regime. Sample period: Q1 1988 - Q4 2019.

**Figure 6: Policy Surprises versus Methods in Recessions and Expansions**  
*(response to a 25 basis points monetary policy shock)*



*Notes:* The vector  $\mathbf{X}_t = [u_t^{mp}, r_t, y_t, \pi_t]'$  defines the endogenous variables (see text for definitions).  $u_t^{mp}$  varies depending upon the nonlinear model employed. The first and the second columns use Romer-Romer surprises and the data set of Tenreiro and Thwaites (2016) over the quarterly sample period 1969Q1 - 2002Q4. The third and the fourth columns use Bauer-Swanson surprises and the data set over the monthly sample period Feb. 1988 - Dec. 2019. Panels A and B provide the state-dependent responses based on the STVAR and TVAR. Panels C and D provide the state-dependent responses based on the smooth transition and threshold local projections. The Romer-Romer surprises are defined as the residuals of an OLS equation following Romer and Romer (2004) and consistent with the nonlinear models. The Bauer and Swanson surprises are defined as the residuals of an OLS equation using the factors of column 4 of Table 1 and consistent with the nonlinear models. As in Tenreiro and Thwaites (2016), the threshold is set at the 20th percentile of the real GDP growth distribution. Each panel shows the median IRFs and the corresponding posterior 68% credible sets. The transition to the other regime is not allowed. The red (blue) lines are associated to the high (low) growth regime.

**Figure 7:** Romer-Romer Surprises in Recessions and Expansions

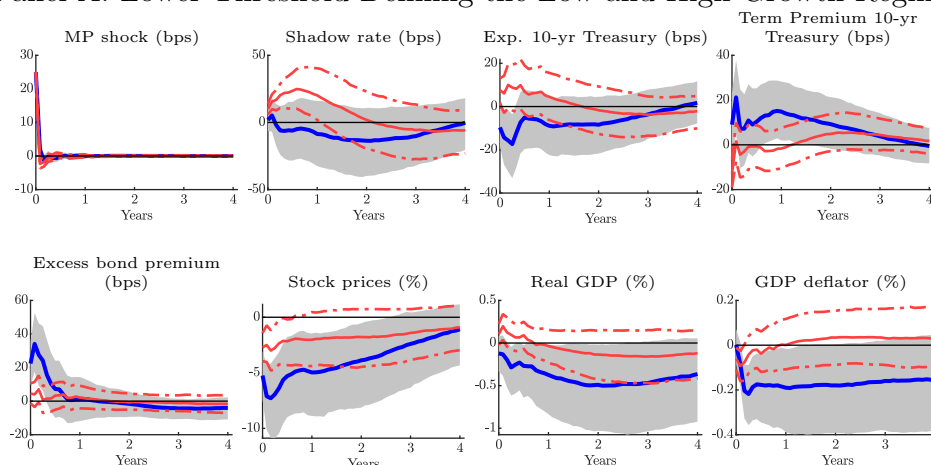


Notes: The orange solid line is the series of monetary policy surprises in Romer and Romer (2004) between March 1969 and December 1996. The black dotted line is the series of monetary policy surprise in Romer and Romer (2004) extended to December 2002. The remaining line is the monetary policy surprises in high growth (expansions, blue) and low growth (recessions, red) regimes. See main text for details. Sample period: Mar. 1969 - Dec. 2002.

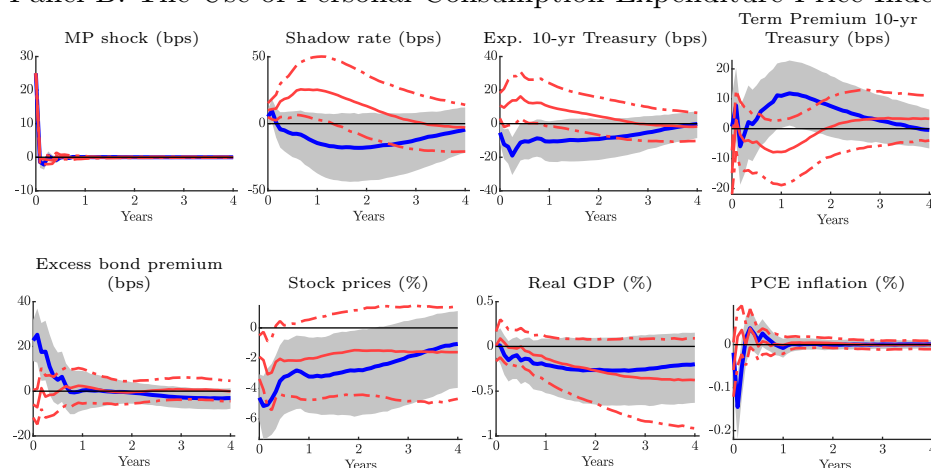


**Figure 8:** Robustness Checks of the Baseline Model: State-Dependent IRFs  
*(response to a 25 basis points monetary policy shock)*

