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LEAN AGAINST THE WIND OR FLOAT WITH THE STORM? REVISITING THE MONETARY POLICY ASSET PRICE NEXUS BY MEANS OF A NOVEL STATISTICAL IDENTIFICATION APPROACH

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Lean against the wind or float with the storm? Revisiting the monetary policy asset price nexus by means of a novel statistical identification approach*

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Abstract

This paper revisits the monetary policy asset price nexus employing a novel identification approach for structural VARs in a framework of non-Gaussian independent shocks. This allows us to remain "agnostic" about the contemporaneous relations between the variables. We provide empirical evidence on the U.S. economy for monetary policy shocks and shocks originating from two asset markets: Equity and housing. Our results indicate that contractionary monetary policy shocks have a mildly negative impact on both asset prices. The effect is less pronounced for equity. Moreover, we find considerable differences in the speed of monetary policy transmission among both asset classes.(JEL C32, E44, E52)

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1 Introduction

The role of asset prices in the business cycle and their effects on consumer prices have received considerable attention from academics and, in particular, policy makers. From a theoretical point of view, this interest is based on several linkages between asset markets and the macroeconomy, pointing to their relevance in the transmission of monetary policy (see Mishkin, 2001, for an overview). These linkages comprise the wealth channel of asset prices affecting consumption (Goodhart and Hofmann, 2008), transmission via investment decisions (Tobin's q), and a credit channel operating via collateral effects on the external finance premium of households and firms (Bernanke et al., 1999; Adrian and Shin, 2010; Borio and Zhu, 2012).[1]

The crisis of 2008–2009 has arrestingly demonstrated that asset prices are not only a transmitter but also an important independent source of shocks. It became again obvious that asset prices might deviate from fundamentals and include bubble components which eventually cause adverse real economic effects (Cecchetti et al., 2000). Therefore, the Great Recession has re-ignited the "lean against the wind" (LATW) debate, addressing the necessity to monitor asset prices more closely and incorporate them into the monetary policy rule.[2] A minimum prerequisite for a LATW strategy to be feasible is that the central bank is capable of effectively steering asset market developments in the desired direction.[3]

The influence of monetary policy on asset valuation is, however, discussed controversially in the theoretical as well as empirical literature. Theoretically, the conventional view that asset prices respond negatively to surprise interest rate hikes has recently been challenged by the notion of rational bubbles (see e.g., Galí, 2014). Empirically, the relationship between monetary policy and asset prices has been widely investigated by means of structural VARs. Recently, Björnland and Leitemo (2009) have criticized the assumption of recursive transmission schemes employed in the earlier literature (e.g., Thorbecke, 1997; Neri, 2004). Björnland and Leitemo (2009) find an active contemporaneous bidirectional linkage between monetary policy and stock prices by employing just identifying long-run restrictions within a five dimensional SVAR for the U.S. economy. In contrast, Lütkepohl and Netšunajev (2014) show within an over-identified heteroskedasticty-based identification scheme that the long-run restrictions in Björnland and Leitemo (2009) are at odds with the data. Galí and Gambetti (2015) even report evidence for a positive response of stock prices to contractionary monetary policy shocks.

In the light of these contradicting findings, we adopt a novel identification approach that allows us to leave the transmission pattern of initial effects largely unrestricted. Hence, we avoid making specific assumptions which might contribute to the controversial findings in the literature, and which are prone to evoking misspecification of the structural system. Specifically, identification is achieved in a non-Gaussian framework by means of detecting independent structural shocks.

In order to demonstrate that the detection of independent shocks provides an instrument to

[1] For instance, house equity withdrawal has been identified as an important determinant of private consumption expenditures in the U.S., especially during the build-up of the subprime crisis (Duca et al., 2010; Bezemer, 2009).

[2] In this paper, we refer to monetary policy in the narrow sense as interest rate setting policies and exclude, for instance, macroprudential regulation.

[3] Besides the prerequisite above mentioned, a LATW strategy requires the ability of central banks to detect the build-up or presence of an asset price bubble. Assenmacher-Wesche and Gerlach (2010) and Smets (2014) provide a thorough discussion on the LATW debate.

distinguish between economically meaningful shocks, we investigate a five dimensional system of U.S. variables comprising a short-term interest rate and stock prices. Such systems have also been investigated by Björnland and Leitemo (2009) and Lütkepohl and Netšunajev (2014) and allow a direct comparison with their findings. As it turns out, the dynamic patterns of stock price and monetary policy shocks recovered in this work are broadly in line with results based on the Björnland and Leitemo (2009) identification approach.

In a next step, we focus on a more informative model in the LATW context. More specifically, we include another asset price in the model, viz house prices. The reasons for doing so are compelling theoretical and empirical arguments which imply that focusing on a single asset class might conceal important aspects of the dynamics between asset markets and monetary policy. In particular, from a theoretical perspective, modeling two asset classes allows one to: *i.*) account for differences in the effectiveness of monetary policy with regard to distinct asset markets; *ii.*) incorporate possible mismatches in the speed of price adjustments in distinct markets which might result in trading one imbalance for another (Assenmacher-Wesche and Gerlach, 2008); *iii.*) unveil possible differences in the effect size and dynamic propagation of different asset price shocks on inflation and output.[4] Including residential property, being the predominant form of wealth for households, is particularly relevant for the policy maker in this context owing to its central role in business cycles (Leamer, 2008, 2015) and for financial stability (IMF, 2015). Moreover, we can expect a difference in the timing of policy effects, since house prices tend to react more sluggishly to news than stock prices.[5] These theoretical considerations have been substantiated by several empirical studies. For instance, results in Björnland and Jacobsen (2013) and Simo-Kengne et al. (2016), highlight a vital role for housing, and reveal major differences between the inertia and persistence profiles of both asset price responses to monetary policy shocks and vice versa the reaction of monetary policy to both asset price shocks (see Section 3.1 for details). Moreover, in a data-rich environment Eickmeier and Hofmann (2013) provide evidence suggesting that interest rate shocks do not affect stock prices, but impact quite strongly on house prices.

Previewing the core results from our multiple asset price model, first we find evidence for an active transmission channel of monetary policy via the asset markets. Monetary policy has a moderately negative impact on real house prices which is sluggish and persistent. Our findings draw a different picture for real stock prices. The effect of interest rate shocks on equity prices is immediate and short-lived, while the degree of uncertainty surrounding the point estimates is considerable. At longer horizons the response turns positive. In the light of these findings, a LATW strategy would be compromised by timing frictions of monetary policy or its ineffectiveness at high economic costs. Hence, interest rate setting appears as a too coarse policy instrument for combating excessive asset price developments, and other instruments, e.g., prudential policies might be superior. Second, our results suggest that the FED reacts to both asset price shocks but more decisively to stock price shocks. Both shocks imply moderate macroeconomic effects, while the effect of the equity price shock appears more pronounced. Third, historical decompositions

[4]Identifying these differences may allow a more effective monetary policy conduct. Relatedly, if shocks stemming from a specific asset market turn out to exert only negligible effects on real economic fluctuations, even differences in the speed of adjustment to monetary policy impulses may not be associated with undesirable consequences.

[5]We are aware that other segments of the real estate market, such as, commercial property, may exhibit different features regarding, e.g., sluggishness and, hence, might lead to distinct conclusions regarding monetary policy transmission. We leave the analysis of additional segments of the real estate sector as an interesting route for future research.

of house prices reveal that deviations of the FED from the (model-implied) policy rule have not contributed considerably to the housing boom preceding the Great Recession.

Section 2 outlines the SVAR model and the statistical identification scheme. Section 3 empirically illustrates the independence based identification scheme by means of a five dimensional model. In Section 4 we present estimation results for the multiple asset price model, while we carry out robustness checks in Section 5. Section 6 concludes.

2 Empirical framework

2.1 The SVAR model

We consider a K-dimensional SVAR model of order p, i.e.,

$$y_t = c_t + A_1 y_{t-1} + \ldots + A_p y_{t-p} + B\epsilon_t, \quad t = 1, \ldots, T, \tag{1}$$

where $y_t = (y_{1t}, \ldots, y_{Kt})'$ is the $(K \times 1)$ vector containing the variables of interest, c_t denotes vector valued deterministic terms and A_i, for $i = 1, \ldots, p$, are $(K \times K)$ coefficient matrices of autoregressive parameters.

