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## Less bang for the buck? Assessing the role of inflation uncertainty for U.S. monetary policy transmission in a data rich environment

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**LESS BANG FOR THE BUCK? ASSESSING  
THE ROLE OF INFLATION  
UNCERTAINTY FOR U.S. MONETARY  
POLICY TRANSMISSION IN A DATA  
RICH ENVIRONMENT**

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# Less bang for the buck? Assessing the role of inflation uncertainty for U.S. monetary policy transmission in a data rich environment \*

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November 30, 2018

## Abstract

We investigate the relationship between inflation uncertainty and monetary policy transmission in the U.S. economy. Monetary policy shocks are identified within the framework of nonlinear structural factor-augmented VARs which allow us to analyze several complementary hypotheses connecting IU with reduced monetary policy effectiveness. We find that the real effects of monetary policy shocks are markedly dampened conditional on high IU. This can be traced back to, inter alia, real-option and precautionary savings effects which distort the traditional interest rate channel. Moreover, policy transmission through the external finance premium and the term structure of interest rates appears strongly dependent on inflation uncertainty and contributes to the reduced policy effectiveness.

**JEL Classification:** C32, E44, E52

**Keywords:** inflation uncertainty, SVAR, monetary policy, sign restrictions, asset prices, smooth-transition

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# 1 Introduction

Recently, theoretical and empirical advances as well as events, such as the Great Recession, have turned the academic spotlight on economic uncertainty as a determinant of fluctuations in real macroeconomic variables. A growing literature, sparked by [Bloom \(2009\)](#) has investigated the role of uncertainty within linear model frameworks (e.g. [Nodari, 2014](#); [Gilchrist et al., 2014](#); [Jurado et al., 2015](#); [Mumtaz and Surico, 2013](#)). Several theoretical mechanisms have been proposed for explaining the role of economic uncertainty in shaping the real economy and imply both positive as well as negative effects (see [Bloom, 2014](#); [Castelnuovo et al., 2017](#), for a detailed review). Seeing strong empirical support for channels that imply detrimental effects of uncertainty on macroeconomic performance ([Fernández-Villaverde et al., 2011](#); [Leduc and Liu, 2016](#); [Bloom, 2014](#); [Gilchrist et al., 2014](#); [Mumtaz and Surico, 2013](#)), we next sketch major theoretical foundations of adverse linkages between economic uncertainty and activity.<sup>1</sup>

Deducting major implications of uncertainty for investment and consumption expenditures from the theory of “real options”, a main line of research has established the well noted “wait-and-see” effect ([Bernanke, 1983](#); [Bloom et al., 2012](#)). Facing sunk costs or partial irreversibility of investment decisions, firms become increasingly reluctant to invest or hire when the success of an investment choice becomes less certain.<sup>2</sup> Another prominent strand of the literature emphasises the role of financial frictions caused by elevated uncertainty and the resulting detrimental impact on investment ([Christiano et al., 2014](#); [Gilchrist et al., 2014](#)). More recently, research has focused on the potentially nonlinear effects of uncertainty shocks on real economic indicators, whereas uncertainty as a conditioning scenario for the transmission of other structural shocks has not yet attracted comparable attention.<sup>3</sup> This applies particularly to monetary policy shocks, notwithstanding the ample evidence for the nonlinear dynamic propagation of such shocks, for instance, along financial conditions or the business cycle ([Tenreyro and Thwaites, 2016](#); [Balke, 2000](#); [Eickmeier et al., 2015](#)). Despite the role of monetary policy shocks in macroeconomic research and the clear-cut implications of the above mentioned theories ([Christiano et al., 1999](#); [Ramey, 2016](#)), only a handful studies have addressed the viability of monetary policy transmission channels under distinguished states of uncertainty ([Eickmeier et al., 2016](#); [Aastveit et al., 2017](#); [Pellegrino, 2017, 2018](#); [Castelnuovo and Pellegrino, 2018](#)). Reviewing the evidence from these studies, aggregate activity, in particular interest-rate sensitive components, will be less responsive to monetary stimuli during episodes of elevated uncertainty either due to the “wait-and-see” effect ([Aastveit et al., 2017](#); [Bloom et al., 2012](#)), worsened financing conditions or reduced nominal frictions ([Vavra, 2013](#); [Bachmann et al., 2013](#)). In this work, we narrow down the broadly defined notion of uncertainty and consider the effectiveness of monetary policy conditional on fluctuations in inflation uncertainty (henceforth IU). To the best of our knowledge this study is the first attempt to explicitly focus on IU in this context.<sup>4</sup>

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<sup>1</sup>Theories predicting positive impacts, are, for instance, cash-in-advance type models or the so-called growth options theory ([Kraft, Schwartz, and Weiss \(2018\)](#), see also [Bloom \(2014\)](#)).

<sup>2</sup>The real-options argument applies similarly to households, which might react with spending cutbacks and postponing expenditures for durable consumption.

<sup>3</sup>State-dependent impacts of uncertainty shocks have been examined, e.g., by [Caggiano et al. \(2014, 2017a,b\)](#); [Mumtaz and Theodoridis \(2017\)](#); [Popp and Zhang \(2016\)](#). These studies consistently find that the propagation of uncertainty shocks is indeed varying with the business cycle or changing over time ([Mumtaz and Theodoridis, 2017](#)).

<sup>4</sup>Numerous studies contribute to the understanding of specific relations, such as the interaction between IU and the yield curve ([Berument et al. \(2005\)](#) and [Fountas et al. \(2006\)](#) provide an detailed review of this literature). The role of IU for nonlinearities in monetary policy transmission has not been analysed yet. This appears surprising even more so when considering the efforts that have been devoted to disentangling notions of intrinsically latent uncertainty along several

Since similar theoretical channels have been identified for the detrimental impact of IU on real economic variables, one might presume that the nonlinear interaction between IU and policy transmission resembles the interaction with other types of uncertainty (Pindyck, 1993; Huizinga, 1993; Fountas et al., 2006; Binder, 2017). However, the specific characteristics of IU may cast doubt on this conjecture. On the one hand, the IU statistics that we employ in this study exhibit fairly distinct time series characteristics compared with typically adopted indicators of economic uncertainty (e.g. stock market volatility). On the other hand, given a well established link between IU and the term premia of fixed income securities (Piazzesi et al., 2006; Rudebusch and Swanson, 2008; Wright, 2011; Ehling et al., 2018), fluctuations in IU might induce additional nonlinearities in monetary policy transmission via the interest rate channel. Accordingly, we expect that elevated (long-term) IU mitigates the stimulative impact of monetary easing and likely invokes asymmetries in regime dependent responses to expansionary or contractionary shocks for both theoretical (Rudebusch and Swanson, 2012) and empirical reasons (Tillmann, 2017).

In order to unravel stylized characteristics of monetary policy transmission conditional on states of low and elevated IU, our empirical strategy consists of extending a linear structural vector autoregressive (SVAR) model along three dimensions: First, following Auerbach and Gorodnichenko (2012) we allow for regime dependence (low vs high IU) in the autoregressive coefficients and the structural relations by means of a smooth-transition (ST-VAR). Second, we consider the responses of a broad set of macroeconomic and financial variables by adopting a factor-augmented VAR (FAVAR, Bernanke et al., 2005; Stock and Watson, 2016) in the spirit of Popp and Zhang (2016). In contrast to small- and medium-scale VARs, the FAVAR specification offers the advantage to alleviate issues linked to deficient information (Forni et al., 2009). In addition, it enables us to deal with a variety of complementary sources of regime dependence and carve out more precise policy implications. Third, we refrain from imposing recursiveness and adopt a sign restriction approach (Faust, 1998; Uhlig, 2005) to identify of the monetary policy shock and characterize the interdependence between monetary policy, asset prices and other financial indicators (see, e.g., Björnland and Leitemo, 2009; Lütkepohl and Netšunajev, 2014; Beckers and Bernoth, 2016).

Modelling the state-conditional monetary policy transmission particularly with regard to (inflation) uncertainty faces two main challenges: First, it requires the sensible choice of a measure for the unobservable notion of uncertainty. Second, since uncertainty typically spikes during recessions (Bloom, 2014), effects of uncertainty have to be disentangled from other sources of regime dependence during recessions (Aastveit et al., 2017). To address these issues, we rely in our baseline specification on a forecasting based measure for the average individual IU which mimics inflation survey deducted uncertainty and has been proven to outperform most other model IU metrics. Moreover, our baseline measure is only weakly related to recessionary phases, in contrast to commonly used general uncertainty indicators.

In the baseline setup, we apply the ST-FAVAR to a monthly US panel of 148 macroeconomic and financial time series for the period from 1977M6 until 2015M12. Our main results are the following. First, we detect a reduced monetary policy effectiveness for steering macroeconomic aggregates, such as, industrial production, consumption and labour market indicators in episodes of high IU. These results admit a role for both IU induced wait-and-see and precautionary savings effects on monetary policy transmission. Second, besides evidence for an attenuated interest rate

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sub-categories, such as financial (Bloom, 2009), fiscal (Baker et al., 2016; Fernández-Villaverde et al., 2011), monetary policy (Tillmann, 2017), micro-firm (Gilchrist et al., 2014), sales growth, and general macroeconomic uncertainty (Jurado et al., 2015).

channel, our analysis indicates that IU interferes with monetary policy transmission through the term-structure. The responses of government bond yields of different maturities and corresponding term spreads suggest that policy rate transmission is modestly attenuated at high levels of IU. While there are regime asymmetries in both expected short-term rates and term premia, these are more pronounced for the former. The response pattern of term premia is similar and the effect is equally, quantitatively negligible in both regimes. This contradicts the view that IU might be an important determinant of (nominal) term premia (Wright, 2011). Third, regime-dependent price flexibility appears to play only a minor role in explaining the asymmetries in monetary policy effectiveness and, hence, puts the model predictions of Vavra (2013) into question. Finally, we find marked evidence for the interaction between IU and financial frictions. Based on several credit risk spreads and indicators for the risk-bearing capacities of the financial sector, we document weaker effects of a monetary easing on the external finance premium during periods of elevated IU. Accordingly, our results suggest that the policy transmissions through the balance-sheet channel or the supply-side bank lending channel are perturbed by IU (Gilchrist et al., 2014; Eickmeier et al., 2016). The remainder of this paper is organized as follows. Section 2 provides details on the theoretical channels which link IU to monetary policy transmission and reviews related empirical work. Section 3 outlines the econometric model (subsection 3.1), estimation and identification strategy (subsection 3.2) and the inflation uncertainty statistic (subsection 3.3). We introduce the data set in Section 4.1, while we present the empirical results from our baseline model in Section 4.2 and discuss its robustness with regard to variations of the data and specification details in Sections 4.3 and 4.4. Finally, Section 5 concludes.

## 2 Uncertainty and monetary policy

Based on the theoretical literature and the available body of empirical evidence, we carve out a set of testable hypotheses on the role of IU for channelling monetary policy signals in this section. First, we review some of the theoretical channels linking monetary policy transmission and uncertainty in general. Second, we discuss the specifics of the interaction between IU and monetary policy transmission. Third, we review empirical studies linking IU and monetary policy. Fourth, we summarize and state hypotheses subjected to empirical modelling in Section 4.

### 2.1 Uncertainty and monetary policy effectiveness from a theoretical perspective

The theoretical literature has put forth several complementary channels to explain the (regime) dependence of monetary policy transmission on the level of uncertainty. We will give a brief overview of four main approaches.<sup>5</sup>

The **first**, and probably most prominent, channel builds on the theory of ‘real options’ and its implications for investment and consumption expenditures (Bernanke, 1983; Pindyck, 1993). According to this theory, firms’ investment choices constitute real options in the presence of adjustment costs (partial irreversibility) and (some) timing flexibility. The option value of delay rises with uncertainty about the success of an investment (i.e., future cash flows). Hence, firms become increasingly reluctant to exercise their options and postpone investment and hiring. This applies likewise to households, which become cautious about spendings on durable consumption goods.

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<sup>5</sup>Bloom (2014) and Castelnuovo et al. (2017) provide a comprehensive review of the uncertainty literature.

As a result, wait-and-see effects reduce the sensitivity of investment and durable consumption to monetary stimuli (Aastveit et al., 2017; Bloom et al., 2012).

The **second broad channel** summarizes mechanisms related to financial frictions and/or risk aversion. Assuming risk aversion and partly binding financial constraints, Christiano et al. (2014) and Gilchrist et al. (2014) develop models which predict a detrimental impact of uncertainty on investment which originates in financial frictions. For instance, in the quantitative business cycle model of Gilchrist et al. (2014) elevated uncertainty leads to both the wait-and-see-effect and higher user costs of capital via the external finance premium. The latter effect is deduced from option pricing theory applied to the liabilities of firms. Accordingly, the payoff scheme of equity holders corresponds to that of a long position call option, while the payoff structure for bondholders resembles that of a short position put option. Therefore, increasing uncertainty (i.e., higher downside risk) benefits shareholders over bondholders who likely demand higher risk premia. Consequently, both credit availability and cost conditions worsen and invoke slowdowns in investment. Complementing insights of Christiano et al. (2014) and Gilchrist et al. (2014) into the balance-sheet channel of firms, Eickmeier et al. (2016) and Alessandri and Bottero (2017) consider the supply side, i.e., the bank lending channel of monetary policy. Introducing state dependent leverage of financial intermediaries into a New-Keynesian DSGE model weakens the bank lending channel during uncertain times.

Two further generally observed properties of agents/effects threaten the viability of transmission channels of monetary policy under elevated general uncertainty. On the one hand, risk averse economic agents might enhance precautionary savings in uncertain times and thereby reduce consumption expenditures of households, or, similarly, raise the cash holdings of firms to shield against future adverse shocks accompanied by lower investment activity (see Basu and Bundick (2017); Alfaro et al. (2016) and Baum et al. (2006, 2009)). This mechanism, however, is ambiguous in the long-term, since increased savings could potentially lower interest rates and reverse the relationship. On the other hand, the price setting behaviour of firms (menu costs, informational frictions) might trigger the transmission of monetary signals according to the state of uncertainty. The main insight from this reasoning is that price changes become more frequent and more dispersed when uncertainty is high entailing a reduction in real policy effectiveness (Vavra, 2013; Baley and Blanco, 2016; Castelnuovo and Pellegrino, 2018).

## 2.2 Inflation uncertainty and monetary policy

So far we have remained silent on how uncertainty is introduced into the models mentioned previously and on the applicability of the channels outlined above to IU. Typically, the variance of firm-specific idiosyncratic shocks introduces uncertainty into the models (e.g., Gilchrist et al., 2014; Christiano et al., 2014). It is common to the majority of theoretical studies that their “uncertainty generating mechanism” is consistent with some form of stock return volatility (Bloom et al., 2012). Based on available theoretical and empirical evidence, however, it is difficult to assess to what extent the model implied linkages between specific types of uncertainty and monetary policy apply as well to IU. Theoretical models that use IU as an explicit source of real option effects are scarce (Huizinga, 1993; Ghosal and Loungani, 1996). There are, however, various indirect linkages among IU and the above mentioned channels. For instance, Huizinga (1993) identifies several variables (e.g., real wages, real output prices and profit rates) which establish a connection between IU and the real net value of investment projects. Moreover, IU likely transmits uncertainty into the real

income of households and could therefore increase precautionary savings (see [Binder, 2017](#), and references therein). In general, IU could be linked to all channels discussed above by influencing the uncertainty about long-term contracts, real interest rates and commodity, factor and input/output prices.<sup>6</sup> With regard to downside risks of firms, however, IU might be of minor importance in comparison with notions of uncertainty aligned with stock market volatility.

Complementing effects on the real-options and precautionary savings channels, we expect that particularly IU alters the traditional interest rate channel of monetary policy via the term structure of interest rates. Yet, the literature, is far from consensual with regard to the quantitative importance of the expectation and term premium components of bond yields, nor with regard to the role of term premia in qualitative, i.e., directional terms ([Rudebusch and Swanson, 2012](#); [Crump et al., 2016](#); [Kliem and Meyer-Gohde, 2017](#)). However, both theoretical models ([Rudebusch and Swanson, 2008](#); [Ehling et al., 2018](#)) and empirical studies ([D'Amico and Orphanides, 2014](#); [Wright, 2011](#); [Ehling et al., 2018](#)), have provided decisive evidence suggesting that IU is a key determinant for (nominal) term premia of fixed income securities and the respective term structure. Moreover, IU clearly leads investors to demand higher term premia.<sup>7</sup> Therefore, it is conceivable that, first, the transmission of a monetary easing to the long end of the yield curve is attenuated due to higher term premia in periods of high IU and that, second, the effect of IU is asymmetric for expansionary and contractionary monetary policy shocks ([Tillmann, 2017](#)). In addition, as argued by [Gürkaynak et al. \(2005\)](#) long-term inflation expectations might lack sufficient anchoring in periods of elevated IU such that inflation expectations could depend on IU.<sup>8</sup> Overall, we expect (i) that the transmission channels are asymmetrically affected by IU, and (ii) that IU shows specific quantitative and possibly qualitative effects in comparison with other uncertainty categories as, e.g., stock market volatility ([Henzel and Rengel, 2017](#)).