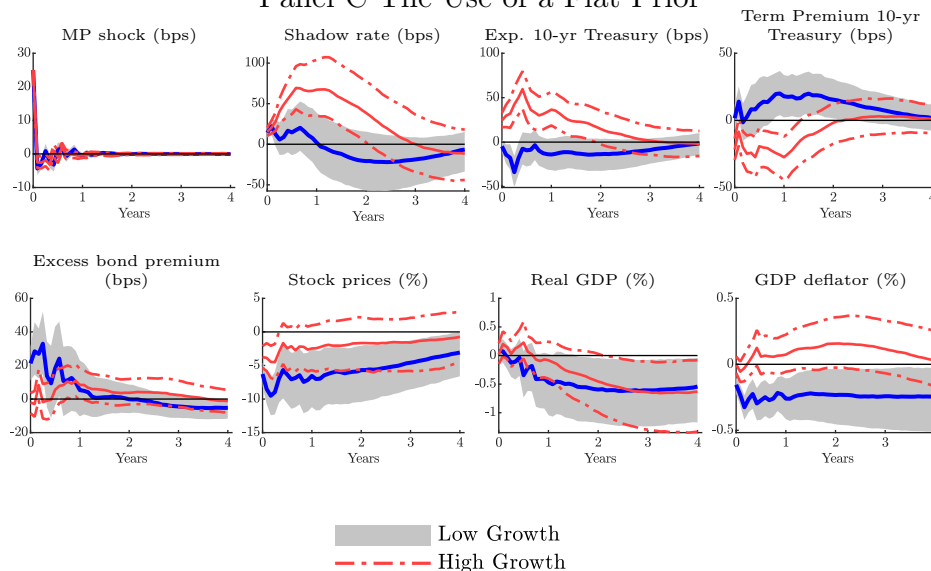
**Panel A: Lower Threshold Defining the Low and High Growth Regimes**