The structural shocks ϵ_t are uncorrelated for all leads and lags with $\mathrm{E}[\epsilon_t] = 0$ and normalized to have unit variance, $\Sigma_\epsilon = I_K$. The non-singular (impact) matrix B specifies the contemporaneous relations between the structural shocks and the reduced form innovations $u_t = B\epsilon_t$, with $\mathrm{E}[u_t] = 0$ and $\mathrm{Cov}[u_t] = BB' = \Sigma_u$. Hence, the reduced form innovations are a linear combination of the structural innovations. Moreover, we assume that the reduced form residuals are causal for y_t, i.e., $\det(I_K - A_1 z - \ldots - A_p z^p) \neq 0$ for all $|z| \leq 1$. The decomposition of $\Sigma_u = BB'$ is not unique and, hence, neither are the structural shocks $\epsilon_t = B^{-1}u_t$. Therefore, drawing on additional theoretical or statistical information, several approaches to identification of the structural matrix B have been proposed in the literature.[6] In this regard we rely on a statistical identification approach, viz the detection of independent structural shocks, which we briefly outline in the next section.

2.2 Independence based identification

In a conventional Gaussian SVAR setting the decomposition of the reduced form covariance matrix $\Sigma_u = BB'$ is not unique. By implication the matrices defined through B and any arbitrary rotation are observationally equivalent. In contrast, if reduced form residuals are non-Gaussian, independence of vector valued structural shocks ϵ_t can be exploited to obtain a unique matrix B (up to column ordering and scaling).[7] For instance, assuming independence, Moneta et al. (2013) adopt independent component analysis (ICA) to rank distinct recursive structural models corresponding to alternative variable orderings in u_t based on loss measures for the violation of joint independence. Recently, allowing for cyclic causality, a more flexible approach has been proposed in Lanne et al. (2017). They suggest to determine B by means of ML estimation, assuming a specific

[6]Popular strategies that resort to a-priori theoretical information are short-run zero restrictions, as, for instance, the recursive ordering of variables (Sims, 1980), long-run restrictions along the lines of Blanchard and Quah (1989) or sign restrictions in the spirit of Faust (1998) and Uhlig (2005). Killian and Lütkepohl (2017) provide a thorough textbook treatment of structural VARs.

[7]More exactly, according to the fundamental result in Comon (1994) at most one Gaussian component of the residual vector is admissible.

parametric distribution for the structural shocks (see also Gouriéroux et al. (2017) for a pseudo ML (PML) approach). In this work we adopt an identification framework where $\hat{B}$ is determined by minimizing a dependence measure in order to obtain mutually (most) independent shocks. We quantify the dependence level by means of the nonparametric distance covariance (dCov) statistic $\mathcal{U}$ of Székely et al. (2007).[8] The estimation procedure involves the nonlinear minimization of $\mathcal{U}(\hat{\epsilon}_t(\theta))$ with respect to a vector of rotation angles θ of Givens rotation matrices defining the set of admissible structural matrices $B(\theta) = DQ(\theta)$ and the corresponding structural shocks $\hat{\epsilon}_t(\theta) = B(\theta)^{-1}\hat{u}_t$. Here D denotes the lower triangular Choleski factor of $\widehat{\Sigma}_u$ and Q is a product of Givens rotation matrices (or, in general, rotation matrices). As shown in Matteson and Tsay (2017), the structural impact matrix $\hat{B}_{dCov} = B(\hat{\theta})$ that corresponds to the minimized dependence statistic for $\hat{\theta} = \text{argmin}_{\theta}\mathcal{U}(\hat{\epsilon}_t(\theta))$ is a consistent estimator for the structural system associated with least dependent structural shocks.

At this point a note on the choice of the dependence criterion is in line: Alternatives are manifold, for instance, the Cramér-von Mises (CvM) statistic of Genest et al. (2007) or the negentropy employed by Capasso and Moneta (2016) based on the efficient algorithm *FastICA* (Hyvärinen and Oja, 1997). However, the simulation exercise in Matteson and Tsay (2017) demonstrates that algorithms based on dCov dominate *FastICA* in terms of MSE. Moreover, our results are broadly confirmed when we substitute the distance covariance with the CvM statistic (for an application of the CvM statistic in an SVAR context, see Herwartz and Plödt, 2016), while dCov offers considerable advantages in computation time. To facilitate inference, we construct confidence intervals for the impulse responses by means of resampling techniques in the spirit of a so-called fixed design wild bootstrap (Gonçalves and Kilian, 2004).

2.3 Labeling of shocks

Even though the shocks detected by means of minimizing the dCov dependence criterion are statistically identified, they lack an immediate economic interpretation. For purposes of shock labeling the analyst can, in general, resort to the following three related tools: Estimated contemporaneous relations ($\hat{B}$), impulse response functions (IRFs) and forecast error variance decompositions (FEVDs). A minimum requirement for the existence of a specific shock consists in a significant contemporaneous reaction of the variable it is associated with. For instance, the federal funds rate should respond, at least on-impact, to a monetary policy shock (Lütkepohl and Netšunajev, 2014). Therefore, we first check the on-impact effects, proceed by analyzing the dynamic responses and compare these with a-priori (theory-based) sign patterns (Herwartz and Lütkepohl, 2014). Lastly, if these criteria are not fully decisive with regard to the uniqueness of economically labeled shocks, we resort to FEVD analysis and consider the percentage of explained variances to discriminate between otherwise similar shocks or to corroborate preliminary labels. Utilizing variance decompositions in this context is in the spirit of Uhlig (2004) suggesting FEVDs for the identification of technology shocks. More recently, Antolín-Díaz and Rubio-Ramírez (2016) use variance decompositions for benchmarking narrative sign restrictions. Similarly, Netšunajev and Glass (2017) give statistically

[8] To assess mutual independence, the dCov statistic measures the dependence level by considering the groupwise distance between the joint characteristic functions and that under independence (for more details see Matteson and Tsay (2017)). In this work, we apply the function *steadyICA* from the R-package *steadyICA* for the computation of the dCov statistic (Risk et al., 2015).

identified shocks an economic interpretation based on their contribution to the forecast error variance of certain variables. In the next sections we apply the outlined labeling strategy to the two model specifications analyzed in this work. Moreover, given labeled shocks, we study the role of monetary policy for asset prices and the reverse causation, i.e., the systematic monetary policy reaction to asset price shocks. Note that we consider positive shocks throughout the empirical sections, i.e., shocks that lead to an increase in the associated variable.

3 Applying independence based identification to the monetary policy stock price nexus

In this section, we illustrate the capabilities of the proposed identification strategy by means of a widely analyzed model specification. We investigate a five dimensional system comprising variables typically included in a monetary policy context augmented with stock prices. In the following we first give a brief overview of the related literature and continue with a discussion of the merits of the adopted identification approach and the respective empirical results.

3.1 Related VAR literature

Initial contributions, such as Patelis (1997), Thorbecke (1997) and Neri (2004), report rather modest effects of interest rate news on stock prices. Furthermore, asset prices react only sluggishly which appears to contradict established financial market theory. These studies, however, share a characteristic shortcoming owing to the recursive identification scheme employed since recursiveness implies that a contemporaneous reaction of monetary policy (and most other macroeconomic variables) to asset price signals is ruled out by assumption.[9] Early bivariate attempts, allow for contemporaneous interactions are Rigobon and Sack (2004) and Rigobon and Sack (2003) for the effect of exogenous interest rate fluctuations on stock prices and for the reverse transmission, respectively.[10]

Björnland and Leitemo (2009) address the issue of potential simultaneity by identifying the monetary policy and stock price shocks through a combination of short- and long-run restrictions in a more informative five dimensional system. Their results suggest a strong negative initial and dynamically persistent response of stock prices to contractionary interest rate signals. Regarding the reverse linkage, from stock markets to monetary policy, Björnland and Leitemo (2009) report a four basis point increase in the federal funds rate on impact, ensuing from an one percent rise in stock returns. The effect peaks during the first year and tapers off over the course of three years. Based on a non-Gaussian framework of mutually independent innovations, Lanne et al. (2017) lend credence to the results in Björnland and Leitemo (2009) by statistical tests for the

[9] Usually, under recursive identification asset prices (e.g., stock prices) are ordered last in the vector of endogenous variables (y_t) and therefore the monetary policy rule is assumed to incorporate information from the asset markets only with a one period (often one month or quarter) delay. The reader may consult Sellin (2001) for a comprehensive survey of the earlier empirical literature.