### 2.3 Empirical evidence on uncertainty and monetary policy effectiveness

The empirical literature on the link between fluctuations in economic uncertainty and monetary policy transmission is still scant. We are aware of only five studies focusing on this subject. [Aastveit et al. \(2017\)](#) and [Pellegrino \(2017\)](#) employ interacted VARs to model the nonlinear relation between uncertainty and the effects of monetary policy on real economic indicators in the US (see also [Pellegrino, 2018](#), for the Euro area). Both studies assess the regime dependence of monetary policy for tranquil and uncertain periods and highlight that monetary policy shocks exert a considerably weaker effect on investment and GDP during uncertain periods. While the evidence supports a real options effect, the price-setting explanation put forth by [Vavra \(2013\)](#) is not corroborated, i.e., the responses of prices do not vary with uncertainty. Moreover, [Pellegrino \(2017\)](#) documents that reduced monetary policy effectiveness is linked to precautionary savings and real option effects for households. [Castelnuovo and Pellegrino \(2018\)](#) investigate the real effects of regime-contingent monetary policy by means of a medium-scale threshold VAR (T-VAR), and find, contrary to [Pellegrino \(2017\)](#) and [Aastveit et al. \(2017\)](#), considerable differences in price flexibility between tranquil and uncertain states (see also [Bachmann et al., 2013](#)).

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<sup>6</sup>[Berument et al. \(2005\)](#) and [Fountas et al. \(2006\)](#) provide a review of the empirical and theoretical literature regarding the effects of IU on interest rates and economic performance.

<sup>7</sup>We emphasize that the model predictions and empirical results are based on long-term IU (see [Ball et al., 1990](#), for a discussion of IU effects at distinguished horizons).

<sup>8</sup>Recently, [Falck et al. \(2017\)](#) have provided evidence that the impact of monetary policy surprises on inflation expectations depends of the degree of inflation disagreement among agents.



Interestingly, all reviewed studies adopt a recursive (or almost recursive) identification scheme to identify the monetary policy shocks.<sup>9</sup> As a result, the notorious “prize puzzle” is observed in the above papers, casting doubts on the employed identification schemes. Moreover, the endogeneity of uncertainty is an issue that has received attention. The results from all four studies are robust to accounting for potential effects of monetary policy shocks on uncertainty.

A study that stands out from those reviewed so far is [Eickmeier et al. \(2016\)](#). Based on a medium-scale T-VAR model, they explore the role of the financial sector for asymmetries in monetary policy effectiveness. Identification is achieved by utilizing either the Gertler-Karadi series of FOMC policy announcements as external instruments or by imposing sign restrictions on the short-run responses. Apart from a considerably weaker policy transmission on real variables during uncertain periods, results of [Eickmeier et al. \(2016\)](#) underpin that the cost of credit (i.e. corporate bond spreads) is strongly state dependent. Accordingly, the financial accelerator weakens as a reflection of reduced risk-bearing capacities of financial intermediaries.

## 2.4 Monetary policy and inflation uncertainty in a nutshell – hypotheses

Given these ample theoretical and empirical hints, we formulate the following hypotheses, exemplarily for an expansionary monetary policy shock. Each hypothesis states i) the underlying theoretical mechanism linking IU and monetary policy transmission; ii) the transmission channel which is affected by this underlying mechanism, and iii) how it can be detected.

1. **Mechanism:** Wait-and-see effect of firms and households on investment, hiring and consumption; **Detection:** During elevated IU regime responses of investment, unemployment (and its components) and durable consumption are dampened. **Channel:** Intertemporal substitution and wealth effects, i.e., the traditional interest rate channel is impaired.
2. **Mechanism:** Precautionary savings of households and firms. **Detection:** In contrast to real-options, regime differences arise also in nondurable consumption expenditures. **Channel:** Affects the interest rate channel directly.
3. **Mechanism:** Enhanced price flexibility and dispersion at high uncertainty; **Detection:** Stronger responses of prices in the high IU regime. **Channel:** Reduced *real* effectiveness of the interest rate channel.
4. **Mechanism:** Investors demand a higher inflation risk premium. Distorted effect of the policy rate on the term-structure of interest rate, i.e., **Detection:** Term premia increase in the high IU regime, and lead to a weaker reaction of long-term sovereign bond yields. **Channel:** Transmission via the term structure;
5. **Mechanism:** IU affects the liabilities of firms by inducing higher down-side risks. **Detection:** Corporate bond risk spreads decrease less strongly during uncertain times. **Channel:** Firm balance sheet channel.
6. **Mechanism:** Financial institutions reduce their leverage and therewith increase their net worth/capital to shield against adverse shocks. **Detection:** i) Condition (5) is fulfilled; ii) Measures of financial intermediaries’ leverage and risk-bearing capacity increase less in high

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<sup>9</sup>[Castelnuovo and Pellegrino \(2018\)](#) adopt a mixture of short and long-run restrictions. However, zero restrictions are imposed on the instantaneous effects of monetary policy shocks on aggregate quantities and prices.

IU regime. Price and quantity effect expected. **Channel:** Bank lending and/or risk-taking channel.

Even though we arrived at a number of clear hypotheses, a caveat arises from the difficulty to disentangle the sources of certain responses (or their state dependence), because they could be attributed to several mechanisms simultaneously. This is particularly pronounced for the discrimination of channels that are related to financial markets. For instance, regime dependence of credit spreads responses might reflect a weakening of the bank lending and/or the firm balance sheet channel.<sup>10</sup> Likewise, regime-dependent fluctuations in mortgage loans probably encompass demand and supply effects, which are difficult to separate from each other. Put differently, the empirical model might just not be sufficiently structural to sharply distinguish between some of the channels, since it is based on highly aggregated data.

### 3 The regime dependent structural factor model

This section provides a detailed account of our empirical framework. We first outline the ST-FAVAR model, and address, second, the estimation of common factors and the identification of monetary policy shocks. Finally, we introduce the employed transition variable. Detailed accounts of the estimation and identification methods are provided in Appendix C.

#### 3.1 The structural smooth-transition FAVAR model

##### 3.1.1 ST-FAVAR in reduced form

The ST-FAVAR model introduces nonlinear relations among the common factors into the FAVAR model of [Bernanke et al. \(2005\)](#). The common factors follow a ST-VAR which allows for two IU regimes ((i.e., low and high, [Auerbach and Gorodnichenko, 2012](#))).<sup>11</sup> We favor the ST-VAR model over alternative approaches - such as the threshold VAR model or a sample splitting strategy - for several reasons. First, although uncertainty might change abruptly we consider instant transitions as overly restrictive even more so since the ST-VAR comprises the threshold VAR as a limiting case. Second, owing to its probabilistic structure a smooth transition specification exploits a larger number of observations for the estimation of regime specific dynamics than sample split VAR models ([Caggiano et al., 2014](#)). Third, utilizing a dynamic factor model rather than only a small-scale ST-VAR model enables us to model the dynamic propagation of monetary policy surprises for a large number of variables. Furthermore, issues of non-fundamentalness might arise from deficient information sets in VAR analysis ([Forni et al., 2009](#); [Forni and Gambetti, 2010](#)) but non-fundamental shocks are unlikely encountered in factor models.

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<sup>10</sup>The balance-sheet and lending channels could be disentangled by looking at the GZ spread and the excess bond premium developed by [Gilchrist and Zakrajšek \(2012\)](#).

<sup>11</sup>The textbook by [Terasvirta et al. \(2010\)](#) provides an excellent treatment of smooth transition vector models.

In state-space representation the ST-FAVAR model consists of the following set of equations:

$$\mathbf{X}_t = \mathbf{\Lambda} \mathbf{C}_t + \mathbf{e}_t, \quad (1)$$

$$\mathbf{C}_t = G(z_{t-1}) \mathbf{A}_l(L) \mathbf{C}_{t-1} + (1 - G(z_{t-1})) \mathbf{A}_h(L) \mathbf{C}_{t-1} + \mathbf{u}_t, \quad (2)$$

$$\mathbf{u}_t \sim N(\mathbf{0}, \mathbf{\Omega}_t), \quad (3)$$

$$\mathbf{\Omega}_t = G(z_{t-1}) \mathbf{\Omega}_l + (1 - G(z_{t-1})) \mathbf{\Omega}_h, \quad (4)$$

$$\mathbf{e}_t \sim N(\mathbf{0}, \mathbf{\Sigma}), \quad (5)$$

$$G(\gamma, c, z_t) = \frac{\exp(-\gamma z_t - c)}{1 + \exp(-\gamma z_t - c)}, \gamma > 0, \text{Var}[z_t] = 1, E[z_t] = 0. \quad (6)$$

Constituting a static factor model, the measurement equation (1) relates the economic and financial indicators collected in the  $N$ -dimensional vector  $\mathbf{X}_t$  to the  $K$ -dimensional common factors  $\mathbf{C}_t$  and the vector of idiosyncratic components  $\mathbf{e}_t$ . The common component  $\mathbf{C}_t = [Y_t \ F_t]$  includes  $r$  observed ( $Y_t$ ) and  $K - r$  unobserved factors ( $F_t$ ), with  $\mathbf{\Lambda} = [\beta \ \Lambda^f]$  denoting a  $N \times K$  matrix of respective factor loadings  $\beta$  and  $\Lambda^f$ . Idiosyncratic components  $\mathbf{e}_t$  are uncorrelated with  $\mathbf{C}_t$  at all leads and lags and with the factor innovations  $\mathbf{u}_t$ , but are allowed to be weakly cross- and autocorrelated. Hence, we consider an approximate factor model.<sup>12</sup>

The nonlinear factor dynamics follow an ST-VAR process of order  $p$  (ST-VAR( $p$ )) specified by the transition equations (2) – (4). Lag polynomials,  $\mathbf{A}_\bullet(L)$  in (2) capture system dynamics during (extremes of) low and high IU regimes, i.e.,  $\bullet = l, h$ . The regime allocation is governed by a logistic transition function  $G(\gamma, c, z_t)$  which depends on i) the smoothness parameter  $\gamma$ ; ii) the location parameter  $c$ ; and iii) the transition variable  $z_t$  which is a measure of IU (see Section 3.3). The transition variable enters the transition function with a lag of one period to avoid contemporaneous feedback and, moreover, it is standardized to ensure that  $\gamma$  is scale invariant. The transition function takes values between 0 (high IU) and 1 (low IU). Hence the ST-FAVAR model is a convex combination of two linear FAVARs and nests the linear FAVAR for  $\gamma = 0$ . According to (3) and (4), the  $K$ -dimensional reduced-form residuals  $\mathbf{u}_t$  are heteroskedastic Gaussian with mean zero and time-varying covariance matrix  $\mathbf{\Omega}_t$ . Comparing our specification with the only two studies that adopt a similar framework, we note that our model differs from Popp and Zhang (2016) who allow for state-contingent factor loadings ( $\mathbf{\Lambda}_\bullet$ ), whereas we assume constant loadings. The model of Dahlhaus (2017) restricts regime dependence to the autoregressive dynamics.

### 3.1.2 The structural ST-FAVAR model

Thus far we have outlined the reduced form model, while we are mainly interested in inference based on causal effects obtained from the structural representation of the ST-FAVAR. To derive the structural impulse response functions (IRFs) we rewrite (2) as,

$$\mathbf{\Pi}_t(L) \mathbf{C}_t = \mathbf{u}_t, \text{ where} \quad (7)$$

$$\mathbf{\Pi}_t(L) = [\mathbf{I} - G(z_{t-1}) \mathbf{A}_l(L)L - (1 - G(z_{t-1})) \mathbf{A}_h(L)L]. \quad (8)$$

From (7) the structural MA representation obtains as

$$\mathbf{C}_t = \mathbf{\Phi}_t(L) \mathbf{B}_t \boldsymbol{\epsilon}_t, \quad (9)$$

<sup>12</sup>As in Bernanke et al. (2005) the observable and unobservable factors might be (contemporaneously) correlated, but  $\text{Cov}[F_t]$  is a diagonal matrix by assumption.

where  $\Phi_t(L) = \Pi_t(L)^{-1}$ . The matrix  $B_t$  which obeys  $\Omega_t = B_t B_t'$  summarizes the time-varying instantaneous relations among structural shocks  $\epsilon_t = B_t^{-1} u_t$  which are distributed as  $WN(0, I_K) \forall t$ . The analysis in this paper is based on dynamic responses assumed to be linear conditional on the IU regime, that is,  $G(\cdot) = 0$  or  $G(\cdot) = 1$  or on linear combinations of both corresponding to a certain quantile of the transition variable  $z_t$  (i.e. IU). This implies that the system is interpreted to stay in one regime for a sufficiently long period. The regime dependent impact matrices satisfy  $B_\bullet B_\bullet' = \Omega_\bullet$ ,  $\bullet = l, h$ , and could be obtained, in a first step, as Cholesky factors of  $\Omega_\bullet$ . Substituting (9) in (1) we obtain the regime-specific IRFs associated with  $X_t$  (disregarding the idiosyncratic component  $e_t$ ) given by  $\Psi_j^\bullet = \Lambda \sum_{i=0}^j \phi_i^\bullet B_\bullet$ ,  $\bullet = l, h$ , where the  $\phi_i^\bullet$  are the MA coefficient matrices corresponding to  $A_\bullet(L)$  for the horizons  $j = 1, \dots, J$ . We can treat the IRFs as conditionally linear, since the transition variable ( $z_t$ ) is not included in the panel of modelled variables. This precludes any feedback from interest rate shocks to IU.

## 3.2 Estimation and identification

To estimate the ST-FAVAR model, we adopt the computationally convenient two-step principal components (PC) approach put forth by [Bernanke et al. \(2005\)](#); [Stock and Watson \(2002\)](#) and [Boivin et al. \(2009\)](#). Accordingly, we first estimate the unobserved factors  $F_t$  nonparametrically by means of principal component analysis (PCA), and subsequently evaluate the ST-VAR conditional on the estimated factors. The two-step PC estimator does not build upon strong distributional assumptions on the factor innovations as it is the case for the two-step ML approach of [Bai et al. \(2016\)](#) or the simultaneous estimation of observation and transition equations by means of Bayesian methods ([Bernanke et al., 2005](#)).<sup>13</sup> [Bernanke et al. \(2005\)](#) document only minor performance leads of the computationally considerably more demanding Bayesian estimation by means of Gibbs sampling.<sup>14</sup>

### 3.2.1 Factor estimation

As elucidated in [Ahmadi and Uhlig \(2015\)](#), it is of utmost importance to include a price index in the VAR model, even in the case of a FAVAR, in order to pin down the exogenous shocks to the policy instrument correctly and to avoid a prize puzzle. Hence, the vector of observable factors comprises the federal funds rate (ffr) and a price index, i.e.,  $Y_t' = [ffr_t, p_t]$ . Based on [Hwang \(2009\)](#), we adopt a two-step approach that avoids an iterative solution. In a first step the estimated coefficients  $\hat{\beta}$  for the observable factors  $Y_t$  are obtained from a regression of  $Y_t$  on  $X_t$ . In a second step the estimate of  $F_t$  is obtained as the first  $K - r$  principal components of the first step residuals  $\hat{U}_t = X_t - \hat{\beta} Y_t$ . Hence, the factor estimate is given by  $\hat{F} = \hat{U}' \hat{\Lambda}^f$ , where  $\hat{\Lambda}^f$  is the matrix of eigenvectors corresponding to the  $K - r$  largest eigenvalues of  $\hat{U}' \hat{U}$ . This ensures that the unobserved factors are identified against any rotational indeterminacy (see [Killian and Lütkepohl, 2017](#), Ch. 16 for alternative identification conditions).

Once the common component has been obtained in the first step, the ST-VAR model describing the factor dynamics can be estimated. Owing to the nonlinear nature of the model, we follow [Auerbach and Gorodnichenko \(2012\)](#) and utilize the MCMC approach proposed in [Chernozhukov](#)

<sup>13</sup>Although the two step ML estimator promises efficiency under correct model specification, it could lack robustness if the distributional assumptions are violated.

<sup>14</sup>[Stock and Watson \(2002\)](#) prove that the PC estimator consistently recovers the space spanned by the factors for large  $N$  if the number of PC factors is at least as large as the true number of factors. Moreover, the problem of generated regressors is negligible if  $N$  is large in comparison with  $T$  ([Bai and Ng, 2006](#)).

and Hong (2003). The location parameter  $c$  is set to zero and an initial estimate of the smoothness parameter  $\gamma$  is obtained from a preliminary ML estimation and then jointly estimated with the remaining parameters in the MCMC sampling procedure.<sup>15</sup> Estimation of  $\Omega_l, \Omega_h, A_l(L), A_h(L)$  is based on 100.000 draws from the Metropolis-Hastings sampler where 40.000 draws are discarded as burn-in.

### 3.2.2 Identification of monetary policy shocks

To identify the exogenous monetary policy shock, we adopt a sign restriction approach, which offers a number of advantages compared with the imposition of short-run zero restrictions (Faust, 1998; Uhlig, 2005). The main advantage is that this identification framework seems particularly suited in the light of ample evidence for contemporaneous simultaneity between monetary policy and financial variables (see, e.g. Björnland and Leitemo, 2009; Lütkepohl and Netšunajev, 2014; Beckers and Bernoth, 2016). A recursive setting in the vein of Bernanke et al. (2005) prevents any instantaneous reaction of monetary policy makers to high-frequency information contained in financial indicators, such as stock prices and bond yields. This seems highly unrealistic, given that these variables are seen to bear information about the stance of monetary policy itself (Rigobon and Sack, 2003). Moreover, in contrast to timing restrictions on the instantaneous relations of structural shocks, sets of sign restrictions usually obtain from predictions of DSGE models and hence benefit from theoretical underpinnings (Canova and Pina, 2005). By assumption, an expansionary monetary surprise has a nonpositive impact on the federal funds rate and two additional short term interest rates (3 and 6 months treasury bond yields), while consumer prices (CPI all items) and output (industrial production) respond nonnegatively. Moreover, the monetary policy shock exerts a nonnegative effect on the monetary base, the monetary aggregate M1 and total reserves. Following Uhlig (2005), all restrictions are imposed on impact and the following six months.