**Panel B: The Use of Personal Consumption Expenditure Price Index**



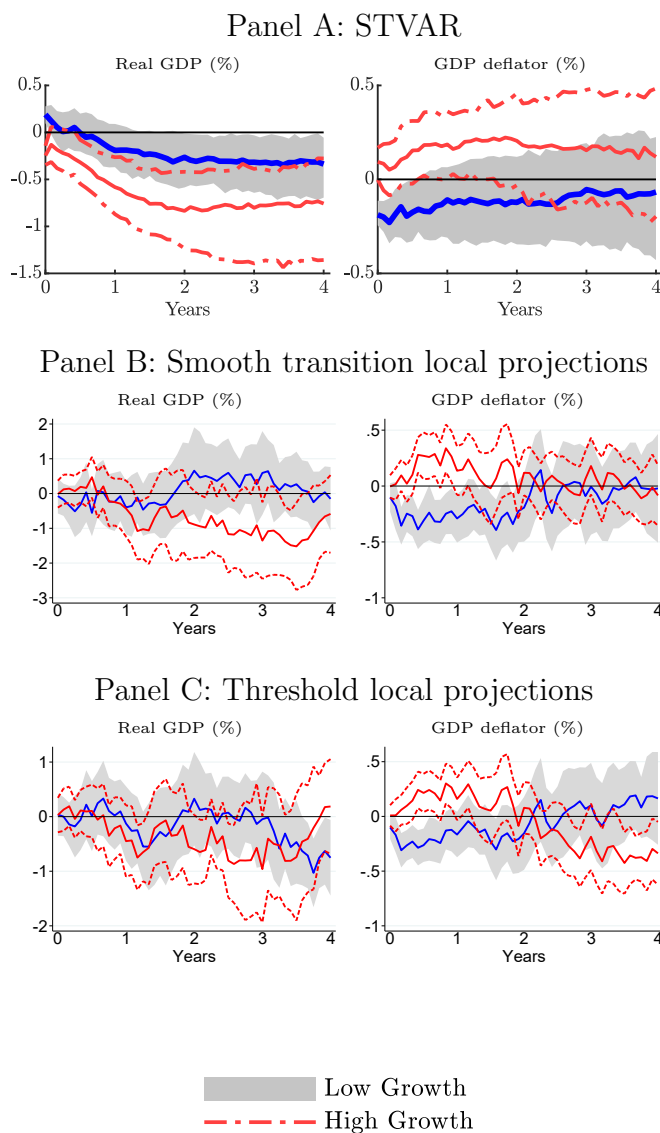
**Panel C The Use of a Flat Prior**



*Notes:* Panel A provides the state-dependent IRFs when setting the threshold of underlying real GDP growth at the worst 25 percent of the periods in the sample. Panel B provides the state-dependent IRFs when substituting the GDP deflator with PCE inflation. Panel C provides the state-dependent IRFs when using a flat prior. See also notes of Figure 3.

**Figure 9:** State-Dependent IRFs using alternative Nonlinear Models

*(response to a 25 basis points monetary policy shock)*



*Notes:* The vector  $\mathbf{X}_t = [u_t^{mp}, r_t, i_t^{en}, i_t^{tpn}, b_t, e_t, y_t, p_t]'$  defines the endogenous variables (see text for definitions).  $u_t^{mp}$  varies depending upon the nonlinear model employed. The monetary policy shocks are defined as the residuals of an OLS equation using the factors of column 4 of Table 1. The monetary policy shocks are obtained from Equation 11 in the case of the smooth transition model and Equation 6 in the case of the threshold model. Both the Sigmoid point of the logistic function and the threshold are set at the median. Each panel shows the median IRFs and the corresponding posterior 68% credible sets (dashed red lines and shaded bands). The red (blue) lines in Panel B are associated to the high (low) growth regime. Sample period: Feb. 1988 - Dec. 2019.

Table 1: Predictive Regressions using Macroeconomic and Financial Data

	Linear BS Surprises (1)	Linear BS & NFCI (2)	Nonlinear BS Surprises (3)	Nonlinear BS & NFCI (4)
Nonfarm payrolls	0.094** (2.390)	0.092** (2.330)	0.108* (1.860)	0.082 (1.540)
Empl. growth (12m)	0.005** (2.090)	0.004* (1.750)	0.005 (1.560)	0.002 (0.610)
$\Delta \log$ S&P 500 (3m)	0.084 (1.430)	0.100 (1.630)	0.083 (1.290)	0.160** (2.400)
$\Delta$ Slope (3m)	-0.010 (-1.370)	-0.010 (-1.380)	-0.017 (-1.600)	-0.020** (-2.040)
$\Delta \log$ Comm. Price (3m)	0.119** (2.350)	0.123** (2.460)	0.139** (2.000)	0.170*** (2.870)
Treasury skewness	0.032*** (2.99)	0.033*** (3.090)	0.032** (2.030)	0.033** (2.360)
$\Delta$ NFCI (7w)	- -	0.014 (0.620)	- -	0.066*** (3.000)
Nonfarm payrolls $\times D$	- -	- -	-0.010 (-0.120)	0.010 (0.120)
Empl. growth (12m) $\times D$	- -	- -	0.002 (0.320)	0.006 (1.230)
$\Delta \log$ S&P 500 (3m) $\times D$	- -	- -	0.003 (0.020)	-0.144 (-1.060)
$\Delta$ Slope (3m) $\times D$	- -	- -	0.010 (0.660)	0.013 (0.850)
$\Delta \log$ Comm. Price (3m) $\times D$	- -	- -	-0.052 (-0.560)	-0.082 (-0.980)
Treasury skewness $\times D$	- -	- -	0.002 (0.090)	-0.003 (-0.170)
$\Delta$ NFCI (7w) $\times D$	- -	- -	- -	-0.135*** (-3.790)
$R^2$	0.162	0.164	0.171	0.223
Sample	1988:2-2019:12	1988:2-2019:12	1988:2-2019:12	1988:2-2019:12
N	322	322	322	322
Wald test (P-value)	-	-	0.990	0.020
Wald test on NFCI (P-value)	-	-	-	0.014