[10] In these studies, the monetary policy and stock price shocks are identified by exploiting (unconditional) heteroskedasticity present in high-frequency (i.e. daily) data, and pronounced instantaneous effects for both directions are highlighted. The analysis in Rigobon and Sack (2003, 2004), however, remains constrained to on-impact responses and does not address longer-term dynamic adjustments following the initial shock. The conclusions drawn in Rigobon and Sack (2003) on the response of monetary policy to stock price shocks are still disputed (see Furlanetto, 2011).

(now overidentifying) short-run exclusion restrictions. The (positive) identified monetary policy shock tends to lower a financial condition index modestly in the short-run. Their results, however, indicate a considerably less persistent response of asset returns compared with Björnland and Leitemo (2009). By means of an overidentified heteroskedasticty-based SVAR on an extended sample (1970 – 2007), Lütkepohl and Netšunajev (2014) find that the long-run restrictions proposed by Björnland and Leitemo (2009) are at odds with the data. Moreover, Lütkepohl and Netšunajev (2014) do not diagnose a significant reaction of the monetary policy instrument ensuing a stock price shock. The authors argue that the stock price shock bears the interpretation of a "news" shock (Beaudry and Portier, 2014), as opposed to the non-fundamental shock detected in Björnland and Leitemo (2009).

3.2 Data and pre-estimation diagnostics

We estimate a $K = 5$ dimensional model on monthly U.S. data covering the period from 1970:1 until 2007:6. This model comprises a linearly detrended log industrial production index (q_t), the annual change in log consumer prices (Δp_t), the annual change in the log of the World Bank (non energy) commodity price index ($\Delta comp_t$), the log of the S&P500 Composite Index deflated by the consumer price index as monthly real returns (Δsp_t) and the federal funds rate (ffr_t).[11] As in Björnland and Leitemo (2009) for this specification the model in (1) comprises an intercept $c_t = c$.

With regard to the variable transformations, we closely follow Björnland and Leitemo (2009) and Lütkepohl and Netšunajev (2014). On the one hand this facilitates comparability to the findings in the reference studies and on the other hand theoretical considerations and empirical findings favor this particular variable choice. First, based on theoretical models (as, for instance, Svensson, 1997) and empirical evidence from a recent meta-study (Rusnák et al., 2013) the output gap should be preferred over industrial production/GDP. Second, annualized inflation (and commodity price inflation) is closer to the targeted price indicator from the perspective of the policy maker than, e.g., monthly inflation rates, and is not subject to time-varying seasonal variations (Björnland and Jacobsen, 2010).[12]

We obtain estimates for the matrices $A_1, \ldots, A_p$ and the residuals u_t in equation (1) by means of least-squares estimation of the reduced form VAR model. Moreover, we include $p = 3$ lags based on the AIC and in accordance with the lag order choice in Lütkepohl and Netšunajev (2014).

We decide on the feasibility of the independence based method by means of tests for normality which rely on applying fourth order blind identification to the model disturbances $\left(\{\hat{u}_t\}_{t=1}^T\right)$. This type of test allows to statistically evaluate the number of Gaussian components in the reduced form VAR residual series and is included in the `R` package *ICtest* by Nordhausen et al. (2017). The null hypothesis of k_1 Gaussian components and $K - k_1$ non-Gaussian components, can even be rejected for $k_1 = 1$ Gaussian components with 1% significance indicating that all components are non-Gaussian.[13]

[11] Most series have been obtained from the St. Louis FRED database, while the stock price index was obtained from the FRED MD project website (McCracken and Ng, 2016).

[12] Since the use of year-on-year changes might induce serial correlation we estimated the VAR on monthly log changes as a robustness check, and find that the main results remain unaffected.

[13] We apply a bootstrap based test implemented in the function `FOBIboot` (Nordhausen et al., 2017). Moreover, note that we arrive at the same test decision using component-wise Jarque-Bera tests.

Before presenting and discussing the identified model, it is noteworthy to point out some (possibly fundamental) advantages of the identification approach relying on a nonparametric independence measure as compared with those developed in Lütkepohl and Netšunajev (2014) and Lanne et al. (2017). Firstly, the heteroskedasticity-based identification by Lütkepohl and Netšunajev (2014) suffers from a lack of statistically significant shifts in volatility of the structural shocks (at least for one shock) and, hence, does not result in a fully identified system. Secondly, the assumption of Gaussian model disturbances, as in Lütkepohl and Netšunajev (2014) is violated for the dataset at hand and consequently conflicts with the objective of studying isolated (that is independent) unit shocks.[14] Thirdly, following an ML (or pseudo ML) approach that explicitly abandons the Gaussianity assumption as, e.g., Lanne et al. (2017) or Gouriéroux et al. (2017), requires a specific parametric (non-normal) distribution (i.i.d.) of the structural shocks which might be in conflict with the data in relatively large samples. Table 1 documents the results from two-sample Kolmogorov-Smirnov tests on the marginal residuals (u_{kt}) subject to several (sub)sampling schemes. These tests provide, in general, evidence against identically distributed model disturbances. More specifically, there are at least two residual components for which we can reject the null hypothesis that subsamples are drawn from the same distribution at 10% significance for all three distinct subsampling schemes. Hence, the assumption of identically distributed shocks (over the time dimension) of (pseudo) ML based estimation and identification is likely to be violated in this class of VAR models.[15]

3.3 Results from the identified model

3.3.1 Shock labeling

We first focus on the contemporaneous responses of real stock prices and of the federal funds rate in order to label the stock price and monetary policy shocks. Figure 1 displays the impulse responses characterizing the five dimensional system (jointly with 68% confidence bands). Similar to findings in Lütkepohl and Netšunajev (2014) only the fifth shock (ϵ_5) exhibits a considerable (and statistically significant) impact effect on the policy instrument. Therefore, we label the fifth shock "monetary policy shock". Labeling the stock price shock is not quite as obvious. In this case two shocks, the first (ϵ_1) and the third (ϵ_3), exert a significant contemporaneous effect on stock prices. The magnitude of the response associated with the first shock is, however, substantially larger than that of the second shock. The instantaneous effects therefore strongly suggest to label the former shock "stock price shock". Figure 2 shows the FEVDs for all shocks and variables. The first shock maximizes the variance of stock returns, explaining almost 100% on-impact and persistently continues to account for about 90% of the variance at all horizons. Therefore, FEVDs confirm the insights from the analysis of impulse response patterns to label the first structural innovation (ϵ_1) stock price shock. Although we are primarily interested in the monetary policy asset price nexus, the independence based approach delivers a dynamic structural system that allows us to identify also commonly analyzed shocks like aggregate demand and supply shocks. Based on standard assumptions, by combining impulse response analysis and FEVD, we can label the shock displayed

[14] Even though Lütkepohl and Netšunajev (2014) explicitly model and utilize conditional heteroskedasticity in the data for identification, conceiving unit shocks as isolated appears misleading as long as they are not mutually independent.

[15] For a more detailed description of the testing procedure and particularly the sampling schemes we refer to the notes below Table 1 in the appendix.

in the second row (ϵ_2) of Figure 1 aggregate supply shock and the shock depicted in the fourth row (ϵ_4) aggregate demand shock.

3.3.2 The link between monetary policy and stock prices

Next, we examine the dynamic responses of the monetary policy and stock price shocks. Beginning with the monetary policy shock, we observe a weak indication of a prize puzzle ensuing from a 50 basis points (bps) increase (one standard deviation) of the federal funds rate.[16]

Showing an inverted hump-shape, and consistent with previous studies (see, e.g., Ramey, 2016), the effect on output is expansionary for the first few months and significantly diminishing at longer horizons. Real stock prices drop immediately and considerably in response to the interest rate shock. The dynamic response pattern of the stock price shock exhibits a rather pronounced effect on output and a decisive systematic monetary policy reaction. The federal funds rate increases by about 25 bps four to five months after the impact, lasting for around five years. Comparing both shocks to those obtained by the identification schemes employed in Björnland and Leitemo (2009) and Lütkepohl and Netšunajev (2014), we note that our results are closer to findings based on the Björnland and Leitemo (2009) identification scheme both qualitatively as well as quantitatively. In summary, we conclude that identification by means of independent shocks employing a nonparametric dependence criterion apparently delivers a structural model which inherits economically meaningful shocks.