## 3.3 Inflation uncertainty as transition variable

### 3.3.1 Predictive horizon and IU assessment

In our ST-FAVAR model the transition indicator  $z_t$  is an IU metric. An important choice that has to be made in order to evaluate the appropriateness of an IU measure concerns the most suitable predictive horizon. A-priori long-term IU matters more for intertemporal decisions while short-term IU is more important for intratemporal decisions. Moreover, long-term IU is more closely related to term premia and monetary policy targets (Ball et al., 1990; Evans, 1991; Caporale et al., 2010). Therefore, all IU measures employed in this work refer to a forecast horizon of 12 months which we consider as long-term IU. Besides the predictive horizon one has to decide on the broad method of measurement. Three main families of methods have been employed to quantify IU (see Grimme et al., 2014, for a summary): survey-based, forecast-based (e.g., VAR models) and model-based measures (e.g., (G)ARCH).<sup>16</sup> We choose as our baseline IU indicator the average of

<sup>15</sup>Treating the ML estimate of the smoothness parameter  $\gamma$  as fixed in the subsequent MCMC procedure does not change results considerably.

<sup>16</sup>From the literature on IU metrics it can be stated that i) disagreement and aggregate uncertainty are theoretically and empirically related; ii) uncertainty partly comprises disagreement; iii) both concepts may nevertheless be related to distinct macroeconomic effects; iv) disagreement and uncertainty measures may diverge considerably especially during most uncertain times (see Lahiri and Sheng, 2010; Zarnowitz and Lambros, 1987; Glas and Hartmann, 2016). Moreover, all methods are subject to specific shortcomings and constitute an additional source of heterogeneity.

individual predictive standard deviations (denoted henceforth by AVEIU) which has been obtained from 14 alternative inflation prediction models (Hartmann et al., 2017). As such it is an ex-ante measure that mimics individual IU based on a set of density forecasts. Accordingly, it accounts to some extent for disagreement and model uncertainty. Specifically, it has received empirical support based on diverse forecasting performance criteria Hartmann et al. (2017).<sup>17</sup>

### 3.3.2 The relationship between IU, the business cycle and other uncertainty indicators

There is by and large consensus in the pertinent literature that many uncertainty indicators surge rather abruptly during recessions and that these exhibit a strong comovement with real activity variables (Bloom, 2014; Castelnuovo et al., 2017). Hence, using IU as a transition variable might be confounded by other sources of regime dependence along the business cycle. For illustration purposes, panel a) in Figure 1 depicts NBER recession dates joint with three distinct IU measures, namely i) the average of individual predictive standard deviations AVEIU; ii) the model-based variance forecasts of permanent shocks to inflation as implied by the stochastic volatility model of Stock and Watson (2007) (UCSV); and iii) the survey-based statistic of Binder (2017) (BINDER).<sup>18</sup>

The main insights from Figure 1 can be summarized as follows. IU metrics take their highest values during the high-inflation era at the beginning of the sample ('Great Inflation' and 'Volcker disinflation') and during or after the Great Recession. Furthermore, only the survey-based measure seems to spike abruptly during recessions, while the remaining two indices (AVEIU and UCSV) lack a close relation to recessions. Rather, they often even decrease during bad times. Hence, there is considerable heterogeneity in IU statistics and the consensual insights – i.e., surge in recessions and close co-movement with real activity indicators – do not apply straightforwardly to all types of uncertainty metrics.

Next we consider a selection of commonly employed financial and general macro-uncertainty indicators displayed in panel b) of Figure 1, namely the general macro-uncertainty factor proposed by Jurado et al. (2015) (denoted JLN12), the option implied volatility index (VIX) and the GZ spread proposed by Gilchrist and Zakrajšek (2012).<sup>19</sup> We observe the typical characteristics shared by many uncertainty measures reported in the literature, that is, a close relation to the business cycle and sudden surges during recessions. All indices attain their maximum values during the Great Recession. Based on eyeballing it appears that the JLN index is closest to the IU measures displayed in panel a) of Figure 1.

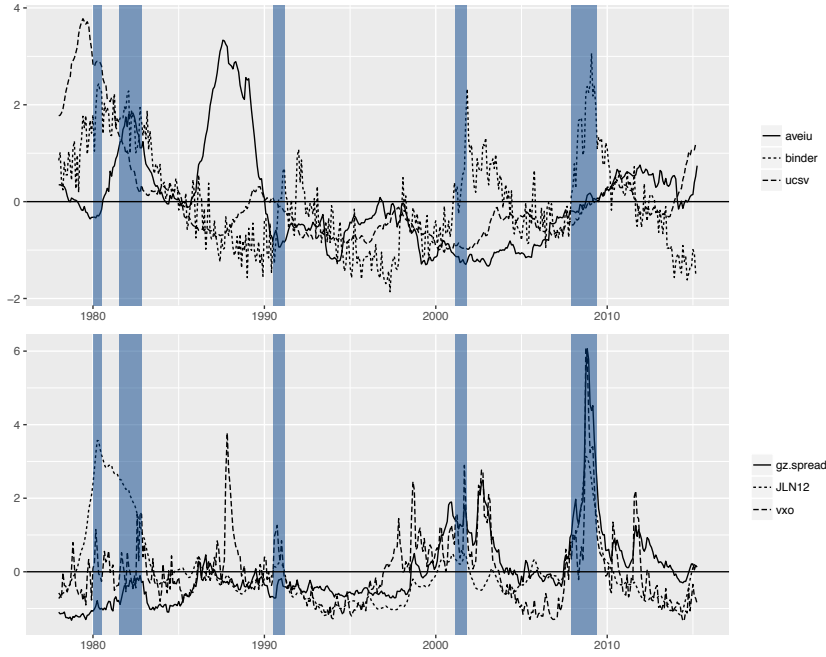
To get a quantitative picture of the degrees of similarity between IU and other uncertainty measures, Table 1 documents respective correlations. Except for survey-based IU statistics (BINDER) (correlation of 0.75 with JLN12) the correlation with the other uncertainty categories is at most moderate and strongest for JLN12 (up to 0.52). The low or even negative correlations of the baseline IU measure (AVEIU) with the VIX and the GZ spread corroborates that we do not capture financial stress related sources of nonlinearity in monetary policy transmission.<sup>20</sup>

<sup>17</sup>More specifically, Hartmann et al. (2017) report that AVEIU outperforms all other common model- and forecast-based IU proxies in terms of forecasting performance with respect to ex-post IU and different loss functions (e.g., RMSE, directional accuracy), in almost all scenarios that the authors investigate.

<sup>18</sup>More detailed descriptions of these IU statistics are provided in Appendix B.

<sup>19</sup>The GZ spread is an unweighted cross-sectional average of around 3500 credit spreads in the respective month. JLN12 is the common variation in the unforecastable component (for a forecasting horizon of 12 months) of a large number of economic indicators. For additional information regarding the computation and exact sources of these uncertainty indices the reader is referred to the data appendix (Appendix B).

<sup>20</sup>The finding of a negative correlation coefficient between, for instance, the VIX index and the UCSV based IU measure



**Figure 1:** Panel a) shows three different IU indices. Panel b) displays three different uncertainty measures. Shaded columns: NBER recessions.

**Table 1:** Correlations between IU and other uncertainty proxies. Lower triangular elements show Pearson's correlation coefficients and upper triangular elements depict Spearman's rank correlation coefficients.

	AVEIU	UCSV	BINDER	JLN12	VXO	GZ spread
AVEIU	1.0000	0.5500	0.0900	0.1500	0.0700	-0.1400
UCSV	0.1800	1.0000	0.4100	0.3400	0.0100	-0.3200
BINDER	0.0200	0.4900	1.0000	0.6200	0.3100	0.0500
JLN12	0.1600	0.5200	0.7500	1.0000	0.4200	0.1400
VXO	0.1000	-0.0600	0.3500	0.3800	1.0000	0.3500
GZ spread	-0.1400	-0.3500	0.2400	0.2300	0.6400	1.0000

In conclusion, we are confident that at least estimation results based on AVEIU are unlikely driven by recession related sources of regime dependence. Moreover, empirical results on the role of IU for monetary policy transmission feature only minor confounding with effects originating in other types of uncertainty. Finally, it is worth noticing, that the correlation between IU proxies is generally rather low [0.02, 0.49] confirming the conjecture made in Section 3.3.1 stating that IU metrics may reflect distinct dimensions of uncertainty and typically diverge considerably. This emphasizes the importance of carefully choosing a suitable IU approximation in order to cover the particular dimensions of uncertainty under scrutiny.

might partly be attributed to the discrepancy in the horizon of the VIX.

## 4 Data, specification and empirical results

This section first describes our dataset and specification details. Second, we show regime dependent dynamic responses resulting from the baseline ST-FAVAR model, highlight evidential underpinnings of the hypotheses raised in Section 2 and examine the robustness of our findings. Finally, we discuss results from a structural ST-FAVAR model applied to an extended panel of time series that is observed with lower time resolution (i.e., quarterly data).

### 4.1 Data and specification

In a baseline setting we analyse an extended version of the FRED-MD database maintained by Micheal W. McCracken and Serena Ng (McCracken and Ng, 2016) that consists of 148 monthly time series covering the period 1977M1 to 2015M12.<sup>21</sup> The panel comprises a broad range of time series relating to real activity, (e.g., output, income and labor market) and various price indices and financial variables (such as interest rates, monetary aggregates and credit related indicators). The panel is standardized and the series are transformed to achieve stationarity prior to factor extraction (see Appendix B for a list of variables and details on data transformations). Based on marginal  $R^2$  statistics, we extract four principal component factors for the baseline model.<sup>22</sup> The four unobserved factors explain 33.6% of the variance in the panel and increasing the number of factors adds little information. For instance, the marginal contribution of the fifth factor amounts to 4.3% of the panel variance.<sup>23</sup> Figure 3 shows the unobservable factors jointly with the two observed factors, i.e., the federal funds rate and the CPI. Tables 2 and 3 document sizeable variance shares explained by the common factors (including, ffr and CPI) for all groups of variables. The model is estimated with a lag order of  $p = 2$  as suggested by the BIC applied to the linear FAVAR. Another important specification step is to test if our nonlinear model is supported by the data. In case that the data generating process is linear the ST-FAVAR lacks identification. The nonlinear model receives support from information criteria.<sup>24</sup> Our baseline IU measure  $AVEIU$  (transition indicator  $z_t$ ) is depicted jointly with the estimated transition function  $G(z_t)$  in Figure 2. Apparently, values of  $G(z_t)$  near zero (high levels of IU) are associated with the beginning of the Volcker Fed chairmanship, the subsequent disinflation period around 1982 and most of the second half of the 1980s. Almost coinciding with the Great Moderation era  $G(z_t)$  is close to unity (states of low IU) between 1989 and 2008. Compared with the minimum in the 1980s, the regime allocations during and after the Great Recession are markedly larger, but still considerably below the inflection point.

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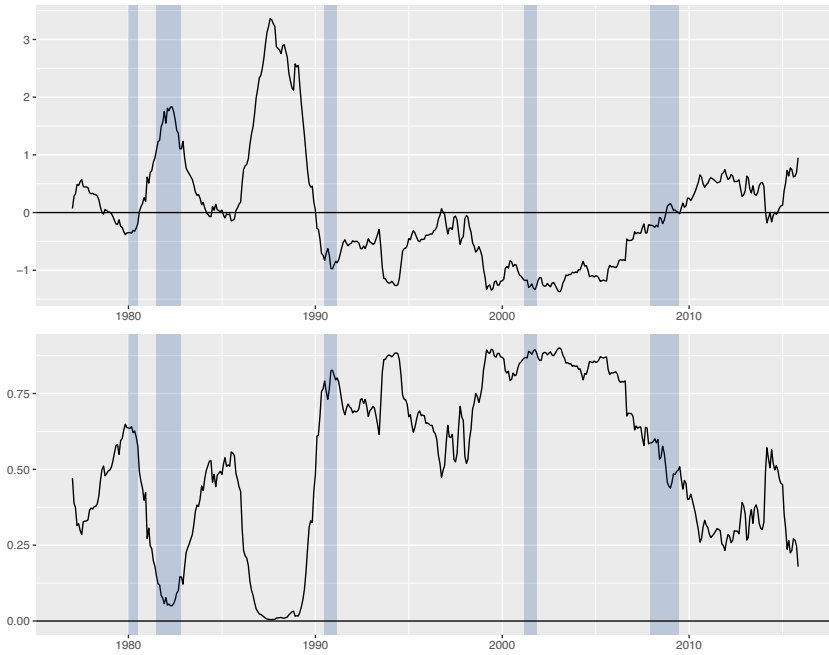
<sup>21</sup>The database is available on John W. McCracken's web page <https://research.stlouisfed.org/econ/mccracken/fred-databases/>.

<sup>22</sup>We also considered the Bai and Ng criterion (Bai and Ng, 2006) which obtains an optimal dimension of eight factors (in addition to the two observable variables). Favouring a somehow overspecified model order, the criterion does not aim at the number of factors to be included in a (non)linear VAR.

<sup>23</sup>We report results for a five factor ST-FAVAR in Section 4.3.

<sup>24</sup>The AIC for the linear (nonlinear) model amounts to -16.59 (-17.02) and the BIC obtains -16.49 for the linear FAVAR and -16.66 for the ST-FAVAR.





**Figure 2:** The upper chart shows the standardized IU proxy (the average of individual predictive standard deviations) and the lower chart displays the transition function  $G(z_t, \gamma)$ . Shaded columns: NBER recessions.

## 4.2 Baseline results

Figures 4 to 7 display the impulse responses of a large number of variables ensuing from an expansionary monetary policy shock which is normalized to reduce the federal funds rate by 25 basis points. As representatives for regimes of low and high IU we show IRFs for the 5th and 95th percentiles of the standardized IU measure. These percentiles correspond to values of the transition function  $G(z_t)$  of .881 (lower percentile) and .026 (upper). In addition, Figure 8 shows the difference between the IRFs conditional on low and high IU for a selection of key variables.

From Figure 4 we note first that the monetary policy shock is associated with a long lasting fall in the federal funds rate (ffr) only in the high IU regime, while it fades out rather quickly conditional on states of low IU and returns to its pre-shock level after about one year. Next, we describe the general effects of the monetary policy shock which are common to both regimes. The unexpected decrease in the federal funds rate invokes a pronounced and persistent increase in industrial production (IP), employment, hours, the help-wanted index and capital utilization. Figure 5 displays the responses of several price indices as well as monetary and credit aggregates. We observe expansionary effects on consumer prices (CPI) in both regimes and over all horizons. Furthermore, the monetary policy shock triggers a strong rise in monetary aggregates and total reserves. Short term (TB3MTHS, TB6MTHS, GB1Y see Figure 7) and long term interest rates (GB5Y, GB10Y) decrease as expected and policy rate transmission weakens with increasing maturity. Throughout, confidence bands surrounding the response profiles shown for the regime of elevated IU are more precise in comparison with results conditioning on states of low IU. Following [Castelnuovo and Pellegrino \(2018\)](#), this might be attributable to more data variation during uncertain times and to the normalization of the monetary policy shock. The standard deviation of the interest rate is about six times as high in the extreme IU regime ( $G(z) = 0$ ) compared with the low IU regime ( $G(z) = 1$ ). In the remainder of this section we focus on state dependent monetary policy transmission and

effectiveness first for variables related with real economic activity and subsequently investigate the nominal and financial side.

#### **4.2.1 Regime dependent real economic effects of monetary policy**

In the following we collect empirical evidence on the hypotheses 1 and 2 (cf. Section 2.4). Regarding potential state dependence of the dynamic responses of real economic indicators (Figures 4 and 8), we make three general observations. First, we detect moderate differentials in monetary policy effectiveness for the majority of time series that are related to production, income and employment. Second, the persistence profiles of impulse responses are of weak regime dependence for most variables. Third, state-specific effects tend to arise in the short- and medium-run (i.e., three months up to three years), while on impact and at long horizons response magnitudes are similar for distinguished states of IU. To give an impression of the loss in effectiveness of monetary policy, we consult Figures 4 and 8 and note that the median peak effects for industrial production and its subindices durable and business equipment is between 30 - 60% at one year after the policy shock has hit the economy. For labour market indicators it is even larger: The effect on unemployment is halved at around 15 months after the impact during high IU. This loss is statistically significant considering parameter and model uncertainty.