Note: This table shows the coefficient estimates  $\beta$  and  $\gamma$  from the predictive regressions  $mps_t = \alpha_0 + \alpha_1 \times D + \beta' X_{t-} + \gamma' X_{t-} \times D + u_t$ , where  $t$  indexes FOMC announcements. Columns 1 replicates the results of Bauer and Swanson (2023a), column 2 controls also for the NFCI, columns 3 and 4 includes an interaction dummy  $D$ , which is equal to 1 if the underlying real GDP growth is above the threshold ( $I\{z_{t-1} \geq z^*\}$ ). Predictors  $X$  are observed prior to the FOMC announcement: the surprise component of the most recent nonfarm payrolls release, employment growth over the last year, the log change in the Standard & Poor's 500 index (S&P 500) from 3 months before, to the day before, the FOMC announcement, the change in the yield curve slope over the same period, the log change in a commodity price index over the same period and the option-implied skewness of the 10-year Treasury yield are obtained from Bauer and Swanson (2023a); the change in the NFCI from 9 weeks before, to the day before, the FOMC announcement is computed using the real time data of Amburgey and McCracken (2023). Bootstrapped t-stat using 50,000 bootstrap replications are reported in parentheses. The Wald test of the joint significance of coefficients is performed on all the interaction terms:  $\hat{\gamma}' = 0$ . The Wald test on  $\Delta$  NFCI is a test on  $\hat{\beta}_{NFCI} + \hat{\gamma}_{NFCI} = 0$ . \*\*\*, \*\*, \* denote statistical significance at 1%, 5% and 10% level.

Table 2: State-Dependent Responses of Asset Prices at Impact and Six Months Ahead

	In the current month				After six months			
	Low		High		Low		High	
Shadow rate (bps)	16%	50%	84%	16%	50%	84%	16%	84%
1-year Treasury, EXP (bps)	0.64	6.42	11.91	5.11	10.60	15.48	-16.57	12.31
2-year Treasury, EXP (bps)	-4.95	1.99	8.99	6.88	16.14	24.75	-28.26	2.09
3-year Treasury, EXP (bps)	-9.85	-2.29	5.74	7.28	17.54	27.74	-27.50	3.00
5-year Treasury, EXP (bps)	-11.79	-4.28	4.10	6.98	17.43	28.28	-28.69	3.32
10-year Treasury, EXP (bps)	-12.31	-5.10	2.88	5.86	16.03	25.61	-26.11	1.86
1-year Treasury, TP (bps)	-10.79	-4.72	1.84	3.53	12.13	19.46	-19.79	1.20
2-year Treasury, TP (bps)	-5.53	-2.70	0.08	-4.34	-1.47	1.64	-1.12	4.33
3-year Treasury, TP (bps)	-6.97	-3.16	0.62	-8.22	-4.09	-0.15	-1.65	6.56
5-year Treasury, TP (bps)	-7.24	-2.68	2.08	-11.40	-6.89	-2.22	-2.35	8.54
10-year Treasury, TP (bps)	-6.92	-0.85	5.69	-16.17	-10.73	-5.35	-2.80	12.16
Excess bond premium (bps)	-6.13	2.21	11.28	-22.04	-14.86	-7.85	-3.51	17.35
log of S&P 500 (%)	13.55	22.70	32.33	-12.86	-6.36	0.35	-1.19	22.03
	-6.35	-4.74	-3.31	-3.63	-2.19	-0.69	-6.35	-1.87

Note: This table shows the response at impact and after six months of the policy shadow rate, the expected Treasury rate and the correspondent term premium at maturity  $n$ , the corporate excess bond premium and the log of the stock market price. The responses of the policy shadow rate, the corporate excess bond premium and the log of the stock market price are taken from the model which uses the 10-year expected Treasury rate and the correspondent term premium. The monetary policy shocks are defined as the residuals of an OLS equation (see column 4 of Table 1), where the monetary policy surprises are regressed on (i) nonfarm payrolls surprises; (ii) employment growth from one year earlier to the most recent release before the FOMC announcement; (iii) the log change on the S&P500 stock price index from three months before the FOMC announcement to the day before the FOMC announcement; (iv) the change in the slope of the yield curve from three months before the FOMC announcement to the day before the FOMC announcement; (v) the log change of commodity prices from three months before the FOMC announcement to the day before the FOMC announcement; (vi) the implied skewness of the ten-year Treasury yield; (vii) the change in the NFICI from three months before the FOMC announcement to the day before the FOMC announcement as well as on the same set of variables interacted with a dummy defining the low economic growth regimes. Monetary policy surprises are provided by Bauer and Swanson (2023a) and the NFICI is provided by the Federal Reserve of Chicago. 16%, 50% and 84% provides the posteriors at the 16%, 50% and 84% percentile, respectively. Sample period: Feb. 1988 - Dec. 2019.

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