4 Monetary policy, equity and housing: Implications for a LATW strategy

In the literature addressing the LATW debate within a SVAR framework, different asset markets are frequently treated as a homogenous market disregarding possible (important) differences (e.g., Galí and Gambetti, 2015; Beckers and Bernoth, 2016). As outlined before there might, however, exist essential differences and associated consequences for the feasibility of a proactive monetary policy conduct. Moreover, real estate prices have been identified along side the credit-to-GDP ratio as most important indicators of financial crises. Incorporating house prices is therefore of utmost importance in a LATW context (IMF, 2015; Leamer, 2015).[17] Taking account of these considerations, we analyze in the following a six dimensional model ($K = 6$). This model comprises, in addition to the variables listed in Section 3.2, the (CPI deflated) Case-Schiller house price index (Δhpi_t) which enters the model in form of log returns. Furthermore, we replace the federal funds rate by the shadow rate compiled by Wu and Xia (2016).[18] The use of a (repeated sales based) house price

[16] A positive price response in the short-run is commonly observed, also for more recent identification approaches, using external information or narrative sign restriction (see for an overview Ramey, 2016). Moreover, informational insufficiency does not seem to play an important role, since also factor augmented VARs (FAVARs) tend to be associated with short-lived price increases, as documented in, for instance, Forni and Gambetti (2010). Our results rather constitute evidence for a cost channel, caused by the transmission of higher financing costs of firms to consumer prices.

[17] For instance, Leamer (2015) advocates for an exceptional role of housing even in more "normal" times. He argues that the informational content in house prices, but more so in residential investment and sales volumes, is superior to any other macroeconomic aggregate in terms of indicating recessions (see also Leamer, 2008).

[18] We opt for the shadow rate, because interest rates have been at the zero lower bound since around the end of 2008. Before 2008 the federal funds rate and the shadow rate are identical.

index for monthly data now limits the sample period to cover 1975:1 until 2015:6. Positioning our modeling approach in the literature, it is worth noting that it has a similar scope as the models considered by, e.g., Eickmeier and Hofmann (2013); Björnland and Jacobsen (2013); Cheng and Jin (2013); Simo-Kengne et al. (2016) who investigate the link between monetary policy and asset prices and consider more than one asset class.[19]

Eickmeier and Hofmann (2013) provide evidence for the interaction between monetary policy and several asset prices in a data-rich environment. Based on a quarterly U.S. sample (1987Q3 – 2007Q4) their results from a factor augmented VAR (FAVAR) point to a missing transmission channel for a number of stock price indices but simultaneously to a quite active channel in the case of property prices.[20] Furthermore, the authors find considerable sample dependence of the strength of monetary transmission. Eickmeier and Hofmann (2013) do not, however, provide evidence for asset price shocks. In contrast, Björnland and Jacobsen (2013), estimate a model incorporating both asset classes (house and stock prices) and investigate also the reverse causation. Adopting the identification approach in Björnland and Leitemo (2009) (1983Q1 to 2010Q1), they find remarkably distinct roles of house and stock markets in the monetary policy transmission process. In contrast to stock prices, shocks stemming from the housing market, move output and prices considerably, but the monetary authority seems to respond with a significant delay of about two quarters to these shocks.

4.1 Diagnostics, estimation and labeling of shocks

We adopt the same estimation and identification approach as outlined in Section 2. Preliminary tests on the Gaussianity of the estimated residuals $\left(\{\hat{u}_t\}_{t=1}^T\right)$[21] indicate that none of the components of u_t are Gaussian (see Section 3.2). Moreover, two-sample Kolmogorov-Smirnov tests summarized in Table 1 cast doubt on the assumption of identically distributed shocks and, hence, speak against the application of (pseudo) ML based methods.

In order to investigate the monetary policy asset price nexus we intend to label three shocks, namely a monetary policy, a stock price and a house price shock. Figure 3 displays the impulse responses for all shocks and variables. Beginning with the labeling of the *monetary policy shock*, we observe only one shock that invokes a noticeable increase in the federal funds rate on impact, i.e., the shock associated with the responses shown in the fifth row (ϵ_5) of Figure 3. Hence, we label this shock monetary policy shock. Examining the variance decompositions depicted in Figure 4 for all shocks and variables confirms this impression. The shock contributes almost 100% to the variance of the federal funds rate on impact and more than 50% at medium horizons. While this variance share appears rather large, it conforms with common findings from the monetary SVAR literature. For instance, Bernanke et al. (2005) and Ahmadi and Uhlig (2015) find that monetary policy shocks contribute 45% and 30-40% (100%) at long horizons (short horizons), respectively.

[19]Important studies which focus on the transmission of monetary policy through (U.S.) house prices only are, for instance, Goodhart and Hofmann (2001); Iacoviello (2005); Del Negro and Otrok (2007); Goodhart and Hofmann (2008). With regard to identification, the dominant strategy in these studies is, however, to assume a recursive ordering of variables.

[20]Eickmeier and Hofmann (2013) employ an identification approach based on a combination of contemporaneous zero and short-run sign restrictions.

[21]We estimate the model with $p = 3$ lags and include an intercept and trend term both according to the AIC, i.e, $c_t = \gamma_1 + \gamma_2 t$ in Equation (1).

Similarly, Jarocinski and Smets (2008) estimate the FEV share explained by the interest rate shock to be around 75% on-impact.

We conjecture that the relatively larger share at longer horizons can be attributed to differences in the sample period between both specifications, i.e., the stock prices only model in section 3 and the extended model considered in this section. This conjecture grounds on two observations. First, our sample includes the Great Recession (GR) and post Great Recession period which undoubtedly are characterized by larger deviations from the policy rule. Second, the explained FEV share is smaller for the five ($K = 5$) variable VAR presented in Section 3. Since we account for unconventional monetary policy measures by using the shadow rate, the higher variance share of the ffr explained by monetary policy shocks is likely caused by less predictable policy actions.

The labeling of the *asset price shocks* is more ambiguous by means of impulse response or structural impact multiplier analysis, although the first (ϵ_1) and the sixth shock (ϵ_6) are suitable candidates for the stock price and house price shocks, respectively.[22] In an attempt to overcome remaining concerns regarding the validity of the economic interpretation of shocks, we resort to FEVDs (Figure 4). We indeed find that the first shock (ϵ_1) maximizes the variance of stock prices and explains about 90% of variation at all horizons. Similarly, the sixth shock (ϵ_6) maximizes the variance of house prices contributing about 70% on impact and 75% at longer horizons. Accordingly, we label the first shock "stock price shock" and the sixth shock "house price shock". It is noteworthy that the correlation between the identified stock price shocks from the multiple asset price ($K = 6$) and the stock price model in Section 3 ($K = 5$) is 0.95. On the other hand, the correlation between the house price and the stock price shock merely amounts to -0.105, which further corroborates the labeling of shocks.[23] The large share of variance explained by the house price shock is in line with recent estimated DSGE model evidence. Iacoviello and Neri (2010) report that the long-run contribution of housing demand and housing technology shocks to house price fluctuations amounts to around 60%. Naturally, conditional on correct identification, the single house price shock identified in the present paper should comprise, at least to a large extent, the multitude of possible, more specific hosing shocks as considered in larger models such as Iacoviello and Neri (2010).[24]

4.2 The effect of monetary policy on housing and equity

An unexpected monetary tightening (ϵ_1 in Figure 3) evokes an insignificant (at the 68% confidence level) fall in prices and a pronounced and persistent decrease in commodity prices. Moreover, a small increase in output is triggered in the very short-run followed by a hump-shaped fall after 4–5 months subsequent to the initial federal funds rate impulse (of about 100 bps). Hence, the monetary policy shock exhibits the typical characteristics reported in the pertinent literature. In accordance with findings in Ramey (2016) the prize and output puzzles are mitigated in the longer

[22] The two asset price shocks are difficult to distinguish by means of theoretically derived sign patterns. Both shocks could potentially show similar dynamic responses with regard to prices and output, i.e. they are both likely associated with non-negative responses of prices and output. In addition, one cannot rely on possible differences in the responses of the federal funds rate to discriminate between shocks, since this is one of the focal questions of this study.

[23] Similarly, even though, we used a different sample period for the illustrative example in Section 3, the monetary policy shocks from both specifications ($K = 5$ and $K = 6$) exhibit a relatively strong correlation of 0.56.

[24] For empirical results from an SVAR see, e.g., Jarocinski and Smets (2008). They find that a house price shock explains between 55% (20%) of the FEV in the short (long run).

sample (compare with the IRFs of the five dimensional model in Section 3). This is likely due to the fact that monetary policy was less erratic during the Great Moderation period and hence a "true" shock was harder to identify. With the onset of the Great Recession this has presumably changed. The contemporaneous impact of the monetary policy shock on real equity prices is mildly negative (about -0.5%) and the response turns positive quickly. The point estimator indicates a sizable positive reaction in the medium-run, which is, however, not statistically different from zero. Turning to the impact of monetary policy on real house prices, we notice that real house prices react more sluggishly to unexpected federal funds rate signals than stock prices: The response of real residential property prices is practically zero during the first quarter, turns negative thereafter and continues dropping over the entire time period of five years. The impact is rather strong peaking at about -0.75% after five years.