Besides these general characteristics we observe several more specific results. Comparing the responses of industrial production of durable and nondurable goods, we note that these are qualitatively similar but the regime dependence of IP of nondurables is even more pronounced than that of durables. As a reflection of their stronger interest-rate sensitivity, the effects on the production of durables are more pronounced in comparison with effects on the production of nondurables. Indicating the presence of both real options and precautionary savings channels, this comparative evidence is in line with hypotheses 1 and 2 and with findings in Pellegrino (2017). In contrast, if there were only the real options channel active, one would expect that the effects on nondurable consumption lack regime-dependence. Both conclusions should be taken cautiously, however, since we are considering only production indices and not the consumption expenditure counterpart.

Turning to labour market indicators, employment, the help-wanted index and weekly average hours increase sluggishly but persistently in response to the expansionary monetary policy shock in both regimes. Unemployment drops for two years before it stabilizes at a total reduction by 0.11 percentage points. While the differences in the stimulative effects on unemployment, employment and utilization between low and high IU states are quite pronounced, the regime dependence in hours and the help-wanted index are weaker in relative terms. Hence, the dynamic propagation of monetary surprises via the labour market is consistent with hypothesis 1, emphasizing the labour market aspects of the wait-and-see effect.

In summary, the investigation of real activity related IRFs suggests that an unexpected monetary easing invokes weaker effects conditional on elevated levels of IU in comparison with macroeconomic states featuring low IU.

#### **4.2.2 Prices, liquidity and asset markets**

In this section we discuss empirical evidence on the hypothesis 3, monetary aggregates and the interaction of IU with the asset price channel of monetary policy. The dynamic responses of several alternative measures of aggregate prices are depicted in Figure 5, and reveal major effect

differentials for monetary policy transmission across IU regimes only for the aggregate CPI and not for the subindices. The CPI (all items) reacts considerably stronger during times of low IU, while its response is virtually zero at high levels of IU. For all other subindices, except for the CPI medical care and CPI apparel regime dependencies do not arise and respective responses share the same qualitative characteristics, i.e., after a significant short-term effect, responses are statistically indistinguishable from zero. We note that the dynamic reactions of the CPI and its subindices are not directly comparable, since the former depicts level responses and the latter are inflation rate responses (this is due to subindices entering the panel in second differences, while the CPI as observable factor enters in levels). Hence, our results do not provide evidence in favour of hypothesis 3 which refers to enhanced price flexibility as a threat to monetary policy transmission Vavra (2013). Combining the findings that precautionary savings of households likely play an important role in altering monetary policy transmission and that prices are more sticky during high IU our results corroborate the theoretical mechanism put forth in Basu and Bundick (2017).

Turning to the responses of monetary aggregates (Figure 5), we note that the monetary policy shock triggers a pronounced and persistent surge in the monetary base (M0), M2 and in total reserves held by monetary institutions in both regimes.<sup>25</sup> Moreover, we detect mild regime dependence since the liquidity effect is stronger during episodes of low IU. As Figure 8 shows exemplarily for M0 and M2, the differences between responses are significant over all horizons, .

Relatedly, the reactions of aggregated loan provision appear not as uniform as the reactions of monetary aggregates. Lending to the business and real estate sectors increases after zero or negative short term effects, whereas lending to consumers declines persistently over all horizons. Regime differences arise for all three loan aggregates. Business and real estate loans react more pronounced (positively) during uncertain times, while consumer loans respond stronger (negatively) conditional on states of high IU. Given the close link between mortgage loans, house prices and real estate investment, the stronger response of real estate loans in times of high IU suggests to further investigate if this corresponds to increased activity in the real estate sector. The reaction of nationwide housing permits, housing starts, the Case-Shiller house price index and a REIT price index are, not only at odds with this conjecture (see Figure 6) but also at odds with conventional wisdom. For both regimes of IU a monetary easing has contractionary effects on housing starts and permits at least in the short run, while effects become statistically insignificant at longer horizons. Looking at real estate prices, we observe an imprecisely estimated negative and persistent response of the CS HPI and REIT index in both regimes with strongly overlapping confidence bands.

Finally, turning to stock markets, we detect slight differences between the regime dependent responses which are depicted in Figure 6. The reaction of the S&P 500, Dow Jones Industrial and its subindices as well as the Wilshire 5000 are very similar qualitatively as well as quantitatively. A monetary easing induces an immediate and persistent decrease in both regimes with a stronger reaction during low IU. These regime differences are, however, only statistically distinguishable for the very short run.

#### 4.2.3 IU, monetary policy and the term structure

Figure 7 shows responses of selected interest rates and term spreads, that allow for an investigation of hypothesis 4, which highlights the adverse role of more persistent or increased term premia for

<sup>25</sup>The presence of a liquidity effect is generally expected for monetary policy shocks (Christiano et al., 1999), and has been confirmed formerly (Bernanke et al., 2005; Eickmeier and Hofmann, 2013; Castelnuovo and Pellegrino, 2018).

monetary policy transmission along the term structure (Crump et al., 2016). At low levels of IU the dynamic propagation of an expansionary monetary policy impulse (25 bps) on the three- and six-months T-bill rates (TB3MS and TB6MS) is characterized by a 22 bps reduction after three months. After two years both rates approach a new steady-state roughly 10 bps below the pre-shock value. In contrast, conditional on states of elevated IU, the short-term interest rates fall by only 10 bps after three months and overshoot after about 15 months to stabilize at a 5 bps higher steady-state. Interestingly the peak response of the one-year rate is stronger than those of the former rates (25 bps after three months in the low IU regime). A similar finding has been reported by Gertler and Karadi (2015) for the ffr and the one-year government bond rate. The response patterns for five- (GB5Y) and ten-year (GB10Y) government bond rates at low levels of IU are characterized by similar persistent, almost ‘linear’ trajectories with a steady-state reduction of 20 and 15 bps for the five- and ten-year rates, respectively. At high levels of IU the responses mirror those of the low IU regime but shifted towards the zero line, with peak (long-term) effects of roughly 12 (5) bps for the five- and 9 (4) bps for the ten-year rate. The IRF differences depicted in Figure 8 show clearly that, first, there is considerable, and statistically significant regime dependence, with monetary policy transmission being weakened at high levels of IU. Second, our evidence does not point to increasing impairment of monetary policy transmission as the maturity lengthens. Next, we consider briefly the term spreads between the long-term rates and the ffr (GB5FFR, GB10FFR) which decline only upon impact in both IU regimes, and increase in the medium run to ultimately stabilize around zero. Asymmetries in monetary policy transmission are noticeable, especially in the medium term. Besides the on-impact tightening in states of low IU, this response pattern accords with the expectation hypothesis of the term structure and confirms that we have not identified a “term spread shock” (Eickmeier and Hofmann, 2013). In summary, our analysis aligns with the expectation that monetary policy transmission weakens with increasing maturities, at least if we compare the medium and long term yields. In addition and in line with hypothesis 4 it provides evidence for the interaction of IU with the interest rate channel.

The presumption of a weakened interest rate transmission channel at high levels of IU rests on theoretical and empirical results which center around IU induced increases of term premia (Wright, 2011; Tillmann, 2017). We explore this aspect further by analyzing term premia and expected average short term rates for one-, five- and ten-year sovereign bonds. These estimates are obtained from the term structure model of Adrian et al. (2013) and associated impulse responses are depicted in Figure 7.

The responses of the expectation component reflect almost one-to-one the responses of corresponding rates, whereas the reactions of term premia are quantitatively negligible (between 1.5 and 2.5 bps peak effect). On impact, the term premium decreases in both IU regimes for one-year government bonds. For five- and ten-year bonds the impact effect of the monetary policy shock on the term premia is zero (positive) for states of high (low) IU and subsequently negative for both regimes, although insignificantly at low levels of IU. The regime differences are only statistically significant for the first few months.

The finding that expected rates respond sizably is in line with results reported in Crump et al. (2016) and Tillmann (2017). The dynamic responses of term premia, however, contrast with the theoretical rationale that long-term IU tends to render long-term bonds less attractive for investors due to the uncertainty about nominal and real future short-term rates, i.e., the mechanism behind hypothesis 4 and the empirical literature. Given the observation that (i) a monetary easing is less effectively transmitted already to short-term rates and (ii) that term premia are not significantly

affected, IU must interact with one or several of the following components of bond yields: expected real rates, inflation expectations or the real term premium.<sup>26</sup> While a further in-depth investigation of the exact link between IU and monetary policy transmission via the term structure is beyond the scope of this work, we point out two conjectures. First, our results are broadly consistent with model predictions and empirical results by [Ehling et al. \(2018\)](#) who demonstrate that disagreement between investors about future inflation could raise, inter alia, nominal and real yields above the potential effect of inflation disagreement on inflation expectations (see also [Falck et al., 2017](#), for the effect of inflation disagreement on inflation expectations). Importantly, it is found that inflation disagreement is mainly linked to nominal yields via its effect on real yields. To the extent that our uncertainty measure correlates with disagreement, this might explain our results. Second, given that regime dependence is present for the responses of short-term rates, one conjecture would be that our baseline IU measure is a yardstick for monetary policy uncertainty which might affect inflation expectations already at short maturities.

#### 4.2.4 IU-dependent financial accelerator

Next, we examine the empirical results with regard to hypotheses 5 and 6, i.e., the interaction of IU with the financial/credit channel of monetary policy. Relevant IRFs are depicted in Figure 7. Looking first at the response patterns of corporate bond yields (AAA and BAA) and corresponding spreads (AAA and BAA minus *ffr*) we observe that these are qualitatively almost identical to those for long-term government bonds. There are differences, however, in quantitative terms. Conditional on both high and low IU, the effects of the monetary policy shock on AAA and BAA corporate bond yields are smaller than the effects on the five- and ten-year bond rates. As IU induced disruption of transmission via the term structure is mirrored in the corporate bond market, the monetary authorities lose largely their ability to influence BAA rated bond rates (AAA rated bonds react only slightly in the very short run). Hence, this indicates that corporate finance conditions only improve considerably in the low IU regime. In order to gain insights into the role of financial market frictions we now consider the dynamic propagation of a monetary policy shock via credit and loan market spreads. Irrespective of the state of IU, an unanticipated monetary easing leads to a considerable contemporaneous drop in credit market spreads (GZ spread and the BAA - AAA spread). The effect is larger conditional on states of low IU (see also Figure 8). While conditional on low IU the response of the GZ spread spikes and turns positive at two months to drop again until returning to its pre-shock value, at elevated IU the response stays positive. In line with hypothesis 5 and model predictions of [Gilchrist et al. \(2014\)](#), the diagnosed regime dependence is particularly strong and longer-lasting for the (more informative) GZ spread, and points to a stronger reduction in the cost of credit due to balance-sheet effects.

Credit spreads are, however, not informative regarding the role of a potential lending and/or risk-taking channel, which might be responsible for parts of the reduction in bond financing costs.<sup>27</sup> Therefore, we investigate the excess bond premium (EBP) compiled by [Gilchrist and Zakrajšek \(2012\)](#) which is orthogonal to the expected default risk of issuers. The response of the EBP basically

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<sup>26</sup>Since the employed data on term premia and expectation components are obtained from the term structure model by [Adrian et al. \(2013\)](#) which does not consider a decomposition into real and nominal term premia, we cannot preclude that a term structure model which allows such a decomposition would result in a more prominent role of the inflation risk premium.

<sup>27</sup>A reduction in the credit spread constitutes a necessary condition but is not sufficient for the existence of a bank lending/capital or risk-taking channel.

mirrors the response of the GZ spread but regime dependencies are less pronounced and arise mainly in the medium run. Our results seem to confirm findings of [Gilchrist and Zakrajšek \(2012\)](#), i.e., large parts of the fluctuations of the GZ spreads can be attributed to the EBP. Consequently, it appears that the lending channel plays a relatively important role for the dampening effect of IU on monetary policy transmission via the external finance premium. Finally, loan market spreads (prime lending rate relative to the three and six months T-Bill rate, denoted BPR6TB) exhibit moderate regime differences, with an overall stronger reduction in bank lending costs at low levels of IU. Accordingly, our findings confirm results in [Eickmeier et al. \(2016\)](#) and support hypothesis 6 stating that IU weakens the bank lending channel and/or reduces the risk-bearing capacity of financial institutions.

Finally, we look at the effect of monetary policy on other uncertainty indices. Here, the stock market implied volatility (VXOCLSx in Figure 6) is considered as a representative quantification of general uncertainty. We observe that expansionary monetary policy shocks affect financial uncertainty considerably stronger during low IU. Uncertainty is first amplified for about two months with the same peak magnitude as the subsequent attenuation which lasts for about 15 months before fading out. When IU is within its 95% percentile, the response pattern is similar, but quantitatively negligible. Consequently, it is hard to draw clear conclusions on the effect of a monetary easing on uncertainty. However, considering the more persistent attenuating effect, our results seem to support findings in [Duca et al. \(2010\)](#) that lax monetary policy reduces both the risk aversion component of the VIX and the expected stock market volatility (see also [Pellegriano, 2017](#), for a similar result).

### 4.3 Alternative model specifications

In this section we discuss the results from two additional models based on the monthly panel of time series. First, we construct a “purged” version of our baseline IU indicator AVEIU similar to [Aastveit et al. \(2017\)](#). We regress AVEIU on the Federal Reserve bank of Chicago’s economic activity and financial conditions index and use resulting residuals to remove existing correlation between the IU indicator and the state of the economy.<sup>28</sup> Second, we consider a model specification with five statistical factors instead of four, all else being unchanged. Figures 9 and 10 report the resultant IRFs jointly with the 68% confidence bands from the baseline model for a selection of key variables. Our baseline findings are in general confirmed by these alternative specifications. Most quantitative differences emerge in the high IU state, whereas responses are remarkably robust conditionally on the low IU regime. Moreover, employing the purged IU measure leads to the most pronounced loss in effectiveness when uncertainty is high compared to the baseline. Responses of the five factor model resemble the four factor model quite closely both in qualitative and quantitative terms. Regarding the distinct mechanisms that have been scrutinized in this paper, we make the following observations: i) Hypotheses 1 and 2 are confirmed by both models and IU induced asymmetries in monetary effectiveness on macroeconomic indicators are particularly pronounced for the purged IU measure; ii) both additional specifications corroborate the conjecture that prices are more flexible conditional on low IU (hypothesis 3); iii) hypothesis 4 is broadly underpinned, even though the model using the purged IU metric indicates that the policy transmission along the yield curve is only reduced during the first year after the shock hits the economy (conditional on

<sup>28</sup>The  $R^2$  from the regression of AVEIU on the two indices amounts to 12% and hence suggest that there is indeed a non-negligible correlation.

states of elevated IU); iv) We observe that the IRFs of the alternative specifications which relate to the excess bond premium and the GZ spread are located mainly inside the confidence bands of the baseline model. Hence, these findings substantiate the view that IU is an influential variable for the transmission of monetary policy shocks via the external finance premium (hypotheses 5 and 6).

#### 4.4 Empirical evidence from a richer dataset

In this section we discuss results from the application of the ST-FAVAR model to an extended, quarterly dataset. Beyond addressing robustness of benchmark results the augmented information set enables further investigation of some hypotheses of interest, especially hypotheses 5 and 6. The quarterly panel spans the period from 1973Q3 until 2014Q4 and consists of 242 time series. In addition to the variables contained in the monthly panel it includes several balance-sheet, additional price and real economic indicators.<sup>29</sup> The number of time series is large relative to the time dimension, i.e.,  $N/T$  is considerably larger than for the benchmark panel. This allows us to assess whether neglecting the uncertainty of factor estimation is important for the conclusions drawn in the last section. Specification details are very similar to the benchmark model. Factor selection obtains again four latent factors. In light of a sizeable reduction of the time dimension model dynamics are formalized with lag order of  $p = 1$  as suggested by both AIC and BIC. Figure 11 displays the impulse responses of selected variables ensuing from a monetary policy shock normalized to reduce the ffr by 10 bps on impact. In the following, we will first discuss some general differences between both specifications and subsequently discuss the alternative theoretical channels. First of all, we find that the baseline main results are broadly confirmed by the quarterly model. We note that the response of the ffr in the low IU regime differs from that of the baseline model, while the responses in the high IU regime are almost identical. During normal times the policy instrument's response turns positive after five quarters and overshoots considerably in the longer run. It appears that this response pattern is partly reflected in the IRFs of some bond yields. Furthermore, we observe that the dynamic responses are generally more hump-shaped (at least conditional on low IU), in contrast to the more persistent IRFs drawn from the baseline model.

Next, we compare the results from the quarterly model with the baseline for specific theoretical channels. First, we investigate the dynamic responses of aggregate activity indicators in Figure 11. Confirming baseline results based on the components of industrial production, real consumption, GDP, private domestic investment and its subindices (such as residential investment) are stimulated considerably stronger conditional on states of low IU. Likewise, the responses of various labour market indicators suggest that firms are more reluctant to hire when uncertainty is high. According to these results the effects of IU on monetary policy transmission are even stronger as diagnosed for the benchmark model. Moreover, the quarterly model allows us to take a more detailed look at GDP components with a different degree of interest rate sensitivity. The reaction of interest rate sensitive components to an unexpected monetary easing is indeed considerably more pronounced in both IU regimes. For instance, investment increases by roughly 1.6% whereas the peak effect of consumption amounts to about 0.2% conditional on states of low IU. Confirming the presence of a precautionary savings channel, the responses of durable and nondurable consumption exhibit considerable regime dependence with more pronounced positive effects detected for tranquil times. The responses of major residential real estate indicators deviate from the baseline results and point to their importance for asymmetries in MP transmission. Consistent with the real-options view,

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<sup>29</sup>Details on the quarterly dataset are provided in the data appendix B.

residential investment and related series (e.g., housing starts) react more pronounced if IU is within its lower percentiles. Taken together, hypotheses 1 and 2 are supported, i.e., we find strong evidence in favour of a prime role of both the real-options and precautionary savings channels in rendering monetary policy less effective in states of elevated IU.