Consequently, our results conform with previous findings stating that equity prices respond immediately but rather short-lived to exogenous interest rate shocks whereas house prices are more persistent and react delayed (Björnland and Jacobsen, 2013).[25] From a theoretical perspective one would expect such results, for at least two reasons. Firstly, the stickiness of house prices can be traced back to imperfections and resulting inefficiencies in the residential property market, such as high transaction and search costs, different timing preferences between buyers and sellers (and associated frictions regarding the adjustment of reservation prices of property owners and potential buyers), social norms and tax considerations (Case and Shiller, 1989; Leamer, 2008; Geltner, 2015). Moreover, the heterogeneity of traded assets implies that the market value of houses must be estimated based on historical sales data which introduces backward looking behavior into the market (Iacoviello, 2010).[26] On the other hand, we conjecture that the slower transmission to house prices in contrast to stock prices partly reflects differences in the financing structure of households and firms: The residential property market is closely tied to long-term mortgage loans which are mostly fixed-rate contracts. An interest rate hike affects mainly new loans, and therefore can be expected to transmit to house prices less rapidly than to stock prices. Stock prices adjust more timely owing to the discount channel and possibly due to a larger share of short-term (variable rate) loans in the corporate sector. These factors appear to explain the initial adjustment following the monetary policy shock and to some extent also the long-run effect on house prices. We conjecture, however, that the double burden faced by households – a rise in debt servicing costs induced by the interest rate hike (presuming there is considerable pass-through) and a simultaneous worsening of the income situation of households – represents an additional explanation for this long-run effect, i.e., the persistence of the decline in house prices at the longer horizons (from four years onwards).

In quantitative terms the responses of the real asset prices shown in Figure 3 are less pronounced than their counterparts in Björnland and Jacobsen (2013), where stock prices fall even by 10% on impact and house prices reach their minimum at -4%. Similarly, variance decompositions in

[25] Studies that scrutinize interdependence between monetary policy and a single asset class are, for instance, Musso et al. (2011) for house prices and Björnland and Leitemo (2009) for stock prices. The former authors find a rather delayed and persistent response of house prices, whereas the latter study reports an immediate, short-lived and moderately negative response of equity prices to monetary policy shocks. Based on modeling multiple asset classes, for instance, Eickmeier and Hofmann (2013) and Assenmacher-Wesche and Gerlach (2008) arrive at similar conclusion as the present study.

[26] Large cycles in sales volumes have been detected as a symptom of the stickiness of house prices as reported in, e.g., Leamer (2008).

Björnland and Jacobsen (2013) suggest that exogenous monetary policy explains up to 40% of the variation in house prices, whereas in the present study monetary policy accounts only for a smaller fraction ($\leq$ 5%). Nevertheless, results in Björnland and Jacobsen (2013) might be considered as an upper bound for the findings in the literature.[27] In Section 5 we present results which suggest that the differences in the magnitude of the responses of real house prices can be attributed to the modeling of distinct data frequencies.

What do these findings imply for the applicability of a LATW policy? Taking the IRFs at face value, an increase of 100 bps in the federal funds rate was associated with a medium-term decrease of around 0.5% in real house prices and a short-term fall in real stock prices of the same magnitude. This coincides with a 0.3% medium-term rise in the output gap. Hence, a moderate reduction of imbalances comes at high economic costs. Accounting for differences in the dynamic propagation of monetary policy in both asset prices and for the fact that the response of equity prices might well be positive at longer horizons, a LATW policy seems ineffective according to our results.

4.3 Exogenous asset price movements and their macroeconomic effects

The first and last rows of Figure 3 show the IRFs for the real equity and house price shocks, respectively. The impulse responses of the federal funds rate to both asset price shocks reveal a similar hump-shape, albeit the reaction to a house price shock is slightly more delayed (in terms of time until the peak effect is reached), and less pronounced compared with the response emanating from an equity shock. The systematic policy response to house price shocks is surprisingly negative in the short run. Concerning the magnitudes of the systematic policy reaction in terms of peak effects, these are somewhat weaker than those found in Björnland and Jacobsen (2013). In both cases, the effect is quite persistent, lasting for more than five years.

The dynamic patterns with regard to prices and output are broadly similar for both asset price shocks, albeit the stock price shock exhibits stronger macroeconomic effects. At shortest horizons the responses of output to both asset price shocks are basically zero (insignificant or even slightly falling), reach a maximum after around one year, and turn negative (stock price shock) or die out in the long term (house price shock).[28] We interpret these macroeconomic effects as an indication for the existence of an effective wealth channel which boosts consumption and/or an effect on investment resulting from an increase in Tobin's q. Judging by the weak instantaneous response of the federal funds rate and the rather strong macroeconomic implications of asset price shocks, it can not be ruled out that the FED reacts mainly to the latter.

The stronger macroeconomic effect and the apparently higher weight of stock price shocks in the monetary policy rule are reflected in the FEVDs. The fraction of variation in the policy instrument explained by asset price shocks is moderate at longer horizons only for the equity price

[27]As already noted by Lütkepohl and Netšunajev (2014) this might be attributable to the long-run restrictions imposed by Björnland and Jacobsen (2013). Based on a structural FAVAR, Eickmeier and Hofmann (2013) report similar magnitudes for real house prices, but they do not find any statistically significant reaction of real equity to monetary policy shocks. In contrast, Del Negro and Otrok (2007) report a lower bound of -0.6% for the response of house prices (for a 15 bps monetary policy shock), based on several set-identified models. It should be noted that these authors consider a residential property price factor and not a specific house price index.

[28]For instance, output peaks at around 0.2% after a stock price shock of one standard deviation, i.e., an initial rise of the S&P 500 index by 3.5%. Note that in Björnland and Jacobsen (2013) real stock prices (real house prices) are normalized to increase by 10% (1%) in the first quarter, for a stock price shock (house price shock).

shock which explains at most 10%. The importance for fluctuations in economic activity and prices is negligible for the house price shock and small (around 5% for output) for the stock price shock. Nonetheless, in conclusion it seems that there are more similarities than differences between both asset price shocks.

4.4 Assessing the role of monetary policy during the house price boom

Finally, we take up the debate concerning the role of loose monetary policy – in the sense of downward deviations from the model implied Taylor-type rule – in the build-up of the price bubble on the real estate markets. For this purpose we present counterfactuals based on historical decompositions. The right-hand side panel in Figure 5 depicts the contribution of structural shocks to the baseline real house price growth (solid line) with the monetary policy shock shut down (dashed line). Accordingly, the difference between the baseline time series and the counterfactual is the model implied contribution of monetary policy shocks to real house price returns. This cumulated effect is shown in the left-hand side panel of Figure 5. The counterfactual housing return series takes on mostly smaller values than the actual series of housing returns, which implies that deviations from the implied policy rule have mostly had a positive effect on house prices during the depicted period. Even though the difference appears negligible, the cumulated effect is non-trivial: Over the years from 2000 until 2006 it amounts to 4 percentage points. Compared with a total increase of 37 pps of the detrended real house price series, however, monetary policy has played a minor role in the recent real estate price boom. This is in line with results in Del Negro and Otrok (2007) and Eickmeier and Hofmann (2013).[29]

5 Robustness

We discuss several robustness checks for the multiple assets baseline model, exploring a variety of modifications with regard to the estimation period, lag order and different data transformations. Firstly, we apply the adopted identification strategy to alternative sample periods, viz, the period 1975M1 – 2007M6, closer to the sample period used in Lütkepohl and Netšunajev (2014), a sample spanning 1980M1 – 2010M6, similar to Björnland and Jacobsen (2013) and a sample from 1987M1 until 2007M6. The latter is motivated by a discussion in Eickmeier and Hofmann (2013) attributing differences in the strength of the transmission channel via house and stock prices to pervasive changes in financial regulations. The abolition of the so-called "Regulation Q" summarizes these regulatory adjustments, which were completed around 1987. Moreover, this shorter period is characterized by a more stable monetary policy regime (Greenspan) and therefore provides an interesting object for robustness analysis.

We cannot find any considerable dependence of the responses of real house and stock prices to monetary policy shocks on the sample period. There is some variation in the strength of the transmission of monetary policy via house prices but not a clear pattern. Particularly, the sample covering the period with financial deregulation (1987M1 until 2007M12) does not confirm findings

[29] It is noteworthy that Eickmeier and Hofmann (2013) report marked differences in the contribution of monetary policy shocks to house prices, depending on the specific house price index used. We compare our results to theirs for the Case-Shiller house price index.

in Eickmeier and Hofmann (2013), according to which the magnitude of the house price response increases substantially. Overall, subsample results support our key conclusions.

Additionally, we conduct robustness tests for *i.*) an alternative frequency (quarterly) with replacing the industrial production gap by GDP growth; *ii.*) a model in log-levels (replacing all variables that were in growth rates); *iii.*) the estimation of the baseline specification with alternative lag orders (i.e., $p = 6$ and $p = 12$); *iv.*) replacing the industrial production gap by the growth rate; *vi.*) including only an intercept term; *vii.*) CPI and commodity price inflation as well as industrial production in monthly log changes; *viii.*) specification (*vii.*) for a lag length of $p = 12$.[30] Likewise, the key conclusions from the baseline model remain intact. Nonetheless, it is noteworthy to point out several global findings from these exercises:

Firstly, the rather sluggish but clearly negative response of house prices to positive monetary policy shocks is quite robust over all specifications. Interestingly, the transmission via residential property prices is most effective for quarterly data peaking at -3% and is, hence, comparable with findings in Eickmeier and Hofmann (2013) and Björnland and Jacobsen (2013). Secondly, the mild impact of interest rate shocks on stock prices observed for the baseline model is confirmed by the robustness checks. In almost all specifications there is only a small initial effect followed by a subsequent statistically insignificant positive response. Frequently, even the short-run negative effect is insignificant though. Thirdly, there tends to be some support for a markedly different systematic reaction of monetary policy. We observe that in some models the response of the federal funds rate to house price shocks is statistically insignificant and/or close to zero.[31] Fourthly, also the main conclusion that the FED reacts more timely and decisively to stock price shocks is found in most specifications. A related occasional finding is that asset price shocks have a long-run effect on output associated with a change in the interpretation of the shocks. For instance, results obtained for the model in log-levels indicate that the stock price shock bears the interpretation of a news shock, and hence lend some support to Lütkepohl and Netšunajev (2014). Finally, we note that test results on the non-Gaussianity of shocks appear to be unaffected by specific filters applied to the variables in levels. In particular, the robustness checks (vii) and (viii) provide evidence against reservations regarding the specific data treatment (annualized time series).

6 Conclusions

We revisit the asset price monetary policy nexus surrounding the lean against the wind (LATW) debate which has gained considerable attention after the recent Great Recession. Acknowledging the existence of several rivaling identification approaches in the SVAR literature on monetary policy transmission through asset prices, we use a novel method for identification. More specifically, the non-Gaussianity of shocks is exploited by means of a dependence measure in order to obtain a just-identified system. We illustrate this identification approach by estimating a five dimensional model based on U.S. data that has been analyzed before by Lütkepohl and Netšunajev (2014) and Björnland and Leitemo (2009). We demonstrate that the present method delivers a rich set of IRFs

[30] For the sake of brevity we do not report the results of these modifications but they are available from the authors upon request.

[31] The response is statistically insignificant for the models estimated on data from 1987M1 until 2007M12 and 1980M1 – 2010M6. For the latter sample period the IR is additionally also basically zero after 10 months.

that correspond to meaningful economic shocks. The dynamic patterns of the monetary policy and stock price shocks resemble more closely findings obtained from the Björnland and Leitemo (2009) identification scheme.

Proceeding from the results of the five dimensional system, we investigate a model that is more informative in the LATW context. In the spirit of Björnland and Jacobsen (2013), we examine a six dimensional model comprising house and stock prices for monthly data spanning the period from 1975M1 – 2015M6. In line with the majority of the literature, we find that monetary policy transmits to residential property markets. Contractionary monetary policy shocks trigger a sluggish but persistent and pronounced decline in real house prices. This finding is robust with respect to alternative choices of the sample period and other modifications of model specification. The transmission of monetary policy via the stock markets is less clear according to our results. We observe an immediate, mildly negative short-lived response to an unit contractionary interest rate shock. However, the response of equity turns positive at longer horizons (although insignificantly). Combined with the results from robustness tests, we conclude that interest rate shocks have most likely only a negligible effect on stock prices. Finally, supporting similar conclusions of Del Negro and Otrok (2007) and Eickmeier and Hofmann (2013), historical decompositions show that loose monetary policy was not a major factor in the recent housing boom preceding the Great Recession.

In summary, our results allow the tentative conclusion that a proactive monetary policy is likely to be impracticable for at least two reasons: Firstly, a central bank conducting a LATW policy has to face either timing frictions of policy transmission, or ineffectiveness of conventional interest rate policies on stock markets. Secondly, monetary authorities encounter rather moderate declines in house prices (and possibly also in equity prices) at simultaneously substantial costs in terms of output. This implies that conventional monetary policy instruments might lack effectiveness to counter imbalances on the asset markets.

In contrast to Björnland and Jacobsen (2013), we find only weak support for a distinct reaction of the central bank to shocks originating from the asset markets. It appears, however, that the FED incorporates equity price shocks with a slightly higher weight into the monetary policy rule. Moreover, both asset price shocks imply similar macroeconomic effects qualitatively, whereas the impact of the stock price shock on output is stronger.

Methodologically this study highlights that the adopted non-Gaussian framework of structural modeling can provide a valuable tool for the analysis of dynamic systems which might otherwise be grounded on controversial a-priori assumptions. Considering the fruitful detection of independent shocks, a promising avenue for future research consists of extending the information set. Basing a similar analysis on a large panel of macroeconomic data and applying factor models (FAVAR) could provide insights into the dynamics of an even richer set of asset prices. The nonparametric independence based approach can help particularly with identifying distinct asset price shocks, for which a-priori assumptions are especially controversial. Another route for future research consists in allowing for non-linearities, taking into account the growing evidence for the regime-dependence of monetary policy effectiveness, i.e., dependence on volatility/financial stress regimes (Eickmeier et al., 2016; Saldías, 2017), stages of the credit (Balke, 2000; Ciccarelli et al., 2015) or of the business cycle (Avdjiev and Zeng, 2014).

7 Appendix

TABLE 1

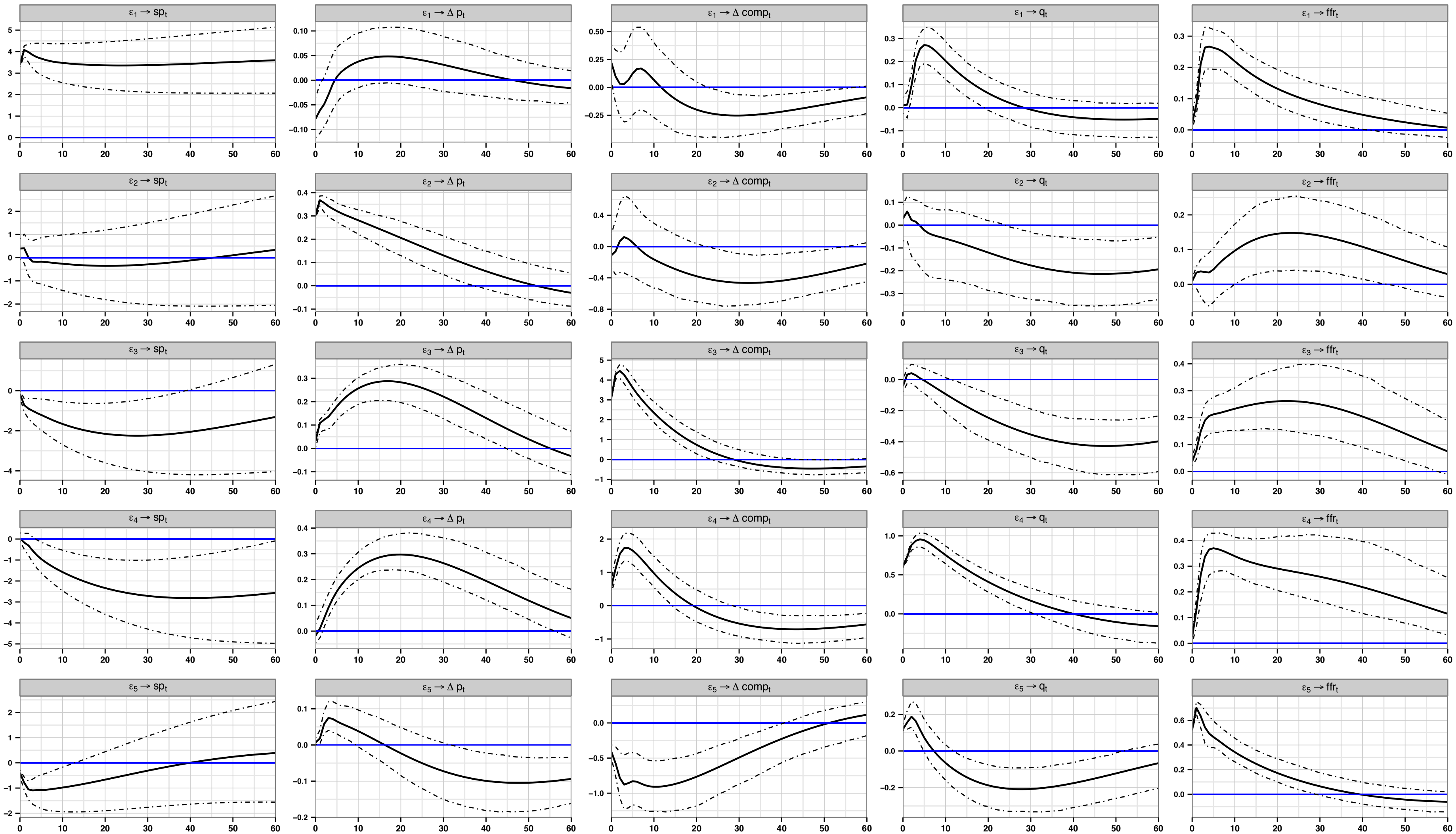
Kolmogorov-Smirnov Test Results for the Stock Price ($K = 5$) and the Multiple Asset Price ($K = 6$) Models