Turning to the responses of government bond interest rates at various maturities, our findings corroborate the conjecture from the baseline model that the interest rate channel might be impaired at all maturities in times of high IU (hypothesis 4). Conditional on states of low IU, interest rates with maturities of three-, six- and twelve-months exhibit similar response patterns and seemingly reflect the response profile of the policy rate, i.e., an initial drop followed by a hump-shaped increase after roughly 2.5 years. During states of high IU, in contrast, short-term rates barely decrease (around 1–2 bps) and show an almost linear response profile. In comparison with the baseline model and consistent with previous studies (e.g. [Gertler and Karadi, 2015](#)) the effects of an unanticipated monetary easing on the five- and ten-year interest rates are short-lived. Both interest rates decrease considerably upon impact only conditional on states of low IU. After about three quarters the responses become statistically insignificant. Conditional on states of high IU, long term interest rates respond only by roughly 1 bp on impact and revert back to their trend after one year. Consistent with an imperfect interest rate transmission, we observe a sizable, transient increase of term spreads (five- and ten-year bond rates minus ffr).

Regarding the channel via increased price flexibility (hypothesis 3), baseline results are confirmed, since prices are more flexible during low IU episodes. The additional information on the household and corporate sector balance sheet series as well as complementary credit risk measures enable a further exploration of the financial channel of monetary policy and its possible regime dependence (hypotheses 5 and 6). In this regard we detect, first, considerable asymmetries in the reaction of household sector net worth, which increases quite strongly and persistently only conditional on states of low IU. In contrast, the non-financial corporate sector net worth drops in the low IU regime for about one year and increases only in the long run, while during uncertain times the response is positive over all horizons but the effect is relatively small. Hence, this indicates the existence of an IU dependent household balance sheet channel. Second, the stronger decrease of money market and credit market spreads (TED, commercial paper and mortgage spreads) diagnosed for regimes of low IU corroborates baseline results. Combined with the corporate credit spreads analyzed in the baseline model, the spreads considered here, cover the credit costs of four important financial markets, namely, the business finance (GZ spread, EBP, commercial paper spread), housing finance (mortgage rate minus ten-year government bond rate spread), consumer finance (commercial paper spread) and the interbank market (TED spread). Accordingly, our results clearly underpin the adverse effects that IU exerts on the transmission channel via the external finance premium.

## 5 Conclusions

We examine the effects of inflation uncertainty (IU) on U.S. monetary policy transmission in a data-rich environment by means of a nonlinear FAVAR model. Employing a factor model allows us to investigate several complementary transmission channels that have been suggested in the theoretical literature to be affected by general uncertainty or specifically by IU. According to our estimates, high levels of IU reduce macroeconomic effects of monetary policy shocks. More specifically, the dynamic propagation of policy shocks through consumption spendings for durable



and nondurable goods and several investment indicators is sizably attenuated during uncertain times. Likewise, stimulative effects of expansionary monetary policy shocks in the labor market are more pronounced at low levels of IU. This conforms with both the presence of wait-and-see effects caused by sizable adjustment costs (Bloom, 2009; Bloom et al., 2012), as well as a buffer-stock/precautionary savings explanation for households and firms (Alfaro et al., 2016; Baum et al., 2009). As a result, the traditional interest rate channel becomes less effective during periods of heightened IU. In addition to these detrimental effects of IU on the interest rate channel, our results suggest that IU might influence monetary policy transmission also at an earlier stage. As indicated by regime-specific responses of yields for government bonds featuring distinct maturities, it seems that the impact of monetary policy shocks on the term structure is attenuated during episodes of elevated IU. In line with conventional theory, expectations about short term rates dominate the transmission of the policy rate to long(er) term interest rates. Surprisingly, the policy rate transmission is already markedly disrupted for short term rates during high IU. Furthermore, the responses of term premia are quantitatively negligible and appears to be largely unaffected by IU. In contrast to Castelnuovo and Pellegrino (2018), we do not find support for the hypothesis that enhanced price flexibility mitigates the real effectiveness of monetary policy (Vavra, 2013). Our findings point also to moderate regime dependence of credit market risk spreads and risk-bearing capacities of financial intermediaries. This accords with state-contingent financial market frictions associated with a higher external finance premium during times of elevated IU (as reported by Eickmeier et al., 2016). As a result, also the balance-sheet and bank lending channels are likely affected by fluctuations in IU.

In summary, consistent with a number of underlying mechanisms our analysis indicates that IU leads to asymmetries in several channels of monetary policy transmission. Furthermore, by looking specifically at IU we gain additional insights along the dimension of different types of uncertainty. Given the low correlation between our IU measure and previously employed uncertainty measures our results tend to support the view that the effects of IU are similar to those of other types of uncertainty.

Our findings are informative also from a modelling standpoint. First, they support frameworks of both nominal price rigidities (Basu and Bundick, 2017) and real frictions on the labor markets (Leduc and Liu, 2016). Second, our results advocate model setups which incorporate financial frictions resting on either financial intermediary balance-sheets, such as uncertainty dependent leverage (Eickmeier et al., 2016) or on firm balance-sheets (Gilchrist et al., 2014). Third, combining deviations from frictionless DSGE models (e.g., real business cycle models) could help to determine the relative importance of distinct mechanisms in general and for the conduct of monetary policy in particular.

In terms of policy implications, our results are supportive for a more aggressive policy conduct during times of enhanced IU. Even though, we document an uncertainty attenuating effect of expansionary monetary policy shocks based on stock market volatility, it is unclear if monetary policy (especially stimulative) reduces likewise IU. Consequently, it might be indicated and perhaps even essential to foster policy predictability. Against this background, particularly endogenizing the response of IU to monetary policy shocks and utilizing fully nonlinear responses a la Koop et al. (1996) (where the IRF is a function of the history of shocks) is a promising route for future research with the potential to improve policy recommendations (for an application, see Pellegrino, 2017). Along the same lines, the analysis of potential consequences of the zero lower bound and unconventional monetary policy measures on the policy IU nexus appears urgent and of high

interest (Basu and Bundick, 2017; Leduc and Liu, 2016). Finally, in the light of a controversial discussion regarding the “right” IU metric, it would be rewarding to extend the focus in terms of alternative measures of IU, such as disagreement statistics or survey based measures.

## Appendices

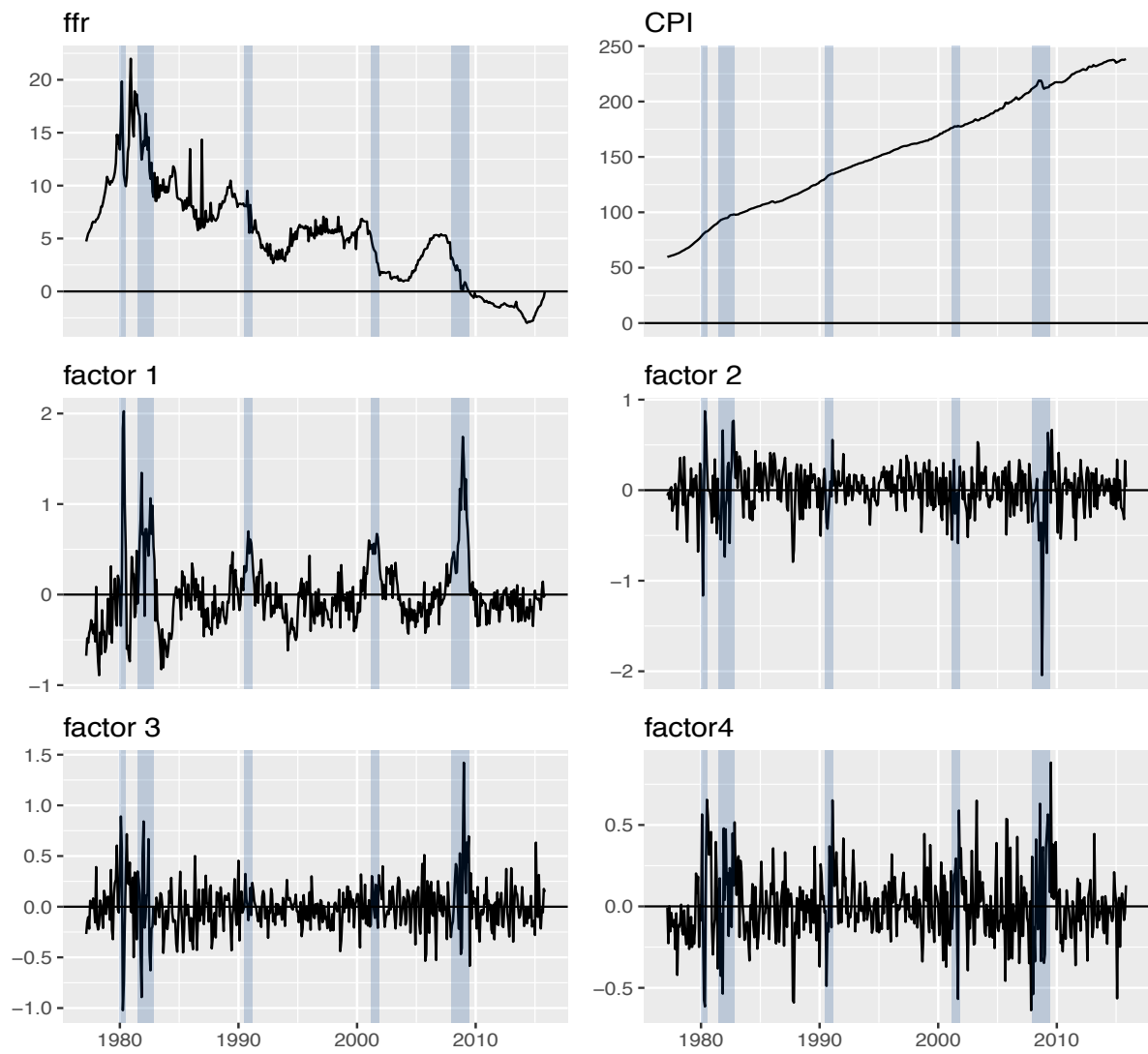
### A Figures and Tables

**Table 2:** Shares of explained variances for groups of variables

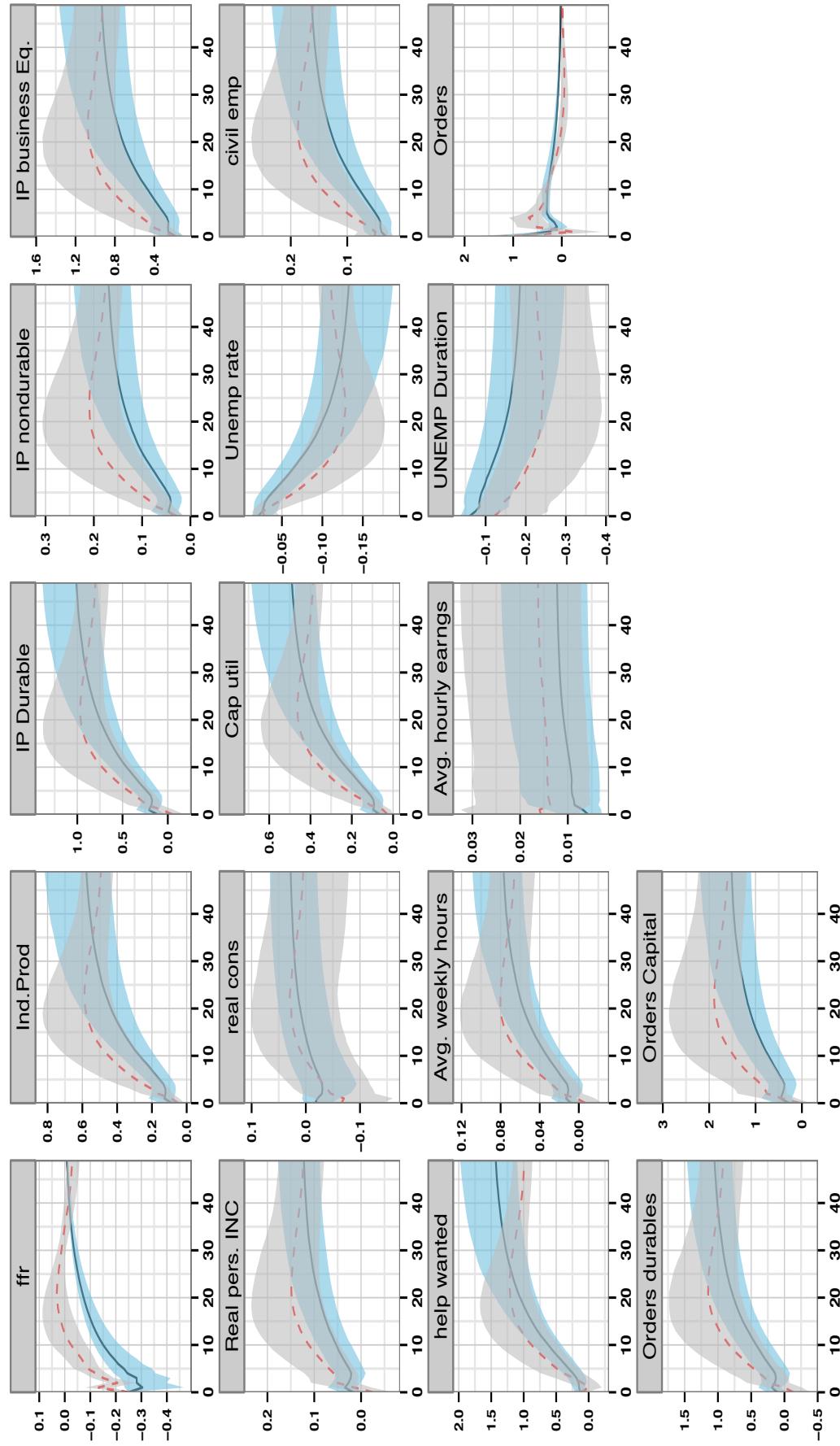
Output		labor		Cons&orders		Orders	
Variables	$R^2$	Variables	$R^2$	Variables	$R^2$	Variables	$R^2$
IPMANSICS	0.7522	USGOOD	0.7451	PERMIT	0.2690	NAPM	0.6211
INDPRO	0.7093	MANEMP	0.7412	PERMITS	0.1443	NAPMNOI	0.5435
CUMFNS	0.6731	DMANEMP	0.7290	HOUST	0.1316	NAPMII	0.3832
IPFPNSS	0.6561	PAYEMS	0.6963	PERMITW	0.1247	ISRATIOx	0.3365
IPDMAT	0.6518	USTPU	0.5630	PERMITMW	0.1021	NAPMSDI	0.3279
NAPMPI	0.5673	USWTRADE	0.5513	HOUSTS	0.0907	BUSINVx	0.3084
IPFINAL	0.5463	NAPMEI	0.5032	PERMITNE	0.0902	AMDMUOx	0.2599
IPMAT	0.5395	SRVPRD	0.4878	HOUSTW	0.0650	AMDMNOx	0.1495
IPBUSEQ	0.5090	NDMANEMP	0.4538	HOUSTMW	0.0485	ANDENOx	0.0704
CMRMTSPLx	0.3979	USCONS	0.4483	HOUSTNE	0.0219		
IPCONGD	0.3647	USTRADE	0.4141				
IPDCONGD	0.3276	UNRATE	0.3594				
RETAILx	0.2913	HWIURATIO	0.3484				
IPNMAT	0.2470	CE16OV	0.2968				
DPCERA3M086SBEA	0.2374	USFIRE	0.2634				
W875RX1	0.1956	UEMP15OV	0.2465				
RPI	0.1412	CLAIMSx	0.2424				
IPNCONGD	0.1270	HWI	0.2381				
IPB51222S	0.0406	AWHMAN	0.1799				
IPFUELS	0.0273	UEMP27OV	0.1655				
		CES0600000007	0.1629				
		AWOTMAN	0.1369				
		UEMP15T26	0.1054				
		UEMP5TO14	0.0933				
		CES1021000001	0.0745				
		USGOVT	0.0518				
		UEMPLT5	0.0467				
		CLF16OV	0.0374				
		UEMPMEAN	0.0239				

**Table 3:** Shares of explained variances for groups of variables (continued)

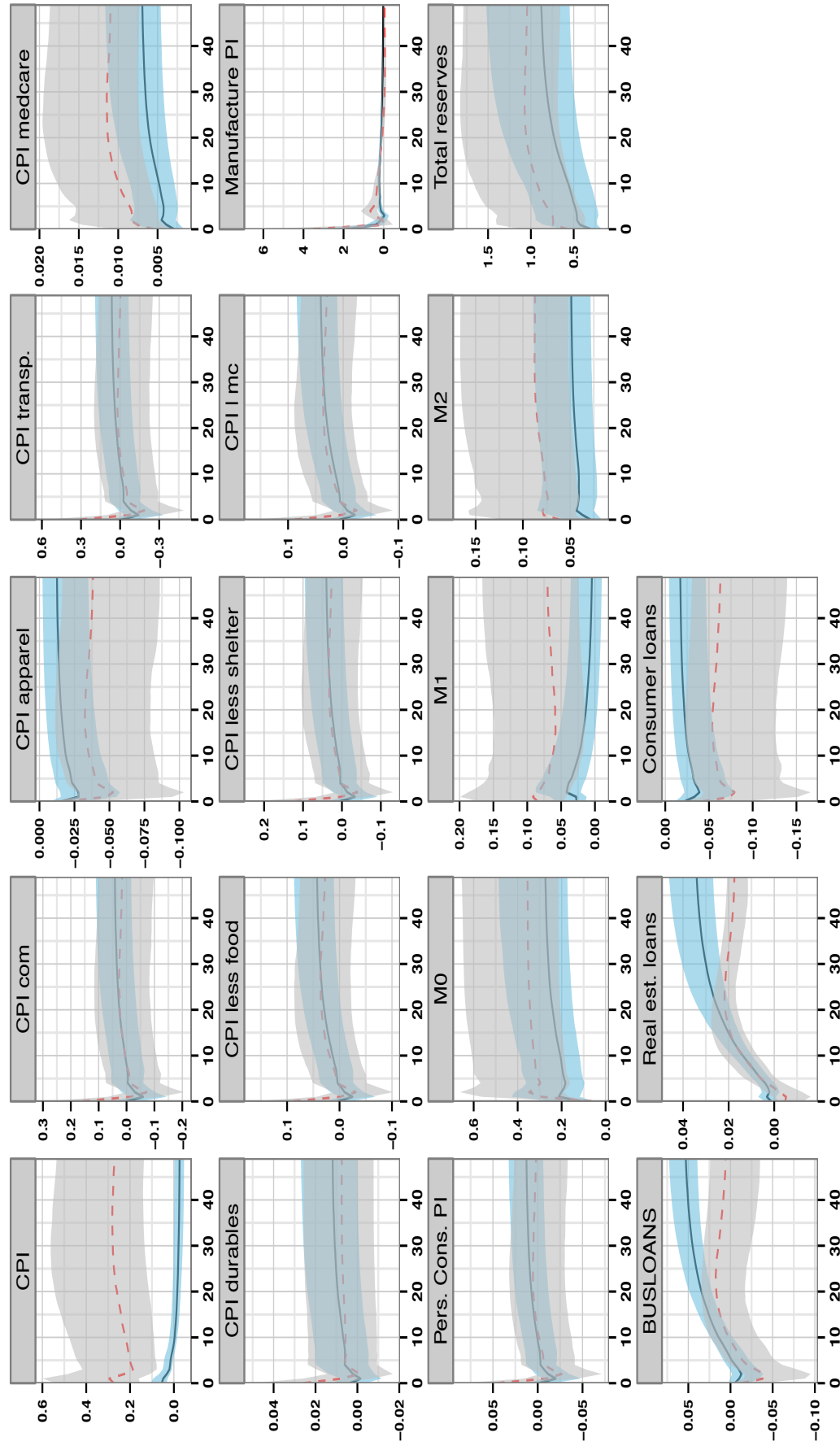
Money		interest		Asset prices		prices	
Variables	$R^2$	Variables	$R^2$	Variables	$R^2$	Variables	$R^2$
M1SL	0.0232	T10YFFM	0.8296	NYSEALL	0.8277	CUSR0000SAC	0.6806
M2SL	0.0291	T5YFFM	0.8237	S.P.500	0.8120	DNDGRG3M086SBEA	0.6580
M2REAL	0.4733	AAAFFM	0.8088	S.P.indust	0.7867	CPITRNSL	0.6420
AMBSL	0.1093	T1YFFM	0.7756	S.P.div.yield	0.7667	CUSR0000SA0L5	0.6372
TOTRESNS	0.0251	BAAFFM	0.7709	WIL5TMK	0.7639	CPIULFSL	0.6224
BUSLOANS	0.0417	TB6SMFFM	0.7522	DJINDUS	0.7459	CUUR0000SA0L2	0.5809
REALLN	0.0106	TB3SMFFM	0.6177	NASCOMP	0.6070	PCEPI	0.5382
NONREVS	0.0342	CP3Mx	0.6035	DJTRSPT	0.5400	NAPMPRI	0.4516
CONSPI	0.0551	GS1	0.5732	S.P.PE.ratio	0.5104	OILPRICEx	0.2051
		TB6MS	0.5705	DJUTILS	0.3436	PPICMM	0.0688
		AAA	0.5660	NYSTRGT	0.2718	CUSR0000SAS	0.0675
		GS5	0.5573	NAREQU..PL..U.	0.1692	DDURRG3M086SBEA	0.0480
		GS10	0.5490	CSUSHPISA	0.0394	CUUR0000SAD	0.0462
		BAA	0.5220	NYBRDTH	0.0136	DSERRG3M086SBEA	0.0252
		TB3MS	0.5186			CPIAPPSL	0.0239
		COMPAPFFx	0.2681			CPIMEDSL	0.0090
		EXCAUSx	0.2503				
		TWEXMMTH	0.1279				
		EXUSUKx	0.0845				
		EXSZUSx	0.0599				
		EXJPUSx	0.0300				



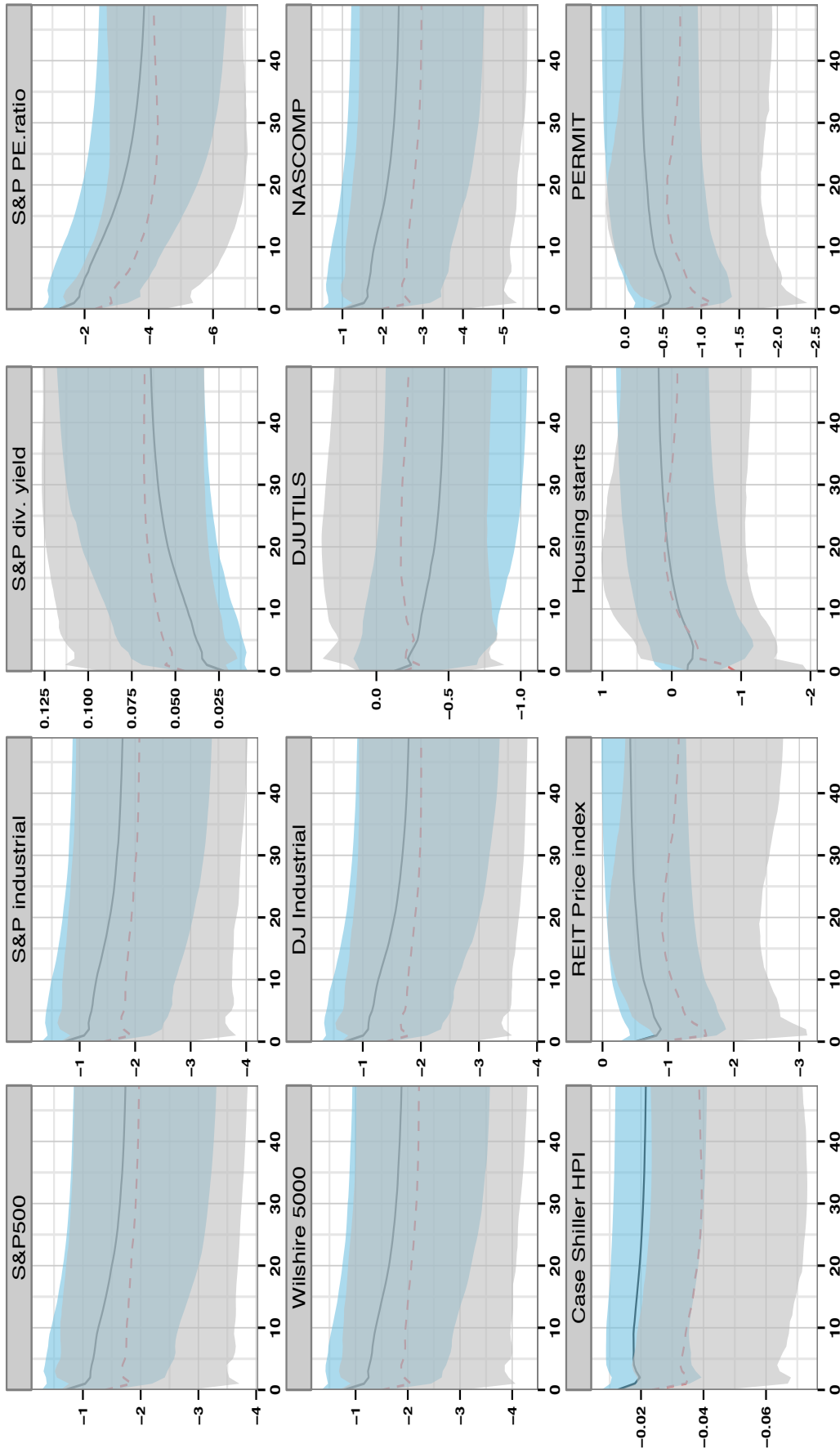
**Figure 3:** Observed and unobserved factors (principal components) for the baseline specification. For more details see Section 4.1 of the main text.



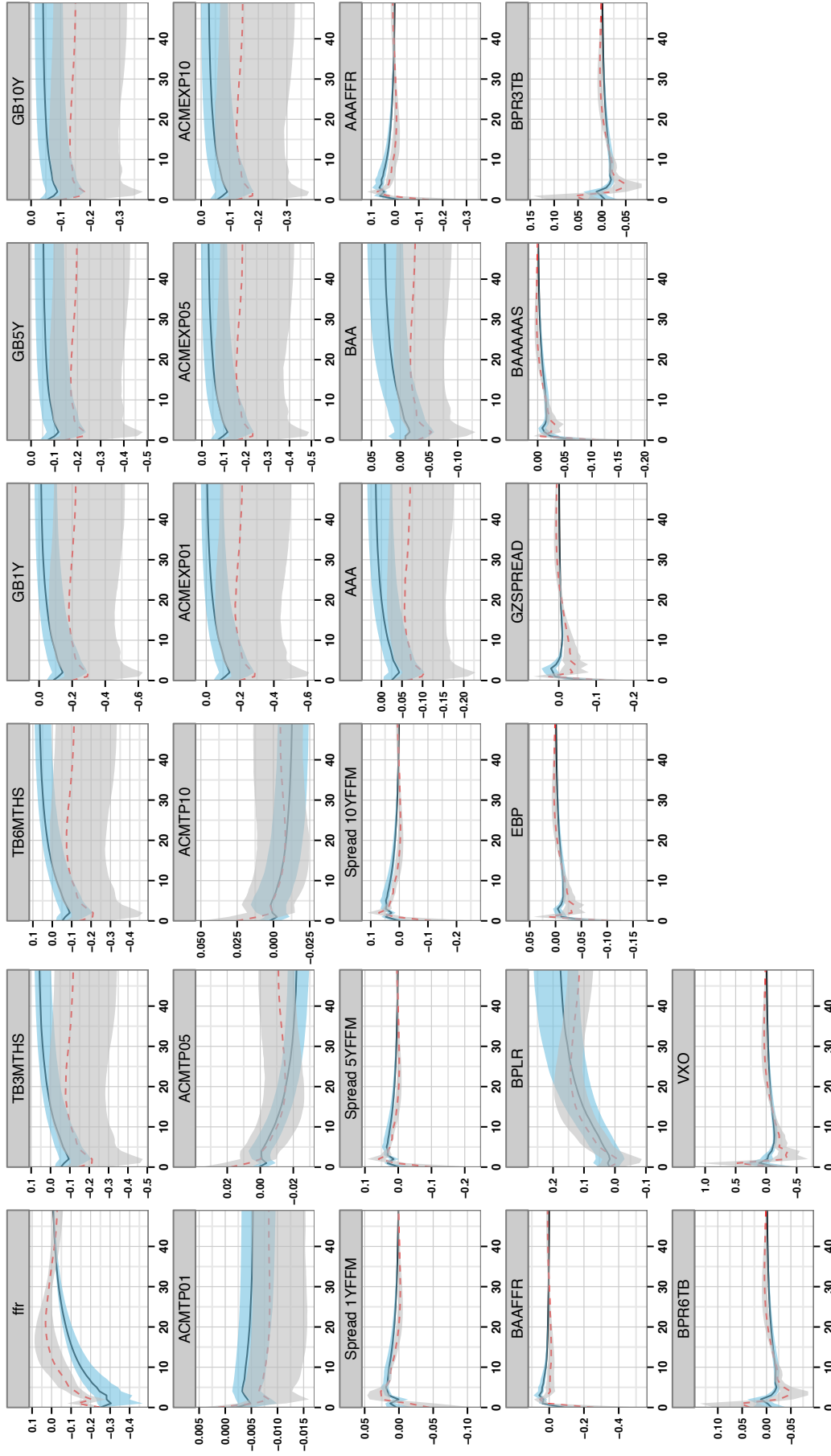
**Figure 4:** Impulse response functions of the federal funds rate, macroeconomic aggregates, such as output related variables and price indices. The charts show the median response jointly with 68% confidence bands for states of high (solid blue line and blue shaded areas) and low IU (dashed red line and grey shaded areas). Most responses are level responses, except for variables which hold the transformation code “6”, i.e., second differences. For the construction of the confidence bands and the respective interpretation, see Appendix C.



**Figure 5:** Impulse response functions of monetary aggregates, credit and real estate. For further notes see Figure 4. The charts show the median response jointly with 68% confidence bands for states of high (solid blue line and blue shaded areas) and low IU (dashed red line and grey shaded areas). Most responses are level responses, except for variables which hold the transformation code “6”, i.e., second differences.

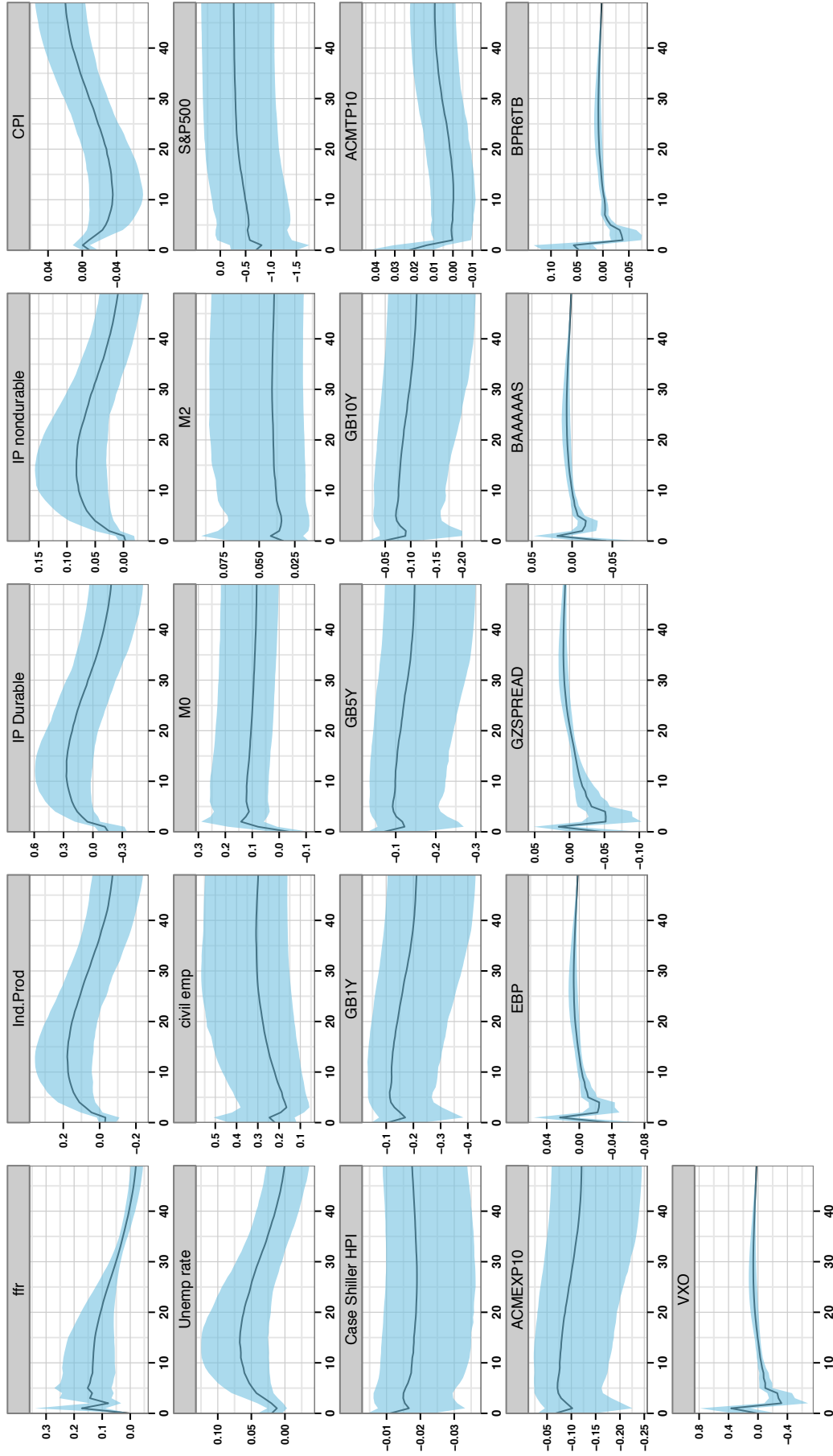


**Figure 6:** Regime-dependent impulse response functions of selected interest rates, term premia, term and credit spreads. The charts show the median response jointly with 68% confidence bands for states of high (solid blue line and blue shaded areas) and low IU (dashed red line and grey shaded areas). Most responses are level responses, except for variables which hold the transformation code "6", i.e., second differences. For further notes see Figure 4.