	u_{1t}	u_{2t}	u_{3t}	u_{4t}	u_{5t}	u_{6t}
		K=5 dimensional model				
Split50	0.5103	0.1050	0.0033	0.0160	0.0001	.
Split75	0.6269	0.4405	0.3203	0.0092	0.0366	.
Split25	0.5083	0.0282	0.0059	0.0046	0.0000	.
		K=6 dimensional model				
Split50	0.4456	0.5901	0.3809	0.3809	0.0013	0.1224
Split75	0.5095	0.1039	0.0251	0.9932	0.0006	0.0038
Split25	0.5953	0.4294	0.5095	0.5518	0.0008	0.8835
Split-crisis	0.4400	0.2034	0.0529	0.7114	0.0030	0.0870

Notes: p-values for Kolmogorov-Smirnov tests of all components of the residual vector u_t and different sampling schemes for both specifications ($K = 5, 6$). The null-hypothesis is that both random variables are drawn from the same distribution. "Split50" denotes the p-values for a test on subsamples obtained by spliting the whole sample at $0.5T$,i.e., $u^{(1)} = \{u_t\}_{t=1}^{0.5T}$ and $u^{(2)} = \{u_t\}_{t=(0.5T+1)}^{T}$, "Split75" devides the series $\{u_t\}_{t=1}^{T}$ into two subsample at $t = 0.75T$, i.e., $u^{(1)} = \{u_t\}_{t=1}^{0.75T}$ and $u^{(2)} = \{u_t\}_{t=(0.75T+1)}^{T}$. "Split25" denotes the opposite sampling scheme, i.e. $u^{(1)} = \{u_t\}_{t=1}^{0.25T}$ and $u^{(2)} = \{u_t\}_{t=(0.25T+1)}^{T}$. Finally "Split-crisis" devides the whole sample into a subsample from 1976 untill 2008 (pre-crisis) and one comprising the crisis and post-crisis periods for the six dimensional model.

FIGURE 1

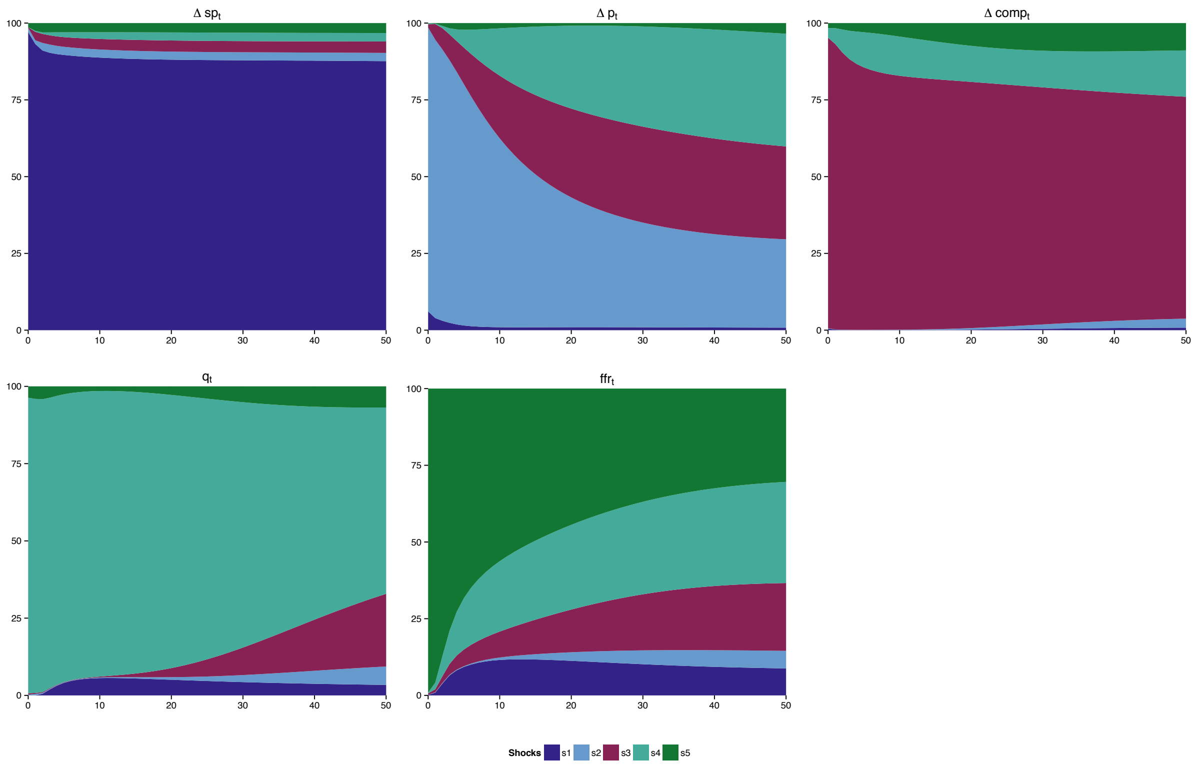
Impulse Responses to a Unit Shock for the Stock Price Model ($K = 5$)



Notes: Solid lines denote the point estimate of the response, the dashed lines limit the 68% confidence bands based on 1000 bootstrap replications. Note that the impulse responses for real stock prices sp_t have been accumulated. ϵ_1 denotes the stock price shock, ϵ_2 the aggregate supply shock, ϵ_4 the aggregate demand shock and ϵ_5 the monetary policy shock.

FIGURE 2

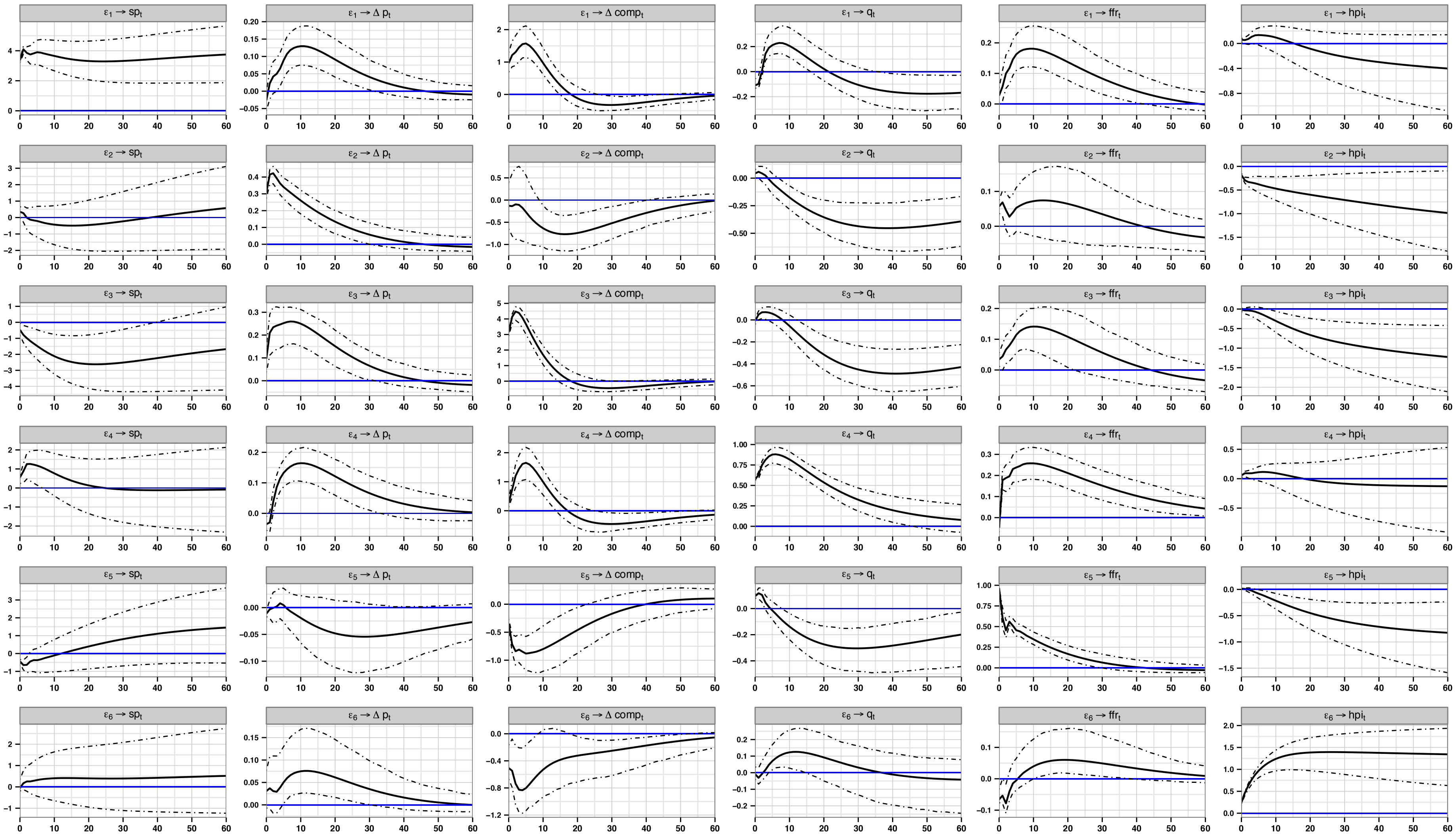
Forecast Error Variance Decompositions for the Stock Price Model ($K = 5$)



Notes: "s1" denotes the stock price shock (blue area), while "s5" stands for the monetary policy shock (green area).

FIGURE 3

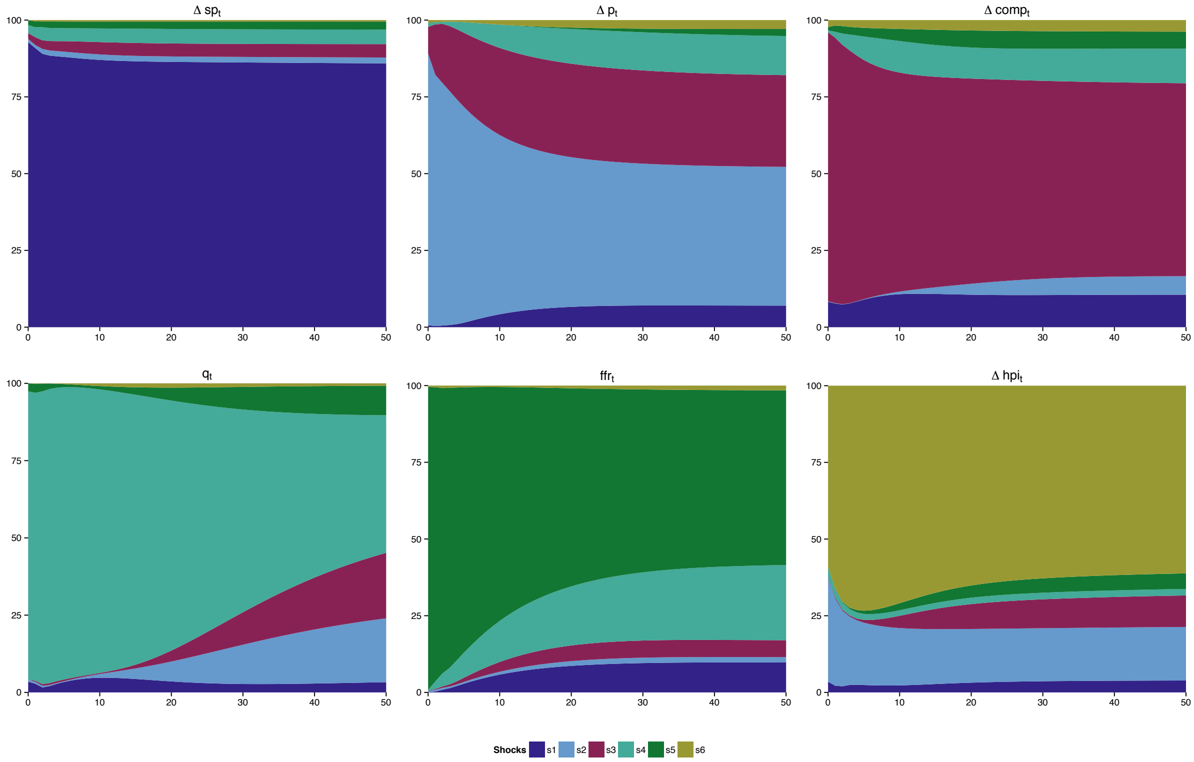
Impulse Responses to a Unit Shock for the Multiple Asset Price Model ($K = 6$)



Notes: Solid lines denote the point estimate of the response, the dashed lines limit the 68% confidence bands based on 1000 bootstrap replications. Note that the impulse responses for real stock sp_t and real house prices hpi_t have been accumulated. ϵ_1 denotes the stock price shock, ϵ_5 the monetary policy shock and ϵ_6 the house price shock.

FIGURE 4

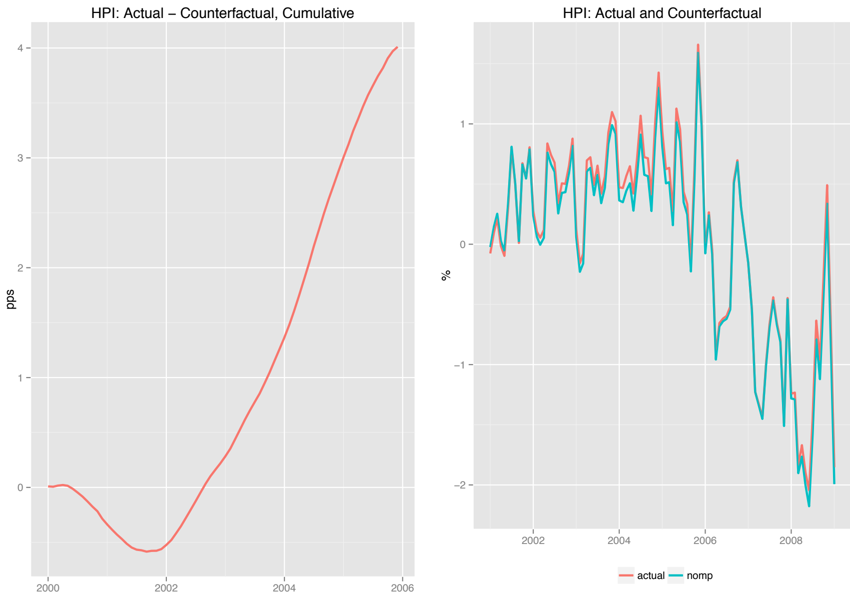
Forecast Error Variance Decompositions for the Multiple Asset Price Model ($K = 6$).



Notes: "s1" denotes the stock price shock (blue area), "s5" stands for the monetary policy shock (green area) and "s6" denotes the house price shock (brown area).

FIGURE 5

Historical Decomposition Based Counterfactual



Notes: Right-hand side panel: Historical decomposition based counterfactual without MP shock ε_5 (dashed line) vs. actual series (excluding deterministic terms) for real house prices (solid line). Left-hand side panel: Cumulated difference between actual and counterfactual.

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