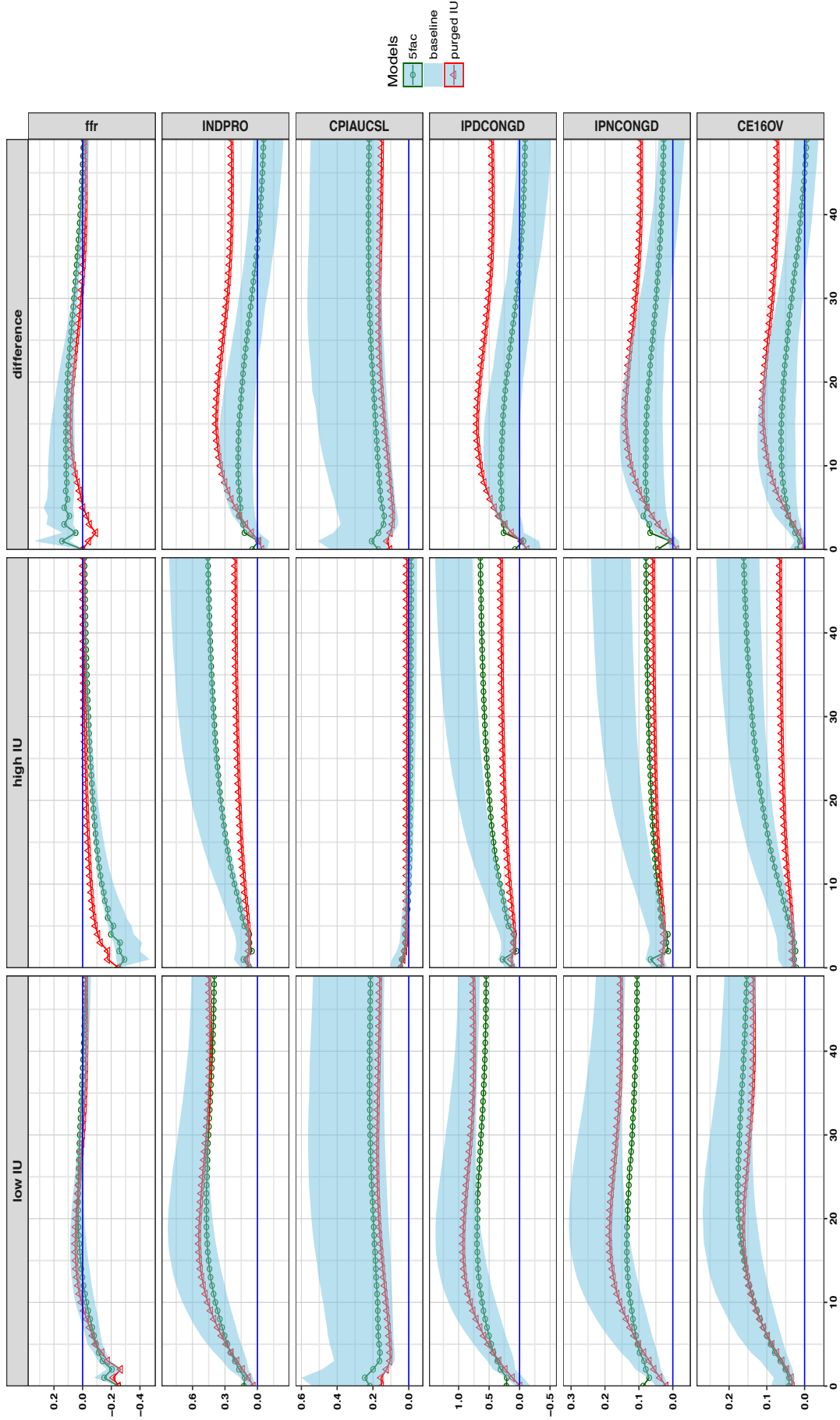


**Figure 7:** Regime-dependent impulse response functions of selected interest rates, term premia, term and credit spreads. The charts show the median response jointly with 68% confidence bands for states of high (solid blue line and blue shaded areas) and low IU (dashed red line and grey shaded areas). Most responses are level responses, except for variables which hold the transformation code "6", i.e., second differences. For further notes see Figure 4.

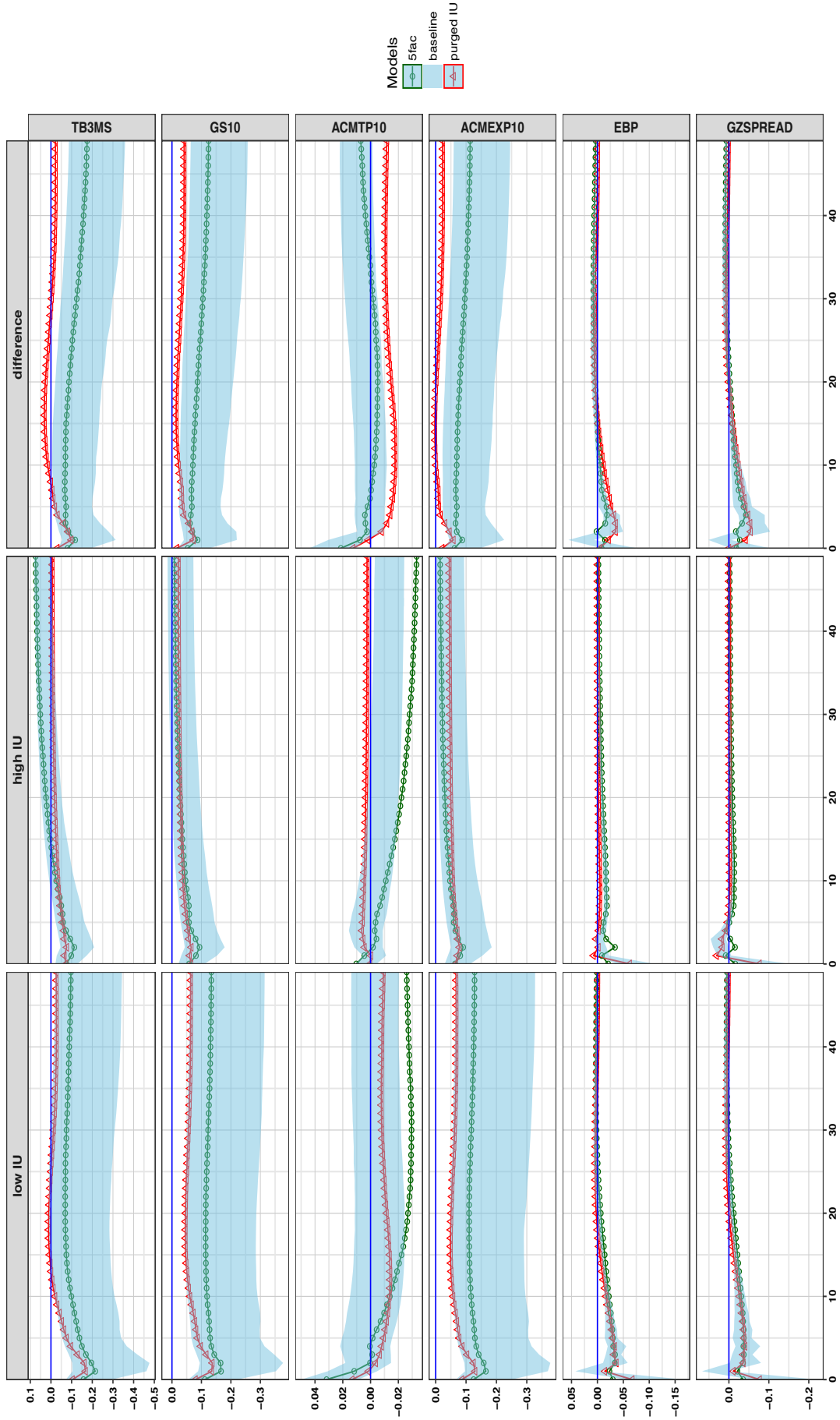




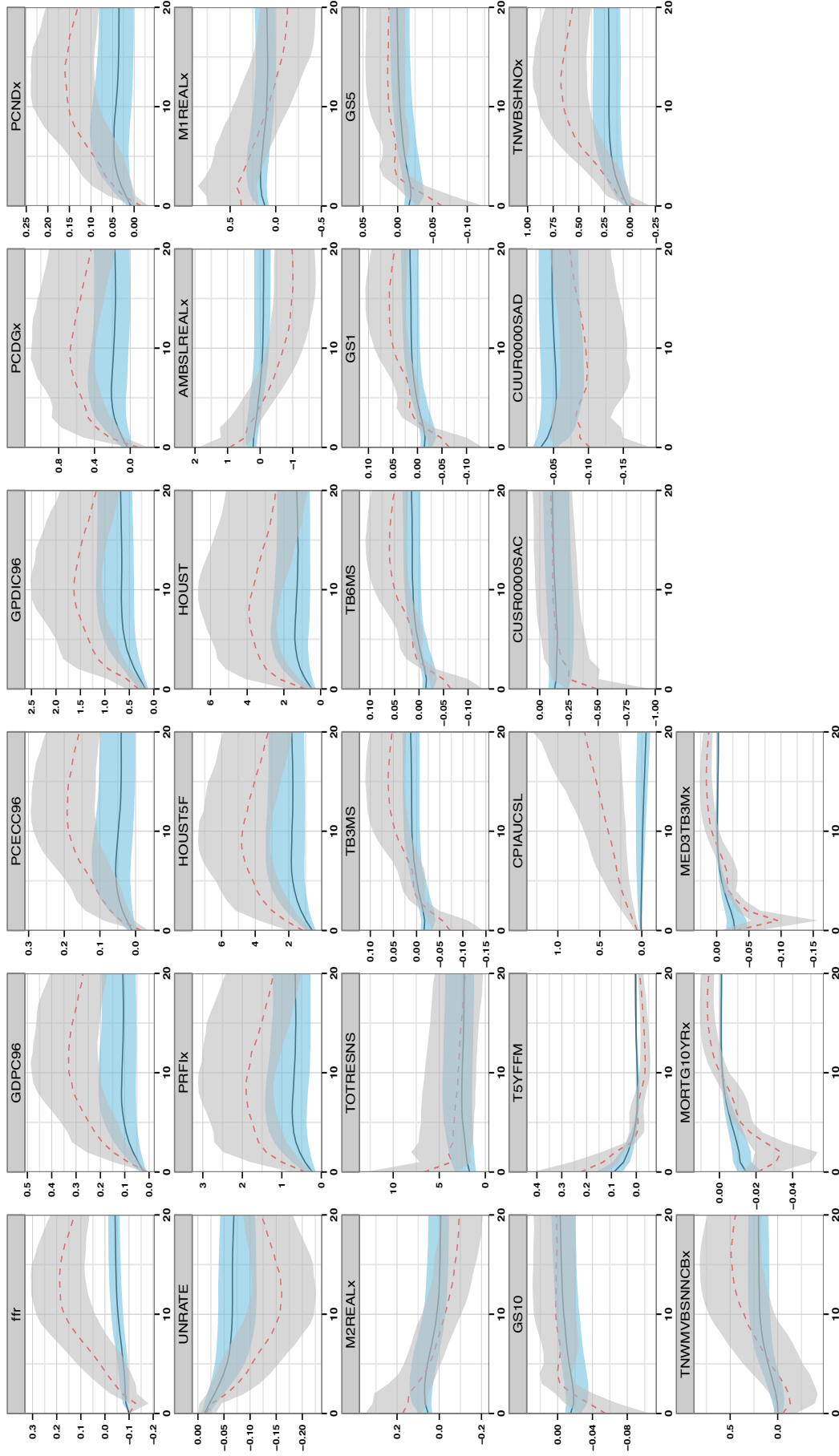
**Figure 8:** Median response jointly with 68% confidence bands for the difference between the IRFs of the low and high IU regime (i.e.,  $IRF_i^l - IRF_i^h$  for  $i = 0, \dots, H$ , where  $H = 50$ ). For further notes see Figure 4.



**Figure 9:** Robustness checks: purged IU measure model (red triangles) vs. five factor model (green circles) vs. baseline model (light blue shaded areas). The chart shows the median responses of the two alternative models jointly with 68% confidence bands for the baseline responses and selected key **economic activity and labor market variables**. The first column displays the low IU, the second column the high IU regime, while the third column depicts the difference between the the regime dependent IRFs (median responses for the alternative models and 68% confidence bands for the baseline responses), i.e.,  $IRF_i^h - IRF_i^l$  for  $i = 0, \dots, H$ , where  $H = 50$ . For further notes see Figure 4.



**Figure 10:** Robustness checks: purged IU measure model vs. five factor model and baseline model. Median response of the two alternative models (triangles and circles) jointly with 68% confidence bands for the baseline responses and selected key interest rates and financial market variables. For further notes see Figures 4 and 9.



**Figure 11:** Regime-dependent impulse response functions of selected indices for the quarterly data set. The charts show the median response jointly with 68% confidence bands for states of high (solid blue line and blue shaded areas) and low IU (dashed red line and grey shaded areas). Most responses are level responses, except for variables which hold the transformation code “6”, i.e., second differences. For further notes see Figure 4.

## B Data appendix

### Inflation uncertainty measures

Besides our baseline IU statistic AVEIU (Hartmann et al., 2017), we specified our smooth transition model alternatively by means of a model-based and a survey-based IU indicator in section 3.3.

- **Model-based:** In the main text we consider the widely-applied model-based measure, namely the forecasts of the variance of permanent shocks derived from the unobserved components stochastic volatility (called UCSV in the main text) model put forth in Stock and Watson (2007). The model assumes that inflation consists of a (permanent) trend and a transitory component/short-term shock. Trend inflation is driven by level shocks and the conditional log variances of both shocks follow independent random walks. The UCSV model is estimated by means of Gibbs sampling. The variance parameter which governs the smoothness of the variance processes is set to  $\theta = (0.2/3)^2$ . For a detailed description of the UCSV model see Stock and Watson (2007) and Hartmann et al. (2017). Naturally, this IU indicator neglects any model uncertainty and has no disagreement component.
- **Survey-based:** The survey-based measure, denoted ‘BINDER’ in the main text, exploits the fact that more uncertain respondents exhibit a specific rounding behaviour when reporting their inflation expectations (“Round numbers suggest round interpretations”) (Binder, 2017). It is based on the Michigan Survey of Consumers and available from 1978M1 onwards for the forecast horizon of one year. This approach allows to construct a statistic relating to aggregate (individual) uncertainty, even in the absence of density forecasts in surveys. Binder (2017) provides ample evidence that this indicator obtains desirable properties for measures of IU. The data is available at <https://sites.google.com/site/inflationuncertainty/>.

### Uncertainty measures

In section 3.3 of the main text we mention the following uncertainty indicators:

- The **GZ spread** is a corporate bond risk spread summarizing the information of secondary market prices of 5,982 senior unsecured bonds issued by a representative sample of 1,112 U.S. non-financial firms developed in Gilchrist and Zakrajšek (2012). The GZ spread covers the entire spectrum of credit quality, and duration mismatch is avoided by constructing a synthetic risk-free security that exactly mimics the cash flows of individual bonds. Gilchrist and Zakrajšek (2012) demonstrate that their credit spread outperforms the forecasting ability of widely used credit spreads (such as the Baa-Aaa corporate bond credit spread). The GZ spread data is available at <https://www.federalreserve.gov/econresdata/notes/feds-notes/2016/recession-risk-and-the-excess-bond-premium-accessible-20160408.html#fig2>.
- The general macroeconomic uncertainty index constructed by Jurado et al. (2015) (JLN12) is the common unforecastable component of a panel of 134 economic time series. Jurado et al. (2015) proceed by separating the forecastable and unforecastable components of economic indicators. Stochastic volatility models are fitted to all individual forecast error series and the final uncertainty index is constructed as the common factor of all individual volatility series. The data is available at <https://www.sydneyludvigson.com/data-and-appendixes/>.

- The **VXO volatility index** as the S&P 100 options implied volatility index is taken from the FRED-MD database available at John W. McCrackens web page <https://research.stlouisfed.org/econ/mccracken/fred-databases/>.

### Monthly panel

The data used in the analysis is taken from [McCracken and Ng \(2016\)](#) and available at John W. McCrackens web page <https://research.stlouisfed.org/econ/mccracken/fred-databases/>. We added several time series, mainly financial and asset market indices to the database. Particularly, the term structure related quantities, such as term premia and average expected short-term rates have been obtained from [https://www.newyorkfed.org/research/data\\_indicators/term\\_premia.html](https://www.newyorkfed.org/research/data_indicators/term_premia.html). The shadow rate as constructed by [Wu and Xia \(2016\)](#) can be downloaded at [https://www.frbatlanta.org/cqer/research/shadow\\_rate.aspx?panel=1](https://www.frbatlanta.org/cqer/research/shadow_rate.aspx?panel=1). All other time series are available via the FRED online database.

In the Tables below the time series are documented as follows. “Index” indicates the position of the variable in the dataset used in the analysis. This might differ from the original data set, since some variables are left out due to missing observations. “Fred” gives the suffix of the time series as it appears in the original database. Each time series contains a short description and “T-code” refers to the following transformations: (1) no transformation; (2)  $\Delta x_t$ ; (3)  $\Delta^2 x_t$ ; (4)  $\log(x_t)$ ; (5)  $\Delta \log(x_t)$ ; (6)  $\Delta^2 \log(x_t)$  and (7)  $\Delta(x_t/x_{t-1} - 1.0)$ . Slow-moving variables are highlighted by an asterisk.

**Table 4: Group 1: Output and income**

Index	Fred	Description	T-code
1	RPI*	Real Personal Income	5
2	W875RX1*	Real Personal Income ex transfer receipts	5
6	INDPRO*	IP Index	5
7	IPFPNSS*	IP: Final Products and Nonindustrial Supplies	5
8	IPFINAL*	IP: Final Products (Market Group)	5
9	IPCONGD*	IP: Consumer Goods	5
10	IPDCONGD*	IP: Durable Consumer Goods	5
11	IPNCONGD*	IP: Nondurable Consumer Goods	5
12	IPBUSEQ*	IP: Business Equipment	5
13	IPMAT*	IP: Materials	5
14	IPDMAT*	IP: Durable Materials	5
15	IPNMAT*	IP: Nondurable Materials	5
16	IPMANSICS*	IP: Manufacturing (SIC)	5
17	IPB51222S*	IP: Residential Utilities	5
18	IPFUELS*	IP: Fuels	5
19	NAPMPI*	ISM Manufacturing: Production Index	1
20	CUMFNS*	Capacity Utilization: Manufacturing	2

**Table 5: Group 2: Labor market**

Index	Fred	Description	T-code
21	HWI*	Help-Wanted Index for United States	2
22	HWIURATIO*	Ratio of Help Wanted /No. Unemployed	2
23	CLF16OV*	Civilian Labor Force	5
24	CE16OV*	Civilian Employment	5
25	UNRATE*	Civilian Unemployment Rate	2
26	UEMPMEAN*	Average Duration of Unemployment (Weeks)	2
27	UEMPLT5*	Civilians Unemployed - Less Than 5 Weeks	5
28	UEMP5TO14*	Civilians Unemployed for 5-14 Weeks	5
29	UEMP15OV*	Civilians Unemployed - 15 Weeks & Over	5
30	UEMP15T26*	Civilians Unemployed for 15-26 Weeks	5
31	UEMP27OV*	Civilians Unemployed for 27 Weeks and Over	5
32	CLAIMSx*	Initial Claims	5
33	PAYEMS*	All Employees: Total nonfarm	5
34	USGOOD*	All Employees: Goods-Producing Industries	5
35	CES1021000001*	All Employees: Mining and Logging: Mining	5
36	USCONS*	All Employees: Construction	5
37	MANEMP*	All Employees: Manufacturing	5
38	DMANEMP*	All Employees: Durable goods	5
39	NDMANEMP*	All Employees: Nondurable goods	5
40	SRVPRD*	All Employees: Service-Providing Industries	5
41	USTPU*	All Employees: Trade, Transportation & Utilities	5
42	USWTRADE*	All Employees: Wholesale Trade	5
43	USTRADE*	All Employees: Retail Trade	5
44	USFIRE*	All Employees: Financial Activities	5
45	USGOVT*	All Employees: Government	5
46	CES0600000007*	Avg Weekly Hours: Goods-Producing	1
47	AWOTMAN*	Avg Weekly Overtime Hours: Manufacturing	2
48	AWHMAN*	Avg Weekly Hours: Manufacturing	1
49	NAPMEI*	ISM Manufacturing: Employment Index	1
124	CES0600000008*	Avg Hourly Earnings: Goods-Producing	6
125	CES2000000008*	Avg Hourly Earnings: Construction	6
126	CES3000000008*	Avg Hourly Earnings: Manufacturing	6

**Table 6: Group 3: Housing starts and permits**

Index	Fred	Description	T-code
50	HOUST*	Housing Starts: Total New Privately Owned	4
51	HOUSTNE*	Housing Starts, Northeast	4
52	HOUSTMW*	Housing Starts, Midwest	4
53	HOUSTS*	Housing Starts, South	4
54	HOUSTW*	Housing Starts, West	4
55	PERMIT*	New Private Housing Permits (SAAR)	4
56	PERMITNE*	New Private Housing Permits, Northeast (SAAR)	4
57	PERMITMW*	New Private Housing Permits, Midwest (SAAR)	4
58	PERMITS*	New Private Housing Permits, South (SAAR)	4
59	PERMITW*	New Private Housing Permits, West (SAAR)	4

**Table 7: Group 4: Orders and inventories**

Index	Fred	Description	T-code
3	DPCERA3M086SBEA	Real personal consumption expenditures	5
4	CMRMTSPLx	Real Manu. and Trade Industries Sales	5
5	RETAILx	Retail and Food Services Sales	5
60	NAPM	ISM : PMI Composite Index	1
61	NAPMNOI	ISM : New Orders Index	1
62	NAPMSDI	ISM : Supplier Deliveries Index	1
63	NAPMII	ISM : Inventories Index	1
64	AMDMNOx	New Orders for Durable Goods	5
65	ANDENOx	New Orders for Nondefense Capital Goods	5
66	AMDMUOx	Unfilled Orders for Durable Goods	5
67	BUSINVx	Total Business Inventories	5
68	ISRATIOx	Total Business: Inventories to Sales Ratio	2

**Table 8: Group 5: Money and credit**

Index	Fred	Description	T-code
69	M1SL	M1 Money Stock	6
70	M2SL	M2 Money Stock	6
71	M2REAL	Real M2 Money Stock	5
72	AMBSL	St. Louis Adjusted Monetary Base	6
73	TOTRESNS	Total Reserves of Depository Institutions	6
74	BUSLOANS	Commercial and Industrial Loans	6
75	REALLN	Real Estate Loans at All Commercial Banks	6
76	NONREVSL	Total Nonrevolving Credit	6
77	CONSPI	Nonrevolving consumer credit to Personal Income	2
127	MZMSL	MZM Money Stock	6
128	DTCOLNVHFNM	Consumer Motor Vehicle Loans Outstanding	6
129	DTCTHFNM	Total Consumer Loans and Leases Outstanding	6
130	INVEST	Securities in Bank Credit at All Commercial Banks	6

**Table 9: Group 6: Interest rates and exchange rates**



Index	Fred	Description	T-code
	FEDFUNDS	Effective Federal Funds Rate	2
82	CP3Mx	3-Month AA Financial Commercial Paper Rate	2
83	TB3MS	3-Month Treasury Bill	2
84	TB6MS	6-Month Treasury Bill	2
85	GS1	1-Year Treasury Rate	2
86	GS5	5-Year Treasury Rate	2
87	GS10	10-Year Treasury Rate	2
88	AAA	Moody's Seasoned Aaa Corporate Bond Yield	2
89	BAA	Moody's Seasoned Baa Corporate Bond Yield	2
90	COMPAPFFx	3-Month Commercial Paper Minus FEDFUNDS	1
91	TB3SMFFM	3-Month Treasury C Minus FEDFUNDS	1
92	TB6SMFFM	6-Month Treasury C Minus FEDFUNDS	1
93	T1YFFM	1-Year Treasury C Minus FEDFUNDS	1
94	T5YFFM	5-Year Treasury C Minus FEDFUNDS	1
95	T10YFFM	10-Year Treasury C Minus FEDFUNDS	1
96	AAAFFM	Moody's Aaa Corporate Bond Minus FEDFUNDS	1
97	BAAFFM	Moody's Baa Corporate Bond Minus FEDFUNDS	1
98	TWEXMMTH	Trade Weighted U.S. Dollar Index: Major Currencies	5
99	EXSZUSx	Switzerland / U.S. Foreign Exchange Rate	5
100	EXJPUSx	Japan / U.S. Foreign Exchange Rate	5
101	EXUSUKx	U.S. / U.K. Foreign Exchange Rate	5
102	EXCAUSx	Canada / U.S. Foreign Exchange Rate	5
149	BPLR	Bank prime loan rate	2
150	EBP	Excess bond premium	1
151	GZSPREAD	GZ spread	1
152	ACMTP01	Term premium for 1-Year Treasury bonds	2
153	ACMTP05	Term premium for 5-Year Treasury bonds	2
154	ACMTP10	Term premium for 10-Year Treasury bonds	2
155	ACMEXP01	Average expected short term rates 1-Year Treasury bond	2
156	ACMEXP05	Average expected short term rates 1-Year Treasury bond	2
157	ACMEXP10	Average expected short term rates 10-Year Treasury bond	2
158	BAAAAAS	Moody's Baa Corporate Bond Minus AAA	1
159	BPR3TB	Bank prime loan rate Minus 3-Month Treasury Bill	1
160	BPR6TB	Bank prime loan rate Minus 6-Month Treasury Bill	1

**Table 10: Group 7: Prices**

Index	Fred	Description	T-code
103	WPSFD49207*	PPI: Finished Goods	6
104	WPSFD49502*	PPI: Finished Consumer Goods	6
105	WPSID61*	PPI: Processed Goods for Intermediate Demand	6
106	WPSID62*	PPI: Unprocessed Goods for Intermediate Demand	6
107	OILPRICEx*	Crude Oil, spliced WTI and Cushing	6
108	PPICMM*	PPI: Metals and metal products	6
109	NAPMPRI*	ISM Manufacturing: Prices Index	1
110	CPIAUCSL*	CPI : All Items	6
111	CPIAPPSL*	CPI : Apparel	6
112	CPITRNSL*	CPI : Transportation	6
113	CPIMEDSL*	CPI : Medical Care	6
114	CUSR0000SAC*	CPI : Commodities	6
115	CUUR0000SAD*	CPI : Durables	6
116	CUSR0000SAS*	CPI : Services	6
117	CPIULFSL*	CPI : All Items less Food	6
118	CUUR0000SA0L2*	CPI : All items less Shelter	6
119	CUSR0000SA0L5*	CPI : All items less Medical Care	6
120	PCEPI*	Personal Cons. Expend.: Chain Index	6
121	DDURRG3M086SBEA*	Personal Cons. Exp: Durable goods	6
122	DNDGRG3M086SBEA*	Personal Cons. Exp: Nondurable goods	6
123	DSERRG3M086SBEA*	Personal Cons. Exp: Services	6

**Table 11: Group 8: Stock market**

Index	Fred	Description	T-code
78	S&P 500	S&P's Common Stock Price Index: Composite	5
79	S&P: indust	S&P's Common Stock Price Index: Industrials	6
80	S&P div yield	S&P's Composite Common Stock: Dividend Yield	2
81	S&P PE ratio	S&P's Composite Common Stock: Price-Earnings Ratio	5
131	VXOCLS	CBOE S&P's 100 Volatility Index: VXO	1
132	USWSBC..A	US Foreign Sales of US Corporate Bonds to US Residents	5
133	USWPBC..A	US Foreign Purchases of US Corporate Stocks	5
134	USWPBC..A	US Foreign Purchases of US Corporate Bonds	5
135	USSHRPRCF	Dow Jones Industrials Share Price Index	5
136	USCBDMGNA	NYSE Debt Balances in margin Accounts	5
137	USCBDMGNA	Free Credit Balances in margin Accounts	5
138	CSUSHPISA	Case Shiller HPI	6
139	NAREQU..PL..U.	FTSE 100/NAREIT All Equity REIT Price Index	5
140	WIL5TMK	Wilshire 5000 Total Market Index	5
141	NYSTRGT	NYSE Total Strength Index	1
142	NYSEALL	NYSE Composite - Price Index	5
143	NYBRDTH	NYSE Total Breadth Index	5
144	NASCOMP	NASDAQ Composite - Price Index	5
145	DJUTILS	Dow Jones Utilities Price Index	5
146	DJTRSP	Dow Jones Transportation Price Index	5
147	DJINDUS	Dow Jones Industrials Price Index	5

### Quarterly panel

The quarterly dataset is the FRED-QD database, a companion to FRED-MD (monthly database), and was retrieved from John W. McCrackens web page <https://research.stlouisfed.org/econ/mccracken/fred-databases/>. For the sake of brevity, we do not reproduce the entire data description tables for the quarterly panel. Information on all originally included time series can be found in the FRED-QD appendix at [https://s3.amazonaws.com/files.fred.stlouisfed.org/fred-md/FRED-QD\\_appendix.pdf](https://s3.amazonaws.com/files.fred.stlouisfed.org/fred-md/FRED-QD_appendix.pdf). We used the data transformation codes proposed in the original data appendix in order to establish stationarity of the panel. Nonetheless, we had to remove three time series (OUTMS, OPHMFG, ULCMFG), due to missing observations at the beginning of the sample.

## C Estimation and sign restrictions

### MCMC sampling

We follow mainly [Auerbach and Gorodnichenko \(2012\)](#) and outline only briefly the estimation procedure. We use the MCMC methods proposed in [Chernozhukov and Hong \(2003\)](#). This method delivers both a global optimum and densities for the parameter estimates. Noting that the ST-FAVAR model is linear in lag polynomials  $A_l(L), A_h(L)$  conditional on the transition function parameters  $c, \gamma$  and  $\Omega_l, \Omega_h$ , a Metropolis-Hastings algorithm is employed to sample  $\Omega_l, \Omega_h$  (and potentially also  $\gamma, c$ ). For a given draw  $A_l(L), A_h(L)$  are estimated by feasible GLS. Starting values for  $\Omega_l, \Omega_h$  and  $\gamma$  are obtained as follows. We consider a second order Taylor expansion of Eq. (2) approximating the ST-FAVAR model and retain respective residuals  $\hat{u}_t$ . We then estimate Eq. (2) by maximum likelihood substituting  $\hat{u}_t$  for  $u_t$ . In the empirical application, we fix the location parameter  $c$  of  $G(z)$  at zero, whereas  $\gamma$  is jointly estimated with the remaining parameters. This results in parameters chains  $\{\Xi\}_{i=1}^S$  of length  $S$ . Simulations are based on 100.000 draws and 40.000 are discarded as burn-in period. During the burn-in period the proposal density is adjusted on the fly with respect to the scale, ensuring an acceptance rate between 0.2 and 0.3. Moreover, the covariance matrix of the proposal density is updated to be proportional to the posterior covariance matrix estimated from the previous draws.

### Sign restrictions

We impose theoretically motivated restrictions directly on the short-run impulse responses in the spirit of [Ahmadi and Uhlig \(2015\)](#) and [Amir-Ahmadi and Ritschl \(2009\)](#) for two reasons: First, it does not require a structural interpretation of the factors and second, this approach narrows down the set of admissible models by imposing restrictions on a large set of responses. For the implementation of sign restrictions we follow closely the method proposed by [Rubio-Ramirez et al. \(2010\)](#). For a comprehensive discussion of set identified SVARs the reader is referred to Ch. 13 in [Killian and Lütkepohl \(2017\)](#). Starting point for the algorithm are regime-specific structural impact matrices  $\hat{B}_\bullet$  obtained as the lower triangular Cholesky factors of  $\hat{\Omega}_\bullet = \hat{B}_\bullet \hat{B}_\bullet'$ . Next, we generate a  $K \times K$  rotation matrix  $Q$  by applying the QR decomposition to  $W$  a matrix with elements independently drawn from a  $N(0, 1)$  distribution, i.e.,  $W = QR$ . As [Rubio-Ramirez et al. \(2010\)](#) show, given that the diagonal of the upper triangular matrix  $R$  has been normalized to be positive, this algorithm amounts to drawing from a uniform distribution over the space of orthogonal matrices  $\mathcal{O}(K)$ . The rotation matrix  $Q$  is used to draw candidate impact matrices  $\hat{B}_\bullet^* = \hat{B}_\bullet Q$ ,  $\bullet = l, h$  for both regimes. Along these lines, the construction of quantile specific candidate impact matrices is straightforward. Specifying the desired quantile  $q_a$ , e.g.,  $F(z_t \leq z^*) = q_a$ , the candidate impact matrix can be obtained as  $\hat{B}_{q_a}^* = \hat{B}_{q_a} Q$ , where  $\hat{B}_{q_a}$  is the lower triangular Cholesky factor of the quantile specific reduced form covariance matrix  $\hat{\Omega}_{q_a} = \hat{G}(\hat{c}, \hat{\gamma}, z_t = F^{-1}(q_a))\hat{\Omega}_l + (1 - \hat{G}(\hat{c}, \hat{\gamma}, z_t = F^{-1}(q_a)))\hat{\Omega}_h$ . Note that the candidate structural impact matrices of both regimes come from the same rotation matrix  $Q$ . We then check if the set of maintained sign restrictions is fulfilled, in which case we retain the draw otherwise it is discarded. Thereby, draws are only retained if only one shock is consistent with the maintained restrictions, avoiding the so-called multiple shocks problem ([Fry and Pagan, 2011](#)).

To account for both estimation and identification uncertainty we sample  $D_2$  reduced form models from the parameter chain  $\{\Xi\}_{i=1}^I$  and generate  $D_1$  candidate structural models for each of

the  $D_2$  parameter draws. Throughout the paper we use 1000 structural models for inference.